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Author: George Symeonidis

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Collusion, Profitability and Welfare: Theory and Evidence

GEORGE SYMEONIDIS

University of Essex

Address for correspondence: Department of Economics, University of Essex, Wivenhoe Park, Colchester CO4 3SQ, U.K. E-mail: symeonid@essex.ac.uk

Highlights:

- I examine how structural industry characteristics affect the magnitude of the welfare effect of collusion
- The theory is consistent with evidence from a natural experiment of policy reform, the introduction of cartel law in the UK in the late 1950s
- Price-cost margins declined after the breakdown of cartels in low-capital and larger-sized industries relative to capital-intensive and smaller-sized ones
- The welfare loss from collusive pricing may be relatively small in certain types of industries where collusion often occurs in practice

Abstract: In a differentiated oligopoly model with free entry, the static welfare loss from collusion is larger the lower the entry cost, the larger the market size and the higher the degree of product differentiation. The cartel overcharge is larger the lower the entry cost and the larger the market size, and is independent of the degree of product differentiation. These theoretical results are consistent with evidence from a natural experiment of policy reform, the introduction of cartel law in the UK in the late 1950s. Price-cost margins declined after the breakdown of cartels in low-capital and larger-sized industries relative to capital-intensive and smaller-sized ones. There is weaker evidence of a fall in price-cost margins in consumer good and advertising-intensive relative to producer good and low-advertising industries. Crucially, these effects are not observed for industries not affected by the cartel law.

A comparison of these findings with evidence on the incidence of collusion suggests that the welfare loss from collusive pricing may often be smaller in industries where cartels tend to form than in those where collusion is more difficult to sustain.

JEL classification: L13, D43, L11, L60

Keywords: Collusion, cartels, free entry, welfare, profitability, UK manufacturing, antitrust policy

1. Introduction

Despite the large theoretical and empirical literature on cartels and collusion, the effect of structural industry characteristics on the magnitude of the welfare loss from collusive pricing has not been much examined. This is surprising, as there are obvious policy implications from identifying the types of industries where collusion can be expected to be most detrimental for welfare. This paper first presents some theoretical results on this question using a simple model of a differentiated product oligopoly. It then confronts the theoretical predictions with evidence from a unique natural experiment: the abolition of British price-fixing cartels in the 1960s.

A central feature of the model, and one that distinguishes it from most of the existing theoretical literature, is that it allows for free (but costly) entry. This is consistent with the fact that, although entry deterrence by cartel members is possible in principle (Harrington 1989, 1991), most cartels have difficulty in effectively deterring entry in practice (Levenstein and Suslow 2006a). For instance, in a survey of 16 major cartels, Levenstein and Suslow (2004) report that entry occurred and was accommodated in 10 out of the 12 cases for which they had relevant information. This is in line with the fact, discussed in more detail in section 3, that the large majority of the British cartels did not restrict entry. It may be asked why collusion should occur, especially when it is illegal, if profits are to be driven to (almost) zero by entry. One reason is that there are short-run gains from collusion among a given number of firms, even if these gains are eventually eliminated by entry. Furthermore, collusion may persist once it is established, even when it no longer generates supra-normal profits for the average cartel member, because its breakdown will result in short-run losses and even exit for many firms.

The literature on the determinants of market structure has emphasised that, under free entry, collusion generally leads to a less concentrated market structure than competition (Selten 1984, Sutton 1991, Symeonidis 2000a, 2000b, 2002a). This raises the question as to whether welfare results derived from models with a fixed number of firms carry over to models with endogenous market structure. Fershtman and Pakes (2000) have shown that semi-collusion has an ambiguous effect on welfare in a model that allows for free entry. However, they do not examine the case of full collusion or how the welfare implications depend on exogenous industry characteristics. Brander and Spencer (1985) have derived more conventional welfare results in a model of collusion that allows for a two-way link between the number of firms and a reduced-form competition measure (a conjectural variations parameter), but do not analyse how these results vary with industry characteristics.

I adopt a reduced-form theoretical approach in the present paper, aiming to illustrate the main forces at play and derive empirically testable predictions rather than carry out an elaborate analysis of collusion. The firms' choice variable, following the decision to enter or not, is the level of output, and perfect collusion is always achieved irrespective of market structure. The assumption of perfect collusion may seem strong, but it is not crucial. What is probably more important is the exogeneity of collusion, in the sense that the ability of cartels to achieve a higher or lower degree of collusion is not significantly affected by structural industry characteristics. This does not seem implausible in the context of the British cartels of the 1950s, which were legal, long-standing and in most cases operated by industrial trade associations facilitating the coordination, monitoring and enforcement of collusion. I will return to this issue when I describe the model in the next section, and again when I discuss the effectiveness of the British cartels in section 3, and finally in my concluding remarks, where I will argue that relaxing this assumption actually strengthens my results.

In this stylised model collusive pricing unambiguously reduces static welfare, as in most of the literature, and my main objective is to identify the types of industries where these adverse welfare implications are more pronounced. A set of clear results emerge: the welfare loss from collusion is larger the lower the entry cost, the larger the size of the market and the higher the degree of product differentiation. Furthermore, the cartel overcharge, defined here as the difference between the collusive and the competitive price allowing for the endogeneity of market structure, is larger the lower the entry cost and the larger the market size, and is independent of the degree of product differentiation.

The second part of the paper confronts these theoretical predictions with evidence from a natural experiment of policy reform. As a result of the 1956 Restrictive Trade Practices Act in the UK, restrictive agreements among firms, covering a wide range of manufacturing industries, were cancelled. This caused a breakdown of collusion across many industries in the 1960s. However, industries which were already competitive were not affected by the law. Furthermore, and consistent with the model, entry had not been restricted in most collusive British industries. Note that the introduction of cartel law in the UK was an exogenous and measurable institutional change. The inter-temporal structure of the data (spanning a time period both before and after the abolition of cartels) and the exogeneity of the institutional change allow us to largely overcome any concerns about potential biases in the estimated effects of collusion caused by its links with other variables.

Although it is difficult to measure changes in welfare directly, it is possible to use information on the evolution of industry price-cost margins in previously collusive industries either as a proxy for changes in static welfare following the abolition of cartels or as a measure of the cartel overcharge. I find that a relative decline in price-cost margins following the breakdown of cartels occurred in precisely the types of industries where my model predicts the welfare loss from collusion and the cartel overcharge to be larger. Crucially, these effects are not observed in industries not affected by the 1956 legislation, which confirms that they were indeed driven by the breakdown of cartels rather than by some other factor.

A number of recent empirical surveys have reported large positive effects of many cartels on prices and profits (Connor and Bolotova 2006, Connor 2014, Levenstein and Suslow 2006a, Boyer and Kotchoni 2015), although this is by no means a general result and several authors find small or non-existent effects (Asch and Seneca 1976, O'Brien et al. 1979, Symeonidis 2002a). However, the present paper does not seek to address the question of the magnitude of price effects of collusion or the effect of competition on profitability in general, and focuses instead on the related but different question of the structural industry characteristics that determine the magnitude of the welfare effect of cartels. Levenstein and Suslow (2004, 2006) point out that there is considerable variation in the extent to which cartels succeed in raising prices and some of this seems to be driven by idiosyncratic factors. Nevertheless, a small number of previous empirical studies have examined how various systematic factors – such as internal cartel organisation and duration, the geographic region of operation, market concentration, and the legal environment – affect the magnitude of cartel overcharges (Griffin 1989, Connor and Bolotova 2006, Bolotova 2009, Bolotova et al. 2007, 2009). One of the most comprehensive studies is Bolotova (2009), which combines theory and evidence from a cross-section of over 400 cartels going as far back as the 18th century. Bolotova, who uses a model of a homogeneous good industry with a fixed number of firms, finds evidence that cartel overcharges are higher for international compared to domestic cartels, in Europe and the United States compared to the rest of the world, in industries with fewer firms and when outside competition is weak, for older cartels, and in environments where competition policy is not strict.

The present research complements the previous literature in several ways. My theoretical results are derived using a differentiated oligopoly model with free entry that allows for a direct link between theory and evidence. My sample is not large, but it is a panel, it is far more homogeneous than those used in most previous studies, and it is virtually free from selection bias, because it contains all British manufacturing cartels of the 1950s and 1960s and not only those that have been detected or prosecuted. My dependent variable is profitability rather than price. Finally, I focus on a set of factors for whose effect on cartel profitability little is known, namely exogenous industry characteristics such as the level of entry costs, market size and the degree of product differentiation.

There are more aspects to welfare than static allocative efficiency and the price-cost margin. Price collusion may influence product quality and variety, innovation and productivity, and it certainly affects market structure (for evidence on these effects in the context of the British cartels, see Symeonidis 2000a, 2000b, 2002a, 2008a). I discuss this issue and several other extensions or potential concerns in the final section of the paper. I also compare my findings with evidence on the industry characteristics that facilitate cartel formation and conclude that the welfare loss from collusive pricing may be relatively small in the types of industries where collusion most often occurs in practice.

2. The model

Consider an industry producing a potentially infinite number of varieties of a horizontally differentiated product. Competition is described by a two-stage game as follows. There are N_0 potential entrants, each with the capacity to produce a single variety of the product. At stage 1, they decide whether or not to enter at an exogenously given sunk cost of entry *f*. At stage 2, those firms that have entered set quantities. N_0 is sufficiently large, so that at any equilibrium of the game there is at least one non-entering firm. Each firm has a constant marginal cost of production *c*.

Preferences are described by the utility function of a representative consumer

$$U = \sum_i \left(\alpha_1 x_i - \alpha_2 x_i^2\right) - \alpha_2 \sigma \sum_i \sum_{j<i} x_i x_j + M \qquad (1)$$

(see Dixit 1979, Vives 1985, Shaked and Sutton 1990, Sutton 1997, 1998, Symeonidis 2002a, 2002b, among others).[1] The x_i are the quantities demanded of the different varieties of the product in question, while $M = Y - \sum_i p_i x_i$ denotes expenditure on outside goods. This utility function implies that the consumer spends only a small part of her income on the industry's product (which also ensures that the maximisation of U has an interior solution) and hence income effects on the industry under consideration can be ignored and partial equilibrium analysis can be applied. The parameter σ, $\sigma \in (0,2)$, is an inverse measure of the degree of horizontal product differentiation: in the limit as $\sigma \to 0$ the goods become independent, while in the limit as $\sigma \to 2$ they become perfect substitutes. The parameter σ is a basic taste parameter

[1] This demand system was introduced by Bowley (1924). A different type of quadratic utility function, which includes the number of firms as a parameter, was proposed by Shapley and Shubik (1969) and was also used in slightly different form in Shubik (1980). However, the Bowley demand system is a natural choice in the present context, where the number of firms is

which can be seen as an industry-specific measure of the degree to which demand is diversified among users with different preferences or requirements. Finally, α_1, where $\alpha_1 > c$, and α_2 are positive scale parameters.

The inverse demand for variety i is given by

$$p_i = \alpha_1 - 2\alpha_2 x_i - \alpha_2 \sigma \sum_{j \neq i} x_j \tag{2}$$

in the region of quantity spaces where prices are positive. Let N denote the number of varieties offered. This is also the number of firms that have entered the industry at stage 1 of the game.

At stage 2, firms compete by setting quantities. Let each firm choose x_i to maximise $\Pi_i = \pi_i + \lambda \sum_{j \neq i} \pi_j$, where $\pi_i = (p_i - c)x_i$ and the parameter λ can take two values, namely either $\lambda = 0$ (corresponding to the Cournot-Nash equilibrium) or $\lambda = 1$ (corresponding to perfect collusion). In a more general formulation, λ could also take intermediate values between 0 and 1 to represent partial or imperfect collusion by way of a reduced-form parameter (see Shubik 1980, Symeonidis 2002a, 2008b). To keep the welfare analysis simple, I only consider a comparison of the Cournot-Nash and the perfect collusion cases here. The welfare results derived below are also valid if collusion is imperfect, at least when the collusive equilibrium is in the neighbourhood of the joint monopoly outcome. Since I mainly focus on comparing welfare properties of different competition regimes, I will take these regimes as exogenous. The exogeneity of λ with respect to the number of firms N and other parameters of the model is consistent with the well-known multiplicity of possible equilibria in models of infinitely repeated games. Furthermore, it is justifiable in empirical contexts where significant

endogenous and therefore its inclusion as a parameter in the utility function would be inappropriate.

changes in the competition regime occur as a result of exogenous institutional changes such as economic integration or the introduction of effective cartel policy.[2]

Solving the system of N symmetric first-order conditions, we obtain the equilibrium quantity sold and profit of each firm in the second-stage subgame as a function of N. In particular, for the Cournot-Nash equilibrium (λ = 0) we obtain

$$x^C(N) = \frac{\alpha_1 - c}{\alpha_2\left[4 + \sigma(N-1)\right]}, \qquad \pi^C(N) = \frac{2(\alpha_1 - c)^2}{\alpha_2\left[4 + \sigma(N-1)\right]^2}, \tag{3}$$

whereas at the perfectly collusive (joint monopoly) point (λ = 1) we have

$$x^{JM}(N) = \frac{\alpha_1 - c}{2\alpha_2\left[2 + \sigma(N-1)\right]}, \qquad \pi^{JM}(N) = \frac{(\alpha_1 - c)^2}{4\alpha_2\left[2 + \sigma(N-1)\right]}. \tag{4}$$

It can be easily checked that in both cases $\pi(N)$ is decreasing in N for $N > 1$, and that $\pi^{JM}(N) > \pi^C(N)$.

The long-run equilibrium number of firms in the industry, N^*, is determined at stage 1 by the free entry condition $\pi(N^*) = f$, assuming for simplicity that N is a continuous variable. Using (3), (4) and the free entry condition we obtain, for $\lambda = 0$,

$$N^{*C} = 1 + \frac{(\alpha_1 - c)\sqrt{2\alpha_2 f} - 4\alpha_2 f}{\alpha_2 \sigma f} \tag{5}$$

and, for $\lambda = 1$,

$$N^{*JM} = 1 + \frac{(\alpha_1 - c)^2 - 8\alpha_2 f}{4\alpha_2 \sigma f}. \tag{6}$$

[2] In principle, one could think of λ as a function of an exogenous institutional variable and a vector of other variables, which may include N. The present model would then be appropriate for a situation where the dominant influence on λ is a change in the institutional variable, since in such a setup it would be plausible to assume that any feedback effects from N and other variables on λ are small and can therefore be ignored.

In both cases f must be sufficiently low to ensure that $N^* \geq 2$. I will often make use of a weaker restriction on f in what follows, namely

$$0 < f < \frac{(\alpha_1 - c)^2}{8\alpha_2}, \tag{7}$$

which implies $N^* > 1$. Since both $\pi^C(N)$ and $\pi^{JM}(N)$ are decreasing in N and $\pi^{JM}(N) > \pi^C(N)$, it follows that $N^{*C} < N^{*JM}$.[3]

This result raises the possibility that static welfare is higher under collusion because the number of varieties is then larger even though each firm sells a lower quantity than in the Cournot-Nash equilibrium. However, this is not the case here. Because of the free entry zero-profit condition, total welfare is equal to consumer surplus, which is given by

$$\begin{aligned} W^* &= N^*(\alpha_1 x^* - \alpha_2 x^{*2}) - \frac{1}{2}\alpha_2 \sigma N^*(N^*-1)x^{*2} - N^* p^* x^* \\ &= \alpha_2 N^* \left[1 + \frac{1}{2}\sigma(N^*-1)\right] x^{*2}, \end{aligned} \tag{8}$$

where use has also been made of the inverse demand function. Substituting the expressions for $x^* = x(N^*)$ and N^* into W^*, we obtain, for Cournot-Nash behaviour,

$$W^{*C} = \frac{1}{4\alpha_2 \sigma f}\left[(\alpha_1 - c)\sqrt{2\alpha_2 f} - \alpha_2(4-\sigma) f\right]\left[(\alpha_1 - c)\sqrt{2\alpha_2 f} - 2\alpha_2 f\right] \tag{9}$$

and, under perfect collusion,

$$W^{*JM} = \frac{(\alpha_1 - c)^2 - 4(2-\sigma)\alpha_2 f}{8\alpha_2 \sigma}\ . \tag{10}$$

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[3] In a more general formulation where λ can take any value between 0 and 1, it is easy to check that $\pi(N)$ is decreasing in N and increasing in λ. It follows from the total differential of the free entry condition that, for any $\lambda \in [0,1)$, $dN^*/d\lambda > 0$, i.e. the equilibrium number of firms increases with the degree of collusion. For $\lambda = 1$, $\pi(N)$ is maximised, and so is N^*.

Thus:

$$\Delta W^* = W^{*C} - W^{*JM} = \frac{1}{8\alpha_2\sigma}\left[(\alpha_1 - c) - 2\sqrt{2\alpha_2 f}\right]\left[3(\alpha_1 - c) - 2(3-\sigma)\sqrt{2\alpha_2 f}\right], \quad (11)$$

which is a positive expression, given (7). We can therefore conclude:

Proposition 1. *Static welfare is lower under perfect collusion than in the Cournot-Nash equilibrium.*

We can now examine how changes in the parameters of the model affect the magnitude of the welfare loss from collusion. We have:

$$\frac{\partial \Delta W^*}{\partial f} = \frac{-\left[(\alpha_1 - c)(6-\sigma)\sqrt{2\alpha_2 f} - 8(3-\sigma)\alpha_2 f\right]}{8\alpha_2\sigma f} < 0, \quad (12)$$

$$\frac{\partial \Delta W^*}{\partial(\alpha_1 - c)} = \frac{3(\alpha_1 - c) - (6-\sigma)\sqrt{2\alpha_2 f}}{4\alpha_2\sigma} > 0, \quad (13)$$

$$\frac{\partial \Delta W^*}{\partial \alpha_2} = \frac{-(\alpha_1 - c)\left[3(\alpha_1 - c) - (6-\sigma)\sqrt{2\alpha_2 f}\right]}{8{\alpha_2}^2\sigma} < 0 \quad (14)$$

and

$$\frac{\partial \Delta W^*}{\partial \sigma} = \frac{-3\left[(\alpha_1 - c) - 2\sqrt{2\alpha_2 f}\right]^2}{8\alpha_2\sigma^2} < 0, \quad (15)$$

for $\sigma \in (0,2)$ and given (7). To summarise:

Proposition 2. *The welfare loss from collusion is larger the lower the entry cost f, the larger the size of the market (i.e. the larger the value of α_1 or the smaller the value of α_2), and the higher the degree of product differentiation (the smaller the value of σ).*

The intuition for Proposition 2 hinges on the way changes in the parameters of the model affect the equilibrium number of firms and total quantity sold in the Cournot-Nash equilibrium and under collusion. Thus, for instance, a fall in f causes the number of firms and total quantity sold to rise under both regimes. However, a fall in f increases $N^{*JM} - N^{*C}$ and $x^{*C}N^{*C} - x^{*JM}N^{*JM}$, i.e. it makes the outcomes under the two regimes more dissimilar. Note that x^*N^* is more sensitive to changes in f under Cournot behaviour than under collusion even though the opposite is the case for N^*. This is because the overall effect of a change in f on total quantity sold is dominated by the effect on the quantity sold by each firm (which is itself also dependent on N) rather than the direct effect on the number of firms. As a result, a fall in f causes the difference between $x^{*C}N^{*C}$ and $x^{*JM}N^{*JM}$ to increase even though it has a weaker effect on N^{*C} than on N^{*JM}. As the difference between $x^{*C}N^{*C}$ and $x^{*JM}N^{*JM}$ increases, so does the difference between W^{*C} and W^{*JM}.

Note that the negative effect of f on the welfare loss from collusion is *not* a mere consequence of the fact that (i) a fall in f increases N^* and (ii) the welfare loss from collusion increases in N. The reason is that the effect of f on N^*, although always negative, is larger in absolute value under collusion than in the Cournot-Nash equilibrium. In other words, as f falls, N^{*C} increases by less than N^{*JM}. Therefore, although W^{*C} and W^{*JM} will both rise, it is not clear which will rise by more. As it turns out, a fall in f increases the difference between W^{*C} and W^{*JM}, even though it raises N^{*C} by less than N^{*JM}. This point is important because it helps to dispel a potential criticism of the theoretical results derived here, namely that it may be "obvious" that the welfare loss from collusion will be greater in a market with more firms (as when market size is larger, the product more differentiated, or the entry cost lower), because the collusive outcome is always the same whereas the Cournot outcome becomes more competitive as the number of firms increases. This criticism is not valid because of free entry: the direction of the change in ΔW^* cannot be known a priori when ΔN^* also changes.

Finally, I calculate how changes in the parameters of the model affect the difference between the joint monopoly price and the Cournot price, and therefore also the difference in the respective profit margins, when we allow for free entry. From (3)-(5) we obtain

$$p*^{C} = c + \sqrt{2\alpha_2 f}, \tag{16}$$

$$p*^{JM} = c + \frac{(\alpha_1 - c)}{2}, \tag{17}$$

and therefore

$$\Delta p* = p*^{JM} - p*^{C} = \frac{(\alpha_1 - c)}{2} - \sqrt{2\alpha_2 f}, \tag{18}$$

which is positive, because of (7). It is easy to verify:

Proposition 3. *The difference between the collusive price and the Cournot price is larger the lower the entry cost f, the larger the size of the market (i.e. the larger the value of α_1 or the smaller the value of α_2), and is independent of the degree of product differentiation.*

Note that the negative effect of market size on $\Delta W*$ obtained earlier is not a trivial scale effect; an increase in market size raises the cartel overcharge $\Delta p*$ too.

Most of these results are likely to carry over to more general settings, except for the independence of $\Delta p*$ from σ, which may be specific to the linear demand model. It is easy to check that the difference between the total Cournot quantity sold and the total collusive quantity sold decreases in σ, but this is exactly balanced by the effect of σ on $\Delta N*$ to yield the independence result for $\Delta p*$. One way to interpret this finding is to say that the degree of product differentiation could have an ambiguous effect on the cartel overcharge in more general settings, but that any such effect would not be strong and might be difficult to detect empirically.[4]

3. Data and empirical methodology

Explicit restrictive agreements among firms were widespread in British industry in the mid-1950s: nearly half of manufacturing industry was subject to price fixing. The agreements were not enforceable at law, but they were not illegal. Most of them provided for minimum or fixed producer prices. There were generally no restrictions on longer-term strategic decisions such as investment in capacity, advertising or R&D. A description of the institutional changes and the evolution of competition from the 1950s to the early 1970s and a detailed survey of restrictive agreements across all British manufacturing industries can be found in Symeonidis (2002a). Here I summarise the evidence and describe the construction of the dataset for this paper.

The 1956 Restrictive Trade Practices Act required the registration of restrictive agreements, including verbal or implied arrangements, on goods. Registered agreements should be abandoned, unless they were either successfully defended in the Restrictive Practices Court as producing benefits that outweighed the presumed detriment or cleared by the Registrar of Restrictive Trading Agreements as not significantly affecting competition. Because the attitude of the Court could not be known until the first cases had been heard, the large majority of industries registered their agreements rather than dropping or secretly continuing them. The first agreements came before the Court in

[4] The independence of Δp^* from σ also depends on the assumption that λ is exogenous, as previously discussed. In circumstances where a higher degree of product differentiation implies a lower degree of collusion, the independence result for Δp^* would not hold. A higher degree of differentiation (a smaller value of σ) would then be associated with a decrease in the cartel overcharge.

1959 and were struck down. This induced most industries to abandon their agreements rather than incur the costs of a Court case with little hope of success.

Was entry generally free in collusive industries? The evidence from the agreements registered under the 1956 Act, several industry reports published by the Monopolies and Restrictive Practices Commission during the 1950s and a large number of case studies discussed in Swann et al. (1973, 1974) suggests that entry was in most cases free, i.e. only subject to standard entry costs that would also apply in the absence of collusion. Most agreements were operated by industrial trade associations and there were often no significant restrictions on association membership, so that entry would not be difficult if the entrant was willing to become a party to the agreement. For instance, only for four out of 14 collusive industries investigated by the Monopolies and Restrictive Practices Commission in the 1950s, which operated some of the most restrictive agreements across all British industries, did the Commission report evidence of restricted cartel membership.[5] Furthermore, the reports of the Monopolies Commission and the case studies in Swann et al. contain information on profitability of firms in collusive industries: profits were more often described as "reasonable" than "excessive", and this was due to the fact that cartel firms could not defend excessive profits against entry. Finally, Symeonidis (2000a, 2002a) has documented that a significant restructuring occurred in previously

[5] Free cartel membership is not the same as free entry into the collusive industry. In many cases, however, the former implies the latter because many of the entry deterrence mechanisms used by cartels, such as collective exclusive dealing or aggregated discounts to distributors, are essentially attempts to restrict competition from outside firms and are not very effective if cartel membership is free. In fact, such practices were used by several British cartels, but they were often not effective in deterring entry. Part of the reason for this may be the fact that although in some industries the existing association members might reject certain applications for

collusive British industries as a result of the 1956 Act, with mergers and exit of firms causing, on average, the five-firm concentration ratio to increase by about 6 percentage points and the number of firms to fall by about 15% between 1958 and the mid-1970s. It would be difficult to explain such a strong impact of cartel policy on market structure if cartel firms had generally been able to maintain high profits by deterring entry, since in that case many more firms would have been able to survive albeit with reduced profitability rather than be driven to exit.

Were the agreements effective? Or could the lack of excessive profits be simply due to ineffective cartels rather than free entry? The effectiveness of an agreement depended on two factors: the extent to which the parties themselves conformed to it and the extent of competition from outside firms, domestic or foreign. Evidence from the registered agreements, the Monopolies and Restrictive Practices Commission reports, the Political and Economic Planning (1957) survey of industrial trade associations and Swann et al. (1973, 1974) suggests that in most industries the agreements had been operated effectively prior to cancellation, the parties typically accounted for a large fraction of the market and contained the largest and best-known domestic firms, and outside competition was usually weak. For instance, Swann et al. (1974) report cartel market shares of 90% or higher in about two thirds of the cases they examine, and 75% or higher in all but two cases. Competition from imports was often limited as a result of tariffs and quantitative controls, differing technical standards, transport costs or international restrictive agreements. Finally, the legality

membership, they would normally accommodate any powerful non-member firm rather than face strong outside competition.

of the agreements and the institutional role of the trade associations that operated them had facilitated the coordination, monitoring and enforcement of collusion.[6]

To what extent did collusion break down following the abolition of cartels? Evidence from various sources indicates that price competition intensified in the short run in many industries. However, in many others, agreements to exchange information on prices, price changes and so on replaced the former explicit collusive arrangements, and price competition emerged only after these information agreements were abandoned in the mid-1960s, following adverse decisions of the Restrictive Practices Court. Price wars occurred in a number of previously collusive industries in the second half of the 1960s, and the final blow came with the provisions of the 1968 Restrictive Trade Practices Act regarding information agreements.[7] In many industries, therefore, competition emerged more than a decade after the introduction of the 1956 legislation. Overall, sooner or later the large majority of industries with restrictive agreements in the 1950s did experience a breakdown of collusion as a result of the 1956 Act.

Although my main source of data on competition are the agreements registered under the 1956 Act, I also use other sources to identify unregistered agreements, including the industry reports of the Monopolies Commission, the 1955 Monopolies

[6] Genesove and Mullin (2001) and Levenstein and Suslow (2006b) describe other instances where communication and monitoring within trade associations has facilitated collusion.

[7] Information agreements were not registrable under the 1956 Act, unless it could be shown that they were being used as a way to continue a formally abandoned explicit collusive scheme. In two important cases brought before the Restrictive Practices Court in the mid-1960s, prices and price changes had consistently been notified in advance of coming into effect, and the prices of different firms had always been identical. The Court concluded that these agreements amounted to the same effect as explicit price fixing, and ruled against the parties. These Court rulings induced several more industries with information agreements to cancel them. A few years later, the 1968 Restrictive Trade Practices Act made certain types of information agreements registrable. By the time they were called up for registration, most had already been abandoned.

Commission report on collective discrimination, the 1949 report of the Lloyds' Committee on resale price maintenance, industry studies contained in Burn (1958) and Hart et al. (1973), the Board of Trade annual reports from 1950 to 1956, and the Political and Economic Planning (1957) survey of trade associations (including unpublished background material for this survey). The use of a diverse range of sources guarantees that any potential measurement error caused by ineffective agreements or unknown cases of collusion in the data is very small.

The data on profitability and other variables are available for industries defined at the three-digit level of aggregation (i.e. "minimum list heading" industries of the UK Census of Production), although I have sometimes used available data at the four-digit industry level. The three-digit level of aggregation defines a total of about 130 UK manufacturing industries and has been often used in studies with British or international data on profitability (for instance, Cowling and Waterson 1976, Sleuwaegen and Yamawaki 1988) or productivity growth (Symeonidis 2008a). All manufacturing industries were classified as collusive, competitive or ambiguous according to their state of competition in the 1950s on the basis of three criteria: the reliability of the data source; the types of restrictions; and the proportion of an industry's total sales revenue covered by products subject to agreements and, for each product, the fraction of the UK market covered by cartel firms. In particular, the various types of restrictions were classified as significant, not significant or uncertain, according to their likely impact on competition. Next, the products that were subject to agreements were assigned to the industry categories used. An industry was classified as collusive in the 1950s if the products subject to *significant* restrictions accounted for more than 50% of total industry sales. It was classified as competitive if the products subject to *significant or uncertain* restrictions accounted for less than 20% of industry sales. I have used the 50% cut-off point because in some cases most core industry products were subject to price fixing, although some were not; clearly, one would expect a significant impact of the 1956 Act in such cases. I have used the 20% cut-off point because in some cases secondary industry products were subject to restrictive agreements, although core products were not. Variations in these cut-off points (for instance, using 60% instead of 50%, or 10% instead of 20%) do not significantly affect the results.[8] I have excluded a few industries which remained collusive or partially collusive throughout the period under study and those with significant government participation or intervention.

This procedure resulted in a panel consisting of 36 industries, listed in Appendix 1, with a clear change of competition regime and five years: 1954, 1958, 1963, 1968 and 1973. The first four of these are the only ones in the 1950s and 1960s for which comparable data on price-cost margins are available, and 1973 is the last year before the oil crisis of the 1970s. I will also use for comparison a control group of 55 industries which were competitive in the 1950s and therefore were not significantly affected by the 1956 law. Note that although the cartel law was introduced in 1956, it was not until 1959 that industries started cancelling their agreements. And since competition was often slow to emerge, as pointed out above, the 1956 Act had an impact well into the late 1960s and even early 1970s.

It is difficult to measure changes in welfare directly, but it is possible to use information on the evolution of price-cost margins across different classes of previously collusive industries either as a proxy for the change in static welfare

[8] In fact, out of 36 industries classified as collusive 20 had agreements covering all or nearly all industry products and most of the others had agreements covering more than 75% of total industry sales. The use of a continuous competition measure instead of cut-off points has proved impractical for a variety of reasons (see Symeonidis 2002a).

following the breakdown of cartels or as a measure of the cartel overcharge. Note that the cartel overcharge is usually defined as the difference between the collusive and the competitive price, whereas I focus in this and the following section on the difference between the collusive and the competitive price-cost margin. However, the qualitative results will always be similar for these two measures, and the quantitative results should also be similar for the *proportional* change in the price and the price-cost margin, at least if marginal cost is largely exogenous. Furthermore, in practice neither the industry price index nor the quantity sold would be a valid measure of welfare over a 20-year period, because changes in the former might largely reflect exogenous changes in costs, while the latter would be mainly driven by demand. I follow the standard definition of the price-cost margin, *PCM*, in the literature, as the net value of output (or value added) minus wages and salaries divided by sales revenue (see, for instance, Collins and Preston 1969, Cowling and Waterson 1976, Machin and Van Reenen 1993). *PCM* was constructed from Census of Production data and any limitations of these are greatly alleviated here by the use of panel data and the fact that my results are based on changes in *PCM*, not levels.

The industry capital-labour ratio, *K/L*, is the best available proxy for the level of entry cost: since the entry cost is primarily the cost of installing capital (plant and machinery), capital-intensive industries are typically those with significant entry costs and vice versa. For product differentiation, I will use two different empirical measures: the advertising-sales ratio and a consumer good-producer good indicator. Both have often been used as proxies for product differentiation in the empirical industrial organisation and international trade literatures, including studies of collusion. Finally, the industry sales revenue is the best available measure of market size. It may be argued that to the extent that some industries comprise a number of submarkets that are largely independent, sales revenue may be affected by the degree of within-industry heterogeneity. The list of industries in Appendix 1 comprises a few to which this objection may apply; note that any measurement error might only cause the estimated coefficients to understate somewhat the magnitude of the effect of market size on *PCM*.

To mitigate potential endogeneity and measurement error problems and to facilitate the interpretation of the results, I will use the primary data to construct a set of binary variables designed to split the sample of industries into categories or groups according to their characteristics. My motivation is as follows. Take, for instance, the advertising-sales ratio: although it is endogenous, it is largely exogenous characteristics that will determine whether it is above or below 1% in any given industry. Thus for an industry below the 1% cut-off point there is not much scope for using advertising to promote a differentiated product. The opposite is true for an industry above the 1% cut-off point. Of course, whether in such a case the advertising-sales ratio will be 3% or 5%, say, may be largely determined endogenously. But a binary variable is not very sensitive to endogenous factors that affect advertising intensity. In fact, a comparison of advertising-sales ratios across various years reveals very few instances in my dataset where an industry moves from below 1% to above 1% or vice versa. A similar argument can be made with respect to the capital-labour ratio or sales revenue. Measurement error is an additional issue both for sales revenue as a proxy for market size, as discussed above, and for the capital-labour ratio: the capital stock figures are estimates rather than primary data, more reliable for measuring changes than levels (see Oulton and O'Mahony 1990), and a binary variable may help reduce the effect of measurement error on the econometric estimates.

I construct four variables: *LOWCAPINT*, which is equal to 1 for industries with 1958 capital-labour ratio lower than the average across all manufacturing industries and

0 otherwise; *ADV*, which is equal to 1 for industries with 1958 advertising-sales ratio higher than 1% and 0 otherwise; *PRODCON*, which takes the values 0 for industries producing primarily producer goods and 1 for industries producing primarily consumer goods; and *MKTSIZE*, which is equal to 1 for industries with 1958 sales revenue higher than the average across all manufacturing industries and 0 otherwise. Each of these variables takes the value of 0 or 1 for each industry, not for each individual observation (industry-year pair). The year 1958 was chosen because it is the last in my sample when all the cartels were still in place, and the scope for a decline in *PCM* in any given industry must depend on initial conditions. The classification of industries would change very little if another year were chosen instead or if the median capital-labour ratio and sales revenue were used rather than the mean to construct *LOWCAPINT* and *MKTSIZE*, respectively. Details on the classification of industries are provided in Appendix 1 and on data sources in Appendix 2.

Following the theoretical results of section 2, I expect *PCM* to unambiguously fall after the breakdown of collusion in industries where *LOWCAPINT* and *MKTSIZE* take the value 1 relative to industries where they take the value 0. I also expect no change or a modest fall in *PCM* in industries where *ADV* and *PRODCON* are equal to 1 relative to industries where these variables are equal to 0.

Descriptive statistics on initial levels of *PCM* are presented in Table 1 for collusive as well as competitive industries and also for groups of collusive industries according to their capital intensity, advertising intensity, producer good-consumer good index and market size. There are no significant differences across the various rows of Table 1, and this applies even more to the 1954-1958 change in *PCM* than to its level. One should not read too much into a comparison of the levels anyway. For instance, the fact that *PCM* is slightly lower, on average, in collusive than in competitive industries is partly due to the fact that low-advertising and low-R&D industries had, on the whole, lower price-cost margins *and* higher incidence of collusion than advertising-intensive and R&D-intensive ones. Much more informative for our purposes are the 1954-1958 changes in *PCM*. One may conclude, on the basis of these figures, that any differences observed after 1958 across industry groups should be attributed to the 1956 cartel law only and are not biased by any divergent pre-existing trends.

4. Econometric evidence

The econometric specifications used in this paper are panel data models with individual-specific (industry) effects. In line with the theory, my basic specification for *PCM* is

$$\begin{aligned}\ln PCM_{it} = {} & \alpha_i + \beta_1 \ln(K/L)_{it} + \beta_2 Y63 + \beta_3 Y68 + \beta_4 Y73 \\ & + \beta_5 LOWCAPINT \times Y63 + \beta_6 LOWCAPINT \times Y68 + \beta_7 LOWCAPINT \times Y73 \\ & + \beta_8 ADV \times Y63 + \beta_9 ADV \times Y68 + \beta_{10} ADV \times Y73 \\ & + \beta_{11} MKTSIZE \times Y63 + \beta_{12} MKTSIZE \times Y68 + \beta_{13} MKTSIZE \times Y73 + u_{it}.\end{aligned}$$

I use a logarithmic transformation of *PCM* to facilitate the interpretation of the estimated coefficients as the proportional change in the cartel overcharge, for the reasons discussed in the previous section. However, I have also performed robustness checks using the untransformed *PCM* as dependent variable. The year dummies, *Y63*, *Y68* and *Y73*, are meant to capture various factors that have affected price-cost margins in this sample of industries over the period examined, in addition to the breakdown of collusion. These include cyclical fluctuations, the progressive opening of the British economy, the UK government's prices and incomes policies in the 1960s and macroeconomic policy. It is difficult to assess these factors at the industry level, but it may be plausible to assume that their effect would have been more or less equally realised across all industries, or at least that there would be no systematic difference across classes of industries according to their capital intensity, degree of product differentiation or market size.

One potentially important determinant of *PCM* that I am able to control for is the capital-labour ratio, *K/L*: I expect a positive effect of this on the price-cost margin. The capital-labour ratio has often been used in profitability studies to control for the fact that the endogenous variable, the price-cost margin or the rate of return on capital, includes the gross return to capital. In the present study the capital-labour ratio is seen as a proxy for the cost of entry. This is not a real difference, however, since the entry cost is primarily the cost of installing capital (plant and machinery).

The coefficients of primary interest are those on the interaction terms. The coefficient on $LOWCAPINT \times Y73$ will test whether the change in *PCM* between the 1950s and 1973 was different in previously collusive industries with $LOWCAPINT = 1$ (i.e. relatively low-capital industries) than in industries with $LOWCAPINT = 0$, after controlling for the other exogenous variables. The coefficient on $ADV \times Y73$ will compare the evolution of *PCM* between previously collusive advertising-intensive industries and low-advertising ones. Finally, the coefficient on $MKTSIZE \times Y73$ will reveal whether the change in *PCM* was different in previously collusive industries with $MKTSIZE = 1$ (i.e. larger markets) than in those with $MKTSIZE = 0$.

My approach is not meant to say anything about the *level* of price-cost margins. Rather I examine differences in the *evolution* of price-cost margins across classes of industries. My benchmark model does not include time dummies or interaction terms for either 1954 or 1958, the two years in the dataset when the collusive agreements were still in place. This approach helps to gloss over any fluctuations in *PCM* over 1954-1958 – which, as shown in Table 1, are essentially random noise and not related to any effects of collusion – and should therefore provide an accurate estimate of the long-run effect of competition on *PCM*. However, I will also present results using 1958 as the benchmark year to bring more sharply into focus the short-run effect of the breakdown of cartels. On the other hand, distinguishing between 1963, 1968 and 1973 is important: since competition emerged slowly in many industries, it is best not to impose any structure on the data regarding the timing of the impact of the law on *PCM*.

Table 2 contains the results from fixed effects regressions with robust standard errors. As the Wooldridge test for serial correlation in panel data sometimes marginally rejects the null hypothesis of no first-order autocorrelation, alternative results obtained from a model with AR(1) disturbances are presented for some specifications. I begin with a regression of ln*PCM* on year dummies and the control variable ln(*K/L*). The coefficient on ln(*K/L*) is positive and statistically significant, as expected, a result that holds across all specifications. The second and third columns report the results from my benchmark specification, while the fourth column also includes ln*NFIRMS* among the regressors – where *NFIRMS* is the number of firms, the best available measure of market structure at the three-digit industry level over the period 1954-1973.[9] The fifth column adds the year dummy and interactions for 1954. All these regressions provide clear evidence of a fall in the long run in the average price-cost margin of industries where *LOWCAPINT*, *ADV* and *MKTSIZE* take the value 1 relative to industries where these variables take the value 0: the coefficients on $LOWCAPINT \times Y73$, $ADV \times Y73$ and $MKTSIZE \times Y73$ are all negative and statistically significant at the 1% or 5% level.

[9] My specification is derived from a theoretical model in which the number of firms and the price-cost margin are both endogenous. However, I also report results from a regression that controls for market structure both as a robustness check and in order to obtain an indication of whether the welfare effects of collusion may be different in circumstances where entry is restricted and the number of firms fixed. The coefficient on ln*NFIRMS* can be interpreted as a correlation coefficient, although instrumenting ln*NFIRMS* with the log of deflated sales revenue gives very similar results.

Furthermore, the estimated effects are large, indicating average decreases in *PCM* of at least 15% in low-capital and larger-sized relative to capital-intensive and smaller-sized industries, as well as an average decrease in *PCM* of up to 15% in advertising-intensive relative to low-advertising industries. The last three columns replace *ADV* with *PRODCON*. Most of the results are similar, but the coefficients on the interaction terms with *PRODCON* are small and not statistically significant.

The coefficients on the interaction terms (excluding those with *PRODCON*) are statistically significant for 1973 but usually not for 1963 or 1968. Several factors may help explain this result. As pointed out in section 3, competition took many years to emerge in a significant number of collusive industries and information agreements were often used to maintain implicit price fixing until the late 1960s. Import competition also intensified in the second half of the 1960s and the early 1970s, and this must have put pressure on firms in industries where domestic competition was still weak because of a history of collusive arrangements. Furthermore, several important changes during the 1960s in the distribution of manufactured goods – such as the abolition of resale price maintenance following the introduction of the 1964 Resale Practices Act, the growth of large powerful retailers, and the rise of retailers' own brands – reduced the role of brand advertising and the bargaining power of manufacturers and put pressure on their price-cost margins (see Pickering 1974, Ward 1973, Mercer 2013). Resale price maintenance, in particular – which affected more than a third of consumer expenditure on goods in 1964 and was not only prevalent in consumer goods but also existed across a range of intermediate products – must have delayed or hindered the emergence of effective price competition among manufacturers in several previously collusive advertising-intensive industries.[10] Its abolition in the second half of the 1960s may be an important factor behind the fall in *PCM* after 1968 in these industries. All in all, it is not surprising that the divergence in the evolution of *PCM* across classes of previously collusive industries is more pronounced in 1973 than in 1963 or 1968: the year 1973 may be the only unambiguously competitive one for many industries in the dataset.

An intriguing feature of the results in Table 2 is the positive sign and statistical significance of the coefficients on the year dummies. This seems counterintuitive, but there are several factors that help to provide an explanation. First, as I have pointed out already, the capital stock figures are estimates rather than primary data. Since *K/L* was increasing across all industries throughout the period, it is correlated with the year dummies. So to the extent that there is measurement error in *K/L*, the effect of capital intensity on *PCM* is partly picked up by the year dummies. This effect is large: in regressions excluding ln(*K/L*), the magnitude of the coefficients on the time dummies more than doubles compared to regressions where ln(*K/L*) is included. Second, various economy-wide influences are captured by the year dummies and some of these may have been pushing price-cost margins upwards over time. Cyclical fluctuations could be one such factor – for instance, the British economy experienced a slowdown in 1958, whereas 1963 and 1968 were boom years. Third, it is by no means clear that there should be a strong negative effect of the breakdown of collusion on price-cost margins in the long run in industries where free entry and endogenous changes in market structure tend to drive profits (or profits of the least efficient firms) to zero irrespective of the competition regime. A full analysis of this point is beyond the scope of the

[10] In fact, a secondary provision of the 1956 Restrictive Trade Practices Act was to strengthen manufacturers' powers to enforce individual resale prices: this was seen at the time as an attempt to maintain a balance between large and small manufacturers in a context where the collective enforcement of prices was no longer possible.

present work, but the evidence reported in Symeonidis (2002a) suggests that a significant restructuring took place by way of mergers and exit of firms in previously collusive industries during the 1960s, which has tended to restore price-cost margins in the long run, at least on average.[11]

Last but not least, the overall rise in industry profitability in the 1960s in my dataset may be more apparent than real. A preliminary analysis of the data revealed a large increase of the average reported *PCM* between 1958 and 1963. In particular, the average change in *PCM* for the 36 collusive industries was 0.03 in 1954-1958, 0.23 in 1958-1963, 0.13 in 1963-1968 and 0.01 in 1968-1973. The corresponding figures for all industries are 0.02, 0.29, 0.06 and 0.01. The 1958-1963 jump in the reported *PCM* is difficult to explain by reference to any institutional or other exogenous factors and may be due to an unknown change in the way information was collected and certain variables (especially net output) computed for the 1963 and later Censuses as compared to the 1958 and earlier Censuses – in which case it can be regarded as a spurious but largely uniform shift across industries.[12] More generally, although some of the changes in the reported *PCM* may be due to measurement error and factors not explicitly included in my specification, any such factors can be plausibly assumed to be uncorrelated with the structural industry characteristics I consider here, and therefore

[11] Furthermore, price-cost margins may have increased during the 1960s and early 1970s in some previously collusive industries for reasons not captured by the symmetric cost model of section 2. In particular, more intense competition may have caused low-cost firms to expand at the expense of high-cost rivals; and since low-cost firms have higher profit margins than high-cost firms, this could have led to an increase in overall profitability in industries with large efficiency differences among firms.

[12] There is no such jump in UK profitability estimates based on company accounts – see, for instance, the statistics published in the November 1974 issue of *Economic Trends*.

would not affect my comparisons of different classes of industries, as described by the coefficients on the interaction terms.

I performed a variety of robustness checks, some of which are presented in Table 3. These included: 1) replacing ln*PCM* by *PCM* as dependent variable in the first two columns of Table 3; 2) in the next three columns, classifying the industries as collusive when the products subject to significant restrictions accounted for more than 60% of total industry sales (instead of 50%); 3) replacing the capital-labour ratio, *K/L*, by the capital stock of the average plant as a measure of entry costs; 4) controlling for R&D intensity interactions with the time dummies; and 5) adding the logs of deflated sales revenue, ln*SALES*, average plant size, ln*SIZE*, and union density, ln*UNION* (taken from Bain and Price 1980), as additional regressors in the last two columns of Table 3 – all three statistically insignificant. The results from these alternative specifications were broadly similar to those in Table 2 except that the coefficient on $ADV \times Y73$ was usually no longer statistically significant at the 5% level. All in all, the econometric findings are consistent with the theory, which predicts an unambiguously positive effect of low entry costs and large market size and a small or ambiguous effect of product differentiation on the cartel overcharge.

One final question may be asked: can we be certain that the effects described by the interaction terms are driven by the breakdown of collusion rather than by some other unspecified factors that apply not only to my sample of 36 previously collusive industries but to all industries? This concern is particularly relevant in light of the observed timing of the competition effect in the data, with much of the impact appearing to occur many years after the introduction of the cartel legislation, as well as the concerns about potential measurement error in *PCM*.

To address this issue I have performed a similar set of regressions for a control group of industries that did not experience a change of competition regime in the 1960s because they had never been subject to restrictive agreements to any significant degree. This sample consists of 55 industries classified as competitive according to the criteria described in section 3. The results from random effects regressions (the preferred model in most cases according to a standard Hausman test – with the results from fixed effects regressions being very similar) are collected in Table 4. The coefficients on the interaction terms are very different in this group compared to collusive industries. None of them is negative and statistically significant, and in fact most have the opposite sign than in Tables 2 and 3. I interpret this as strong evidence that the results described by the coefficients on the interaction terms in Tables 2 and 3 are indeed driven by the breakdown of cartels.[13]

5. Discussion

In a model of a differentiated product market with linear demand and free entry, collusion reduces static social welfare and the welfare loss is larger the lower the entry cost, the larger the size of the market, and the higher the degree of product

[13] A comparison of Tables 2 and 4 suggests that if one were to argue that differences between collusive and competitive industries are a more accurate measure of the impact of the breakdown of cartels on *PCM* than the coefficients in Table 2, the case for a negative effect of entry costs and a positive effect of market size on the cartel overcharge would be strengthened, whereas the case for an effect of product differentiation on the cartel overcharge would be weaker. As a further robustness check I estimated regressions with the entire set of collusive and competitive industries using a difference-in-differences approach. To this end I defined a binary variable which takes the value 1 for industries with a change of competition regime and 0 otherwise and used two-way and three-way interaction terms to capture differences across classes of industries according to their structural characteristics as well as differences between collusive and competitive industries. The results confirmed my earlier findings.

differentiation. The cartel overcharge is larger the lower the entry cost and the larger the market size, and is independent of the degree of product differentiation. The econometric estimates suggest average decreases in price-cost margins following the breakdown of collusive pricing in 1960s Britain of at least 15% in low-capital and larger-sized relative to capital-intensive and smaller-sized industries, respectively. There is weak evidence of a price-cost margin fall in advertising-intensive and consumer good relative to low-advertising and producer good industries. Note that the somewhat mixed evidence on the effect of product differentiation on the cartel overcharge is consistent with an unambiguously positive effect of product differentiation on the welfare loss from collusion.

Admittedly, my theoretical model is stylised and the econometric analysis subject to data limitations that could somewhat "blur" the results – in particular, I cannot capture variations in the "degree of collusion" across industries or differences in the exact timing of cartel breakdown, since these cannot be determined from the available information. Nevertheless, the two sets of results, theoretical and empirical, are remarkably consistent with one another, and the contrast between the evolution of price-cost margins in collusive and competitive industries provides further support for my conclusions.

One might ask whether the theoretical results obtained from the present approach, which assumes free entry, are different from those that would be obtained if the number of firms were fixed. One reason why this is an important question is that short-run profits may be driving the decision of firms to collude under free entry, as pointed out in the Introduction. Another reason is that entry is sometimes restricted either by institutional factors or by the strategic behaviour of cartel firms. And still another reason is that some cartels, unlike the British ones of the 1950s, are either short-lived or subject to frequent price wars, and in such circumstances high profitability during collusive periods may not attract entry.

It is easy to check that when the number of firms is fixed in the model of section 2, both the cartel overcharge and the welfare loss from collusion increase in the number of firms and the size of the market. Note that to the extent that industries with low entry costs are less concentrated, lower entry costs will be associated with larger welfare losses from collusion even in the short run or when entry is restricted. However, the model with fixed number of firms also predicts that the welfare loss from collusion is larger for intermediate levels of product differentiation, unlike the free entry case. On the other hand, the econometric results are not at all sensitive to controlling for market structure, suggesting that the distinction between free and restricted entry may not be crucial in practice. All in all, from an antitrust perspective, most of my findings, and especially those on the role of entry costs and market size, are likely to be also of relevance for cartels operating in different circumstances than the British cartels of the 1950s.

A potential concern with the econometric analysis is that the price-cost margin may be an imperfect proxy for welfare. In particular, differences in the evolution of the price-cost margin across industries would not accurately reflect welfare differences if they were mainly driven by endogenous changes in productivity or labour costs rather than price and industry sales. To check whether my empirical results might be subject to this problem, I have run regressions similar to those presented in Table 2 but with labour productivity and real wages rather than *PCM* as dependent variables. None of the interaction terms in these regressions was statistically significant. In other words, any changes in productivity or labour costs caused by the breakdown of British cartels were *not* more pronounced in certain types of industries than in others. It follows that the observed differences in the evolution of *PCM* across types of industries in my data accurately reflect welfare differences.

I should emphasise that the theoretical and empirical results of this paper relate to static welfare only. In assessing how the overall welfare implications of collusion may differ across industries, one would need to take into account the production and use of innovations, non-price variables such as product quality and variety, productivity and market structure. This task is beyond the scope of the present research. The available evidence from productivity regressions, as reported above and in Symeonidis (2008a), suggests that the British cartels slowed down labour productivity growth in the 1950s and that this effect was equally strong in all classes of industries. Furthermore, collusion led to excess entry (Symeonidis 2000a, 2000b, 2002a) and the 1956 law caused a significant rise in concentration and fall in firm numbers in low-advertising/R&D industries as well as in advertising-intensive and R&D-intensive industries.

Another potential concern is that the assumption of perfect collusion may not be realistic. This will not matter much if the degree of collusion varies across industries due to idiosyncratic factors, without being correlated with any of the structural industry characteristics examined here. The assumption of perfect collusion is not crucial for the theoretical results provided the degree of collusion is exogenous with respect to the parameters of the model. For the econometric analysis, what is important is that the ability of cartels to raise prices above competitive levels is not correlated with exogenous industry characteristics. If, however, such a correlation existed, then the empirical results could be difficult to interpret. For instance, would the relative decline of price-cost margins in previously collusive low-capital industries relative to capital-intensive ones that we observe in the data reflect the mechanism proposed in the theoretical model, i.e. the higher welfare loss from perfect collusion? Or would it rather result from the ability of cartels operating in low-capital industries to raise prices closer to the joint monopoly level than cartels operating in capital-intensive industries?

I cannot conclusively prove that my interpretation of the empirical results, which is in accordance with the predictions of the model, is correct, but I can offer two arguments to support it. The first is the context of the British cartels of the 1950s, described in section 3: the cartels were legal, long-standing, effective, usually comprising all or most of the important firms and facing limited outside competition, and they were typically operated by industrial trade associations which facilitated the coordination, monitoring and enforcement of collusion. In such circumstances, any differences across industries with respect to the degree of collusion may not be large.

The second argument is that even if the British cartels were able to achieve prices closer to monopoly prices in certain types of industries than in others, the high-degree-of-collusion industries would probably be those where collusion was more likely to occur and be sustainable. Several empirical studies indicate that collusion is more likely to occur in capital-intensive industries or those with high entry costs (see, for instance, Dick 1996). The evidence also suggests that collusion is hindered by product differentiation (Asch and Seneca 1975, Scherer and Ross 1990). Thus to the extent that the ability to raise prices considerably above competitive levels is positively correlated with the ability to form and sustain a cartel, low-capital, advertising-intensive and consumer good industries would have been subject to a lower degree of collusion than capital-intensive, low-advertising and producer good industries before the abolition of cartels. Consequently, they would have experienced a weaker effect on prices and welfare once collusion broke down. And yet what we see in the data is the exact opposite. Clearly, the evidence does not support the different-degrees-of-collusion story. Alternatively, if this story were true, then my econometric results would underestimate, if anything, the effects described in the theoretical model. I would then have to conclude that these effects are so strong that they dominate any others in the data.

In previous work I analysed the factors facilitating collusion across British manufacturing industries (defined at the four-digit level of aggregation) in the 1950s. I used the capital-labour ratio, *K/L*, as a proxy for the level of entry costs and two different measures of product differentiation: the advertising-sales ratio, *ADS*, and the industry type, distinguishing between producer good, consumer good and "intermediate" industries that manufacture significant quantities of both types of goods. I did not consider market size, since this is not generally thought to be directly related to cartel formation and stability on theoretical or empirical grounds. The mean of ln*K/L* was 1.23 for 71 collusive industries but only 0.64 for 80 competitive ones in 1958. Also, 62 of the 71 collusive industries but only 49 of the 80 competitive ones had *ADS* < 1% in the mid- to late 1950s. Finally, 49 (11) of the 71 collusive industries versus 29 (47) of the 80 competitive ones were manufacturers of producer (consumer) goods. Symeonidis (2003) provides more details and an econometric analysis that confirms these descriptive statistics. In summary, and consistent with earlier studies, I found that the incidence of collusion is facilitated by significant entry costs and product homogeneity. The present research suggests that the first and possibly the second of these characteristics are associated with a relatively lower welfare loss from collusion. Thus an interesting implication of the results of the present paper is that the welfare loss from collusive pricing may often be smaller in industries where cartels tend to form than in those where collusion is more difficult to achieve and sustain.

Table 1. Initial conditions in industries affected by the 1956 Act and in industries not affected

	Mean (standard deviation) of *PCM*, 1954	Mean (standard deviation) of *PCM*, 1958
Competitive industries (n = 52)	0.191 (0.071)	0.193 (0.075)
All collusive industries (n = 36)	0.169 (0.047)	0.172 (0.044)
Collusive industries with *LOWCAPINT* = 1 (n = 12)	0.184 (0.041)	0.196 (0.038)
Collusive industries with *LOWCAPINT* = 0 (n = 24)	0.162 (0.049)	0.159 (0.042)
Collusive industries with *ADV* = 1 (n = 7)	0.194 (0.049)	0.192 (0.036)
Collusive industries with *ADV* = 0 (n = 29)	0.164 (0.045)	0.167 (0.045)
Collusive industries with *PRODCON* = 1 (n = 11)	0.169 (0.047)	0.172 (0.044)
Collusive industries with *PRODCON* = 0 (n = 25)	0.191 (0.071)	0.192 (0.075)
Collusive industries with *MKTSIZE* = 1 (n = 23)	0.166 (0.051)	0.164 (0.040)
Collusive industries with *MKTSIZE* = 0 (n = 13)	0.175 (0.039)	0.184 (0.049)

Note: The figures are based on industries with available data for both 1954 and 1958. n indicates the number of industries.

Table 2. Regression results for ln*PCM* in industries with a change in competition regime. Fixed effects estimation

ln(*K/L*)	0.181	0.191	0.156	0.179	0.212	0.171	0.108	0.194
	(2.58)	(2.61)	(1.57)	(2.42)	(2.82)	(2.19)	(1.03)	(2.42)
ln*NFIRMS*	-	-	-	−0.052	-	-	-	-
				(−0.95)				
Y54	-	-	-	-	0.087	-	-	0.090
					(0.95)			(0.96)
Y63	0.078	0.162	0.217	0.153	0.198	0.167	0.228	0.204
	(2.47)	(2.43)	(2.96)	(2.28)	(2.27)	(2.43)	(3.02)	(2.31)
Y68	0.099	0.156	0.219	0.137	0.189	0.172	0.246	0.205
	(2.27)	(2.29)	(2.64)	(1.97)	(2.19)	(2.46)	(2.86)	(2.34)
Y73	0.070	0.245	0.317	0.223	0.272	0.265	0.356	0.292
	(1.04)	(2.36)	(3.29)	(2.10)	(2.37)	(2.41)	(3.52)	(2.42)
LOWCAPINT × *Y54*	-	-	-	-	−0.118	-	-	−0.103
					(−1.39)			(−1.24)
LOWCAPINT × *Y63*	-	−0.102	−0.172	−0.100	−0.161	−0.099	−0.158	−0.152
		(−1.55)	(−2.23)	(−1.52)	(−1.87)	(−1.49)	(−2.00)	(−1.74)
LOWCAPINT × *Y68*	-	−0.045	−0.113	−0.040	−0.108	−0.034	−0.085	−0.091
		(−0.92)	(−1.32)	(−0.81)	(−1.43)	(−0.72)	(−0.95)	(−1.22)
LOWCAPINT × *Y73*	-	−0.150	−0.218	−0.144	−0.213	−0.157	−0.207	−0.213
		(−2.24)	(−2.64)	(−2.16)	(−2.42)	(−2.37)	(−2.30)	(−2.44)
ADV × *Y54*	-	-	-	-	0.020	-	-	-
					(0.30)			
ADV × *Y63*	-	0.047	0.060	0.044	0.057	-	-	-
		(0.94)	(0.78)	(0.86)	(0.92)			
ADV × *Y68*	-	−0.007	0.007	−0.010	0.003	-	-	-
		(−0.15)	(0.08)	(−0.22)	(0.05)			
ADV × *Y73*	-	−0.149	−0.133	−0.152	−0.141	-	-	-
		(−2.19)	(−1.53)	(−2.20)	(−1.88)			
PRODCON × *Y54*	-	-	-	-	-	-	-	−0.070
								(−1.08)
PRODCON × *Y63*	-	-	-	-	-	0.017	−0.031	−0.016
						(0.32)	(−0.47)	(−0.28)
PRODCON × *Y68*	-	-	-	-	-	−0.053	−0.106	−0.084
						(−1.11)	(−1.46)	(−1.65)
PRODCON × *Y73*	-	-	-	-	-	−0.040	−0.095	−0.070
						(−0.58)	(−1.24)	(−0.97)
MKTSIZE × *Y54*	-	-	-	-	−0.051	-	-	−0.029
					(−0.56)			(−0.32)
MKTSIZE × *Y63*	-	−0.096	−0.125	−0.093	−0.120	−0.090	−0.103	−0.103
		(−1.42)	(−1.62)	(−1.35)	(−1.35)	(−1.32)	(−1.32)	(−1.17)
MKTSIZE × *Y68*	-	−0.072	−0.099	−0.064	−0.097	−0.062	−0.071	−0.077
		(−1.30)	(−1.19)	(−1.15)	(−1.22)	(−1.18)	(−0.84)	(−1.00)
MKTSIZE × *Y73*	-	−0.169	−0.198	−0.160	−0.194	−0.194	−0.204	−0.209
		(−2.17)	(−2.36)	(−2.01)	(−2.01)	(−2.50)	(−2.37)	(−2.18)
R^2	0.42	0.49	0.41	0.49	0.50	0.47	0.39	0.48
R^2_{LSDV}	0.84	0.86	-	0.86	0.86	0.85	-	0.86
Hausman statistic	13.53	22.64	-	23.08	23.47	15.02	-	17.10
Prob-value	0.001	0.001		0.003	0.01	0.02		0.01
Wooldridge test	2.63	4.43	-	4.39	3.95	4.68	-	3.92
Prob-value	0.114	0.042		0.043	0.055	0.037		0.056
AR(1)	-	-	0.13	-	-	-	0.14	-
No. of industries	36	36	36	36	36	36	36	36
No. of observations	178	178	142	178	178	178	142	178

Notes: Columns 1-2, 4-6 and 8: fixed effects estimation, t-statistics based on robust standard errors in parentheses. Columns 3 and 7: fixed effects estimation with AR(1) disturbances, t-statistics in parentheses.

Table 3. Robustness checks for industries with a change in competition regime. Fixed effects estimation

	Dep. variable: *PCM*		Dependent. variable: ln*PCM*				
ln(*K/L*)	0.039	0.032	0.181	0.139	0.088	0.155	0.142
	(2.96)	(2.55)	(2.31)	(1.71)	(0.76)	(2.09)	(1.42)
Y54	0.011	-	0.103	-	-	-	-
	(0.78)		(1.05)				
Y63	0.032	0.028	0.232	0.190	0.255	0.188	0.239
	(2.34)	(2.51)	(2.55)	(2.72)	(3.17)	(2.53)	(3.16)
Y68	0.029	0.028	0.216	0.188	0.262	0.217	0.272
	(2.25)	(2.61)	(2.39)	(2.61)	(2.89)	(2.80)	(2.99)
Y73	0.049	0.049	0.305	0.293	0.382	0.408	0.481
	(2.42)	(2.46)	(2.54)	(2.58)	(3.60)	(3.10)	(3.54)
LOWCAPINT × *Y54*	−0.017	-	−0.135	-	-	-	-
	(−1.21)		(−1.43)				
LOWCAPINT × *Y63*	−0.021	−0.013	−0.221	−0.153	−0.217	−0.097	−0.165
	(−1.46)	(−1.02)	(−2.42)	(−2.12)	(−2.40)	(−1.39)	(−2.15)
LOWCAPINT × *Y68*	−0.010	0.002	−0.138	−0.048	−0.105	−0.032	−0.095
	(−0.79)	(0.24)	(−1.64)	(−0.90)	(−1.02)	(−0.64)	(−1.11)
LOWCAPINT × *Y73*	−0.034	−0.026	−0.227	−0.167	−0.221	−0.146	−0.213
	(−2.17)	(−2.01)	(−2.30)	(−2.16)	(−2.08)	(−2.18)	(−2.51)
ADV × *Y54*	0.004	-	0.031	-	-	-	-
	(0.31)		(0.45)				
ADV × *Y63*	0.019	-	0.076	-	-	0.048	0.062
	(1.74)		(1.16)			(0.92)	0.80)
ADV × *Y68*	0.011	-	0.016	-	-	0.003	0.019
	(1.07)		(0.27)			(0.07)	(0.23)
ADV × *Y73*	−0.016	-	−0.131	-	-	−0.131	−0.115
	(−1.01)		(−1.70)			(−1.94)	(−1.30)
PRODCON × *Y63*	-	0.007	-	0.032	−0.012	-	-
		(0.82)		(0.57)	(−0.18)		
PRODCON × *Y68*	-	−0.008	-	−0.052	−0.101	-	-
		(−1.01)		(−1.00)	(−1.29)		
PRODCON × *Y73*	-	−0.002	-	−0.042	−0.095	-	-
		(−0.22)		(−0.57)	(−1.16)		
MKTSIZE × *Y54*	−0.002	-	−0.084	-	-	-	-
	(−0.15)		(−0.84)				
MKTSIZE × *Y63*	−0.021	−0.018	−0.154	−0.105	−0.133	−0.091	−0.117
	(−1.48)	(−1.56)	(−1.63)	(−1.44)	(−1.53)	(−1.27)	(−1.51)
MKTSIZE × *Y68*	−0.016	−0.011	−0.117	−0.058	−0.081	−0.052	−0.074
	(−1.29)	(−1.33)	(−1.34)	(−0.97)	(−0.85)	(−0.93)	(−0.88)
MKTSIZE × *Y73*	−0.041	−0.043	−0.209	−0.194	−0.216	−0.157	−0.179
	(−2.45)	(−3.04)	(−1.95)	(−2.19)	(−2.22)	(−1.97)	(−2.13)
ln*SALES*	-	-	-	-	-	−0.103	−0.080
						(−1.46)	(−0.87)
ln*SIZE*	-	-	-	-	-	0.069	0.096
						(0.72)	(0.73)
ln*UNION*	-	-	-	-	-	−0.351	−0.427
						(−1.79)	(−1.72)
R^2	0.53	0.51	0.48	0.45	0.39	0.50	0.44
R^2_{LSDV}	0.87	0.86	0.86	0.85	-	0.86	-
Hausman statistic	23.05	14.66	19.63	12.18	-	22.76	-
Prob-value	0.001	0.02	0.002	0.03		0.007	
Wooldridge test	2.83	3.54	3.53	4.21	-	4.75	-
Prob-value	0.101	0.068	0.069	0.048		0.036	
AR(1)	-	-	-	-	0.14	-	0.12
No. of industries	36	36	33	33	33	36	36
No. of observations	178	178	164	164	131	178	142

Notes: Columns 1-2, 3-4 and 6: fixed effects estimation, t-statistics based on robust standard errors in parentheses. Columns 5 and 7: fixed effects estimation with AR(1) disturbances, t-statistics in parentheses.

Table 4. Regression results for ln*PCM* in competitive industries. Random effects estimation

ln(*K/L*)	0.073	0.191	0.068	0.092
	(2.00)	(2.58)	(1.99)	(2.57)
ln*NFIRMS*	-	-	–0.058	-
			(–1.21)	
Y63	0.141	0.099	0.087	0.095
	(6.43)	(2.14)	(1.92)	(2.14)
Y68	0.127	0.009	0.001	0.002
	(3.98)	(0.14)	(0.01)	(0.04)
Y73	0.102	–0.046	–0.052	–0.093
	(2.28)	(–0.51)	(–0.59)	(–1.05)
LOWCAPINT×*Y63*	-	0.065	0.071	0.072
		(1.43)	(1.58)	(1.40)
LOWCAPINT×*Y68*	-	0.153	0.157	0.165
		(2.60)	(2.71)	(2.53)
LOWCAPINT×*Y73*	-	0.194	0.204	0.186
		(2.77)	(2.93)	(2.57)
ADV×*Y63*	-	–0.018	–0.023	-
		(–0.48)	(–0.62)	
ADV×*Y68*	-	–0.034	–0.042	-
		(–0.66)	(–0.86)	
ADV×*Y73*	-	–0.021	–0.025	-
		(–0.37)	(–0.42)	
PRODCON×*Y63*	-	-	-	–0.019
				(–0.48)
PRODCON×*Y68*	-	-	-	–0.030
				(–0.56)
PRODCON×*Y73*	-	-	-	0.069
				(1.12)
MKTSIZE×*Y63*	-	0.013	0.024	0.018
		(0.35)	(0.63)	(0.47)
MKTSIZE×*Y68*		0.073	0.086	0.081
		(1.46)	(1.78)	(1.59)
MKTSIZE×*Y73*	-	0.063	0.078	0.081
		(1.02)	(1.28)	(1.35)
Constant	–1.77	–1.78	–1.49	–1.78
	(–36.65)	(–36.80)	(–5.92)	(–36.43)
R^2	0.29	0.37	0.37	0.38
Hausman statistic	0.11	12.85	11.61	7.39
Prob-value	0.95	0.03	0.07	0.19
No. of industries	55	55	55	55
No. of observations	272	272	272	272

Note: t-statistics based on robust standard errors in parentheses.

APPENDIX 1: List of collusive industries with a change of competition regime

Industry	LOWCAPINT	ADV	PRODCON	MKTSIZE
grain milling	0	0	0	1
bread and flour confectionery	0	0	1	1
biscuits	0	1	1	1
milk and milk products	0	1	1	1
cocoa, chocolate and sugar confectionery	0	1	1	1
paint	0	1	1	1
surgical bandages and sanitary towels	1	0	0	0
steel tubes	0	0	0	1
iron castings	0	0	0	1
aluminium and aluminium alloys	0	0	0	1
copper, brass and other copper alloys	0	0	0	1
industrial engines	0	0	0	0
industrial plant and steelwork	0	0	0	1
surgical instruments and appliances	1	0	0	0
electrical machinery	1	0	0	1
insulated wires and cables	0	0	0	1
electrical appliances primarily for domestic use	0	1	1	1
batteries and accumulators	1	1	0	0
electric lamps, electric light fittings and wiring accessories	1	1	1	1
hand tools and implements	1	0	1	0

wire and wire manufactures	0	0	0	1
metal hollow-ware	1	0	0	0
miscellaneous metal goods	1	0	0	1
spinning and doubling on the cotton and flax systems	0	0	0	1
jute	0	0	0	0
rope, twine and net	1	0	0	0
carpets	0	0	1	1
canvas goods and sacks	1	0	1	0
textile finishing	0	0	0	1
asbestos manufactures	0	0	0	0
refractory goods	0	0	0	0
building bricks and other non-refractory goods	1	0	0	1
pottery	1	0	1	0
miscellaneous building materials and mineral products	0	0	0	1
paper and board	0	0	0	1
linoleum, plastics floor covering, leathercloth	0	0	0	0

APPENDIX 2: Data sources and construction of variables

The data sources for competition were described in the text. Data on net output, sales revenue, wages and salaries, and employment were obtained from the industry reports of the Census of Production (various years). The figures are for all firms with at least 25 employees. Small corrections were sometimes made to ensure comparability over time.

Estimates of capital stock, defined as plant and machinery, at the three-digit level of aggregation are available from O'Mahony and Oulton (1990). These are net stock estimates constructed on the assumption of fixed and "short" asset lives and exponential depreciation rates. For a few cases where employment data were available at a more disaggregated level, I adjusted the O'Mahony and Oulton capital stock estimates on the basis of Census of Production data on the fraction of investment on plant and machinery accounted for by each "principal product" within any given three-digit industry.

Data on manufacturers' advertising expenditure in the UK for 1958 were taken from the *Statistical Review of Press and TV Advertising*. The figures were adjusted to correct for the underreporting of press advertising and the failure to take into account discounts for TV advertising and costs of production of advertisements. The *Statistical Review* contains information mostly for consumer good industries; however, nearly all the industries for which data are not reported could be easily classified as low-advertising industries.

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