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Price-Concentration Analysis in Merger Cases with Differentiated Products

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Abstract This paper considers the empirical assessment of the relationship between prices and number of firms in local markets in geographic or, more generally, characteristic space and its use as evidence in merger cases. It outlines a structural, semi-nonparametric econometric model of competition in such markets, examines its testable implications in terms of price-concentration relationships, and demonstrates that the model is non-parametrically identified. This general approach to price-concentration analysis in differentiated product markets is illustrated in a small-scale application to cinemas in the UK. The application highlights the main decision points faced by an authority when assessing the weight that can be attached to this type of analysis as evidence.

JEL L11, C31

Keywords Differentiated products; local competition; non-parametric identification

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1 Introduction

The analysis of the relationship between prices for goods and services prevailing in markets and the structure of these markets is at the heart of virtually every merger assessment. In an important class of cases, markets can be viewed as local in geographic or, more generally, characteristic space, and products and services offered on them as differentiated accordingly. Antitrust authorities are interested in the assessment of the relationship between prices and concentration on such differentiated product markets as an investigative technique, as it can address the key issue in merger cases: Will an increase in concentration bring about higher prices? This paper makes three contributions to this kind of competition analysis. First, it provides a structural econometric model of competition in such markets in which the price-concentration relationship of interest can be empirically assessed. Second, it illustrates this methodology in an application to cinemas in the UK. And, third, it highlights the critical decision points for an antitrust authority when evaluating the results of an application of this methodology as evidence.

Econometric price-concentration studies evolved from classic structure-conduct-performance analyses (Schmalensee (1989)) and have gained wide prominence in the context of horizontal merger analyses, both in the US[1] and in Europe.[2] Price concentration studies also often feature in the empirical industrial organization literature.[3]

Price concentration studies are typically conducted at the industry level, using

[1]See, e.g., Baker (1999), Pautler (2001).

[2]In the UK, the Monopolies and Merger Commission and its successor, the Competition Commission, used price concentration analysis to investigate competition between funeral parlors (1995) and supermarkets (2000). The review of competition in the banking sector conducted by Cruickshank also conducted a price concentration study.

[3]See e.g. Bresnahan and Reiss (1991), Pinske et al. (2002); Manuszak and Moul (2006a), re-examining the classical *FTC vs. Staples* decision; Manuszak and Moul (2006b) for retail gasoline; Borenstein (1989, 1990), Morisson and Winston (1990), Brueckner et al. (1992), Evans and Kessides (1993), Kim and Singal (1996), Singal (1996) for airfares on specific routes; Neumark and Sharpe (1992), Hannan (1992), Cyrnak and Hannan (1999) for banking; Davis (2005, 2006) for US movie theaters; Mazzeo (2002b) for motels along US interstate highways.

cross-section or panel data.[4] The general objective of a price concentration study is to investigate how concentration is related to market power, i.e. the ability of firms to price above marginal cost. To this end, there needs to be a number of independent markets for the same products where concentration varies sufficiently while other parameters remain relatively constant or can be accounted for with reasonable accuracy (particularly costs). In general, measures of concentration used will be Herfindahl-Hirschman Indices (HHIs) and C3 or C4 concentration ratios.[5] Each may have their own advantages (see Bishop and Walker (2002) for a discussion). In a common class of cases, where the candidate market is defined so that firms are similar in size[6], the number of firms (or independent fascias) may be used instead of concentration measures based on market shares. Using the number of firms instead of other, e.g. market-share based, concentration measures has the advantage of removing one source of endogeneity in the relationship as the number of firms is not a function of endogenous firm output or revenues (see Evans et al. (1993)).

In price-concentration studies it is not necessarily a problem that the firms involved sell more than one good or that these products are differentiated either in the geographical or characteristics space, as long as the product mix or the actual product specification does not vary significantly across the local markets of the cross-section. This often applies to retail mergers. A merger between retail chains may bring under the same ownership outlets that were previously competing in a given local area. In these cases the main differentiating factor is likely to be the geographical location of the outlets, next to other dimension of differentiations of the product offering.

Increasing public availability of pricing data and a desire on the part of regulators and courts for quantifiable evidence are certain to enhance both interest in and scrutiny of econometric price-concentration analyses. This will typically require

[4]In principle, they can also be conducted at the firm level, using time series data.

[5]The HHI is the sum of the squared market shares of the firms in a market, and the C3 and C4 concentration ratios are the market shares of the three or four largest firms in the market.

[6]This can often be the case in EU and UK cases involving supermarkets, where the relevant market is typically defined by reference to the stores' size in square feet. Movie theaters were also segmented in the past by size (number of screens and seats) by the UK's Office of Fair Trading and Competition Commission.

balancing rigorous sophistication with practicability and robustness. Practicability often takes precedence and induces reduced form approaches, in the absence of a structural competition model (e.g. in *FTC vs. Staples*; cp. Baker (1999)). There are at least two fundamental risks associated with this research strategy: The structural interpretation of estimation results from the perspective of competition economics is unclear, and there is no theoretical guidance regarding potential joint endogeneity of firms' decision parameters and, if suspected, the choice and validity of instruments. The methodology advocated in this paper outlines a structural econometric model for competition in differentiated product markets that overcomes these shortcomings of reduced form models. The econometric model builds on classical theoretical models for differentiated product markets (Hotelling (1929), Salop (1979)) and the econometric framework for differentiated product industries due to Berry (1994) and Berry, Levinsohn and Pakes (1995) and the literature following their approach[7]. It models a two-stage game of firm entry and price setting. The model is nested in the framework of Bresnahan and Reiss (2006) and provides an alternative to Mazzeo's (2002a, 2002b) approach to correcting for endogenous market structure.[8] The model is semi-nonparametric, making only minimal functional form assumptions. Section 2 of this paper examines the testable implications of the model and demonstrates that the relationships of interest are nonparametrically identifiable on the basis of instruments that have a sound theoretical justification.

This approach, in general, stipulates a general nonparametric, nonlinear instrumental variable regression estimation methodology.[9] Nonparametric robustness is clearly a desirable feature of any econometric model. Nonparametric estimation, however, requires sample sizes significantly larger than those that may be available in applications. In small-scale applications, therefore, parametric and possibly linear instrumental variable methods may be more advantageous and practicable. This is illustrated in an application to local multiplex cinema markets in the UK. A small number of relatively similar multiplex cinema operators (fascias) are active in these

[7]The model presented in this paper follows Berry and Waldfogel (1999) which, itself, is cast within the framework of Berry (1994).

[8]A comparison of the two models is provided at the end of section 2.

[9]Currently available methods include control functions (Blundell and Powell (2003)) or series approximations (Newey and Powell (2003)).

markets, and adult ticket prices and number of operators and cinemas in them are readily available from public sources, as are various socio-demographic characteristics of localities. The application demonstrates the biases that may result from an alternative, purely reduced form approach which is presented for the purpose of comparison. The application also shows that the modelling strategy is critical for empirical assessment as it puts regulators in a better position when assessing how much weight to place on such studies as evidence in competition inquiries.

This paper argues that, contrary to some criticism (e.g. Baker and Bresnahan (1992)), price-concentration studies can be generally useful to assess the impact of mergers in differentiated product industries. It also argues that the structural form approach is superior to reduced form analyses as it enables the authority to make a more informed judgement.

The paper proceeds as follows. Section 2 outlines the general structural, semi-nonparametric econometric model, including its testable implications, and examines its non-parametric identification. Section 3 considers a particular parametric specification that is applied to local multiplex cinema markets in the UK. After a brief description of the data, it also presents estimation results and compares them to estimates obtained from an alternative, reduced form model. Section 4 discusses implications for applied competition analysis. Section 5 concludes.

2 Econometric Model

This section outlines a structural, semi-nonparametric econometric model of differentiated product markets. The model follows Berry and Waldfogel (1999). It maps the classical framework for the analysis of differentiated products, based on conditional indirect utilities (Berry (1994), Berry, Levinsohn and Pakes (1995)), onto two stage game in which firms, on the first stage, decide whether or not to enter a market and, on the second stage, compete on price, conditional on having entered.

Let consumers in a market be indexed by i, and suppose there are n firms serving the market on the second stage of the firms' strategic game. It is assumed that the

conditional indirect utility of consumer i derived from product j, produced by firm j, $j \in J = \{1, \cdots, n\}$, is

$$U_{ij} = \delta(p_j, \mathbf{x}_j) + v_i(\sigma) + (1 - \sigma)\epsilon_{ij},$$

where p_j is product j's price, $\mathbf{x}_j$ is a vector of other product characteristics, and ϵ_{ij} is the idiosyncratic utility that consumer i derives from product j and that is not accounted for by $(p_j, \mathbf{x}_j')'$. The nesting parameter σ governs the degree to which the consumer derives utility from any of the products $j \in J$ relative to an outside option, i.e. not to consume any one of them.[10] Assuming that the idiosyncratic utility components ϵ_{ij} are independent across i and j and identically distributed extreme value type I, the functions $\delta(p_j, \mathbf{x}_j)$ represent the average utility derived from product j, and product j's market share is given by

$$s_j(\mathbf{p}, \mathbf{x}; n) = \frac{\exp(\delta(p_j, \mathbf{x}_j)/(1 - \sigma))}{D} \frac{D^{1-\sigma}}{1 + D^{1-\sigma}},$$

where $D = \sum_{k \in J} \exp(\delta(p_k, \mathbf{x}_k)/(1 - \sigma))$, $\mathbf{p} = (p_1, \cdots, p_J)'$ and $\mathbf{x} = (\mathbf{x}_1', \cdots, \mathbf{x}_J')'$. Notice that the first ratio corresponds to the probability of choosing product j conditional on choice set J, and the second ratio to the probability of choosing any of the J products as opposed to the outside option.[11] When $\sigma = 1$ then the share of the inside goods is one half, regardless of the number of differentiated inside varieties n.

Suppose that consumers are uniformly distributed in the market, with mass $\mu(\mathbf{z})$, where $\mathbf{z}$ denotes a vector of factors determining market size.[12] These factors might include socio-demographic characteristics such as population and income related measures. Suppose, furthermore, that firms face marginal cost c and fixed cost f

[10] Note that the consumer does not derive any idiosyncratic, product specific utility from the goods $j \in J$ if $\sigma = 1$. The distribution of $v_i(\sigma)$ collapses to zero when σ approaches zero. Cardell (1997) provides a detailed discussion of the distribution of $v(\sigma)$. The parameter σ approximates the correlation between the two nests, i.e. of inside goods and the outside option, respectively. See also McFadden (1978) as well as the discussion in Maddala (1983).

[11] In the terminology of nested logit models, $\ln(D)$ is referred to as the inclusive value of the inside goods.

[12] Consumer heterogeneity can be incorporated in a variety of ways. For example, the nesting parameter σ can be allowed to vary across consumers. Letting $P(\sigma)$ denote the distribution of σ, product j's market share is then $s_j(\mathbf{p}, \mathbf{x}; n) = \int_\sigma \frac{\exp(\delta(p_j, \mathbf{x}_j)/(1-\sigma))}{D} \frac{D^{1-\sigma}}{1+D^{1-\sigma}} dP(\sigma)$.

and compete on price. Hence, given prices p_k for firms $k \neq j$, firm j solves

$$\max_{p_j}(p_j - c)s_j(p_j, \mathbf{p}_{-j}, \mathbf{x}; n)\mu(\mathbf{z}) - f,$$

where $\mathbf{p}_{-j} = (p_1, \cdots, p_{j-1}, p_{j+1}, \cdots, p_n)$. A necessary condition for a symmetric Bertrand-Nash equilibrium $p^\star$ is

$$\begin{aligned}\frac{p^\star - c}{p^\star} &= -\left[\frac{p^\star \frac{\partial}{\partial p_j} s_j(p_j, \mathbf{p}^\star_{-j}, \mathbf{x}; n)|_{p_j = p^\star}}{s_j(\mathbf{p}^\star, \mathbf{x}; n)}\right]^{-1}, j = 1, \cdots, n \\ &= \frac{\beta_j^\star}{p^\star}\left[\frac{1}{1-\sigma}\left(1 - \frac{\exp\left(\frac{\delta(p^\star, \mathbf{x}_j)}{1-\sigma}\right)}{D^\star}\right) + \left(1 - \frac{D^{\star(1-\sigma)}}{1 + D^{\star(1-\sigma)}}\right)\frac{\exp\left(\frac{\delta(p^\star, \mathbf{x}_j)}{1-\sigma}\right)}{D^\star}\right]^{-1},\end{aligned}$$

where $\beta_j^\star = -\frac{\partial}{\partial p_j}\delta(p_j, \mathbf{p}^\star_{-j}, \mathbf{x}_j)|_{p_j = p^\star}$ and $D^\star = \sum_{k \in J} \exp(\delta(p^\star, \mathbf{x}_k)/(1-\sigma))$. If symmetry is further strengthened to also involve characteristics other than price, so that $\mathbf{x}_j = \mathbf{x}^\star$ for all $j \in J$, the last expression reduces to

$$\frac{p^\star - c}{p^\star} = \frac{1}{p^\star}\left[\frac{1}{1-\sigma}\frac{n-1}{n} + \frac{1}{n}\frac{\exp(\delta(p^\star, \mathbf{x}^\star))}{\exp(\delta(p^\star, \mathbf{x}^\star)) + n^{1-\sigma}}\right]^{-1}.$$

Consider the extreme cases of $\sigma = 1$ and $\sigma = 0$. In the former, firms have to price equal to marginal cost, because positive profits may trigger entry, but due to constant total market share of the inside goods, the entrant will cannibalize the market share of the incumbents and thereby reduce their profits. In the latter, the equilibrium price cost margin

$$p^\star - c = \left[1 - (\exp(\delta(p^\star, \mathbf{x}^\star)) + n)^{-1}\right]^{-1}$$

decreases with the number of firms n; in this case the expansion of the market share due to an additional product among the inside goods is maximal.

These models induce general, semi-nonparametric structural pricing equations of the form

$$p^\star = \pi(c, n, \mathbf{x}),$$

which depend on the jointly endogenous variable n.[13] Partitioning $\mathbf{x} = (\tilde{\mathbf{x}}', \xi)'$ into observed and unobserved components $\tilde{\mathbf{x}}$ and ξ, respectively, the econometric version

[13] While the function $\delta(\cdot)$ is left non-parametric, the model makes parametric assumptions about ϵ_{ij} and σ.

of the structural pricing equation is

$$p^\star = \pi(c, n, \tilde{\mathbf{x}}, \xi).$$

Firms' decision processes involve two stages. In the first stage, firms' pricing decisions take the number of firms in the market, n, as given. When deciding whether or not to be active in this market in the first stage, firms will take into account the size of the market, $\mu(\mathbf{z})$, the costs of operating in the market, (c, f), and the (prospective) price that can be earned in equilibrium. Assuming all firms make this decision simultaneously, the equilibrium number of firms is such that economic profits in this market are zero, conditional on the prospective price. Ignoring integer constraints, for any given price $p > c$, the equilibrium number of firms $n^\star$ satisfies

$$(p - c)\mu(\mathbf{z}) - n^\star f = 0,$$

which implies

$$n^\star = (p - c)\mu(\mathbf{z})/f = \eta(p, c, f, \mu(\mathbf{z})),$$

for some function $\eta(\cdot)$ which is increasing in p and $\mu(\mathbf{z})$ and decreasing in c and f.

To account for the possibility of only partial information about $\mathbf{z}$, partition this covariate vector $\mathbf{z} = (\tilde{\mathbf{z}}', \zeta)'$, into observed and unobserved factors determining market size, $\tilde{\mathbf{z}}$ and ζ, respectively. Then, $n^\star = \eta(p, c, f, \tilde{\mathbf{z}}, \zeta)$ is the econometric structural entry equation.

The system of equations

$$\begin{aligned} p^\star &= \pi(c, n, \tilde{\mathbf{x}}, \xi) \\ n^\star &= \eta(p, c, f, \tilde{\mathbf{z}}, \zeta). \end{aligned}$$

forms the structural econometric model. Since these relationships are only observed at equilibrium values for $p^\star$ and $n^\star$, identifying the structural functions $\pi(\cdot)$ and $\eta(\cdot)$ is impeded by a classic econometric simultaneity problem, because the number of firms in endogenous in the pricing equation, and price is endogenous in the number of firms equation.

In order to uncover the structural relationship, one needs instruments for the equation of interest. Instruments for the respective endogenous right-hand side

variable can be any exogenous variables that are not included in the equation of interest. Hence, firms' fixed costs f can act as instrument for the number of firms in the pricing equation. Moreover, if the exclusion restriction $\tilde{\mathbf{z}} = (\tilde{\mathbf{z}}_1', \tilde{\mathbf{z}}_2')'$, $\tilde{\mathbf{x}} = (\tilde{\mathbf{x}}_1', \tilde{\mathbf{x}}_2')'$ and $\tilde{\mathbf{z}}_1 \notin \tilde{\mathbf{x}}$ and $\tilde{\mathbf{x}}_1 \notin \tilde{\mathbf{z}}$ hold, then $\tilde{\mathbf{x}}_1$ can act as instruments for p in $\eta(\cdot)$, and $\tilde{\mathbf{z}}_1$ as instruments for n in $\pi(\cdot)$, and both functions are identified. Hence, assuming $(\xi, \zeta) \perp (\tilde{\mathbf{x}}_1, \tilde{\mathbf{z}}_1)$, given a sufficient amount of data, both functions can, in principle, be estimated non-parametrically.[14] In cases in which data are not abundant, parametric versions of these functions can be estimated by instrumental variable methods.

The functions $\pi(\cdot)$ and $\eta(\cdot)$ form the structural part of the model. They can be solved for the reduced form of the model, which yields equilibrium price and number of firms as functions of the exogenous variables of the model, i.e. the market size, taste and cost parameters. Antitrust authorities's interest typically focuses on the structural relation $\pi(\cdot)$.

The econometric model outlined in this section is closely related to Mazzeo's (2002b) approach to correct for endogenous market structure. In the relationship between price and number of firms, Mazzeo includes a correction for self-selection of firms into the market. The selection equation that induces this term via a correlation of its residuals with the residuals of the equation of interest is governed by exogenous covariates; in Mazzeo's model, it relates to whether a given number of firms can achieve non-negative profits in the market. The correction term then is an estimate of the conditional expectation of the residuals in the equation of interest, given the range of residuals in the exogenous selection equation that are consistent with the observed number of firms being just profitable. The instrumental variable methodology proposed in this paper can also be interpreted in terms of the inclusion of a correction term. It is well known that, in the linear regression model, the 2SLS coefficient estimates can be obtained by including the residuals of the first stage regression into the second stage estimation. Here, the correction term has the interpretation of the conditional expectation of the residuals of the equation of interest, given the residuals from the reduced form regression equation for the endogenous

[14] E.g. using the control function approach of Blundell and Powell (2003) or the series approximation approach due to Newey and Powell (2003).

regressor.

3 An illustration: Cinemas in the UK

3.1 The industry and the data available for analysis

The UK cinema industry is described in detail in a recent report by the UK Competition Commission.[15] This section summarizes the main features of the industry and the data used for the analysis.

The Competition Commission report provides the following descriptive statistics about the industry. There exist several large movie exhibitors (or fascias) in the UK, with a total revenue of over GBP 900 million in 2004. These cinema operators are often owned by private equity houses. The current industry structure results from a sequence of mergers and acquisitions, primarily in the 1990s. Over 70 percent of UK screens are now owned by four exhibitors: Terra Firma (UCI, Odeon; 924 screens in January 2005), Blackstone (Cineworld and UGC; 787 screens), Vue (incl. Warner Village; 409 screens), and National Amusements (Showcase; 237 screens). 73 percent of all UK screens are now in multiplex cinemas, with just below 2500 screens in total.

For the analysis presented in this paper, only multiplex cinemas in England with at least 5 screens were considered. The London metropolitan area was excluded, as it is considered to have very different market features compared to those prevailing in the rest of England. This leaves 153 such multiplexes in England. With regard to these cinemas, the following data were available: Saturday night adult ticket prices[16], the number of competing multiplexes and fascias in 10-minute and 20-minute drive-time around the cinema, the minimum drive time to the closest multiplex, the

[15] This report is available on the Competition Commission's website, at http://www.competition-commission.org.uk.

[16] There is considerable variation in prices, across facias and across locations within a fascia; the mean (median) price is GBP6.04 (GBP5.95), with a minimum (maximum) price of GBP3.50 (GBP7.50) and a standard deviation of GBP0.58.

population and the deprivation index[17] of the area that the cinema is located in.

3.2 Model Specification

In light of the econometric model, prices - which in this market are typically held in place for one year - and number of multiplexes or fascias in local markets, defined by drive-time, are jointly endogenous. Hence, the latter must be instrumented when estimating the pricing equation $\pi(\cdot)$. The population size of a local area is likely to affect the size of the market, but is unrelated to consumers' conditional indirect utilities. In other words, local population can take the role of $\tilde{\mathbf{z}}_1$ in the discussion of the previous subsection. Moreover, in line with the discussion in the preceding section, if consumers' willingness to trade-off product characteristics does not depend on income, then the deprivation index is part of $\tilde{\mathbf{z}}_1$ as well and can serve as additional instrument. The illustration presented in this section maintains this hypothesis. An alternative view, explored below as well, might stipulate that consumers' heterogeneity in choosiness is a function of income, in which case the deprivation index is part of $\mathbf{x}$ and, hence, cannot act as an instrument for n in the estimation of $\pi(\cdot)$. Finally, planning applications could serve as an instrument for n as well, as they indicate firms' desire to enter a market.[18]

Due to the relatively small number of cross-sectional units, i.e. multiplexes, a parametric specification is estimated.[19] For simplicity, this specification is a straightforward linear model,

$$\ln(p_k) = \alpha + \beta n_k + \xi_k,$$

where k indexes multiplexes and n_k is the number of competing multiplexes or fascias

[17]The Overall Index of Multiple Deprivation 2004; the index has been constructed by the Social Disadvantage Research Centre at the University of Oxford and is published by the Office of the Deputy Prime Minister. The index combines several indices of deprivation, including income, employment, health deprivation and disability, education, skills and training, barriers to housing and services, crime and living environment.

[18]We are indebted to Peter Davis who suggested this instrument. Unfortunately, no complete set of observations on local planning applications was available for this application.

[19]It is worth mentioning, however, that Newey and Powell (2003) illustrate their methodology in a small scale simulation with similar sample size.

Pairwise Correlations

	mc1	mc2	fascia
di	0.412	0.411	0.452
pop	0.461	0.448	0.560

Table 1: Pairwise correlations: instruments deprivation index (di) and population (pop), vs. number multiplexes in 10-minute (mc1) and 20-minute (mc2) area, and number of fascias in 10 minutes.

in a 10-minute or 20-minute drive-time area, respectively. The model is estimated by 2-Stage-Least-Squares, using the local population and deprivation index as instruments. Table 1 presents the estimated correlation between the instruments and the respective right-hand-side endogenous variables.

3.3 Estimation Results

Table 2 summarizes the estimation results. Columns 1 and 2 provide 2SLS estimation results for the model with the number of multiplex cinemas in 10-minute and 20-minute drive-time areas as measure for market concentration. Columns 4 and 5, for comparison, show the corresponding OLS estimates. The comparison reveals the downward bias (in absolute value) that arises when the endogeneity of number of competitors (n) is not taken into account.

Columns 3 and 6 present 2SLS and OLS estimates for the model using the number of fascias in 10-minute drive time areas as measure for local concentration. In general, there may be good reasons for choosing either the number of firm-level establishments or the number of fascias as the appropriate covariate. The former estimates the effect of a merger as a result of eliminating one competing outlet or product type, while the latter does so by reducing the number of independently owned groups of outlets or product types by one. Again, OLS is seen to suffer from a severe downward bias in absolute value.

For comparison, the relationship between prices and number of fascias is also estimated non-parametrically. For this, the non-parametric function of interest,

Log Price Regressions (1)

	2SLS			OLS		
C	1.802	1.824	1.977	1.758	1.762	1.886
	(0.017)	(0.021)	(0.029)	(0.010)	(0.011)	(0.017)
MC1	-0.111			-0.032		
	(0.025)			(0.011)		
MC2		-0.033			-0.009	
		(0.007)			(0.003)	
FASCIA			-0.109			-0.057
			(0.017)			(0.0091)

Table 2: OLS and 2SLS point estimates; standard errors in parenthesis

$p^\star = \pi(n, \xi)$, is left unrestricted. No parametric linearity restrictions or additivity assumptions with regard to the econometric errors ξ are imposed. This makes this model a candidate for an application of the control function approach due to Blundell and Powell (2003). Their approach uses the independence assumptions $(\xi, \zeta) \perp (\tilde{\mathbf{x}}_1, \tilde{\mathbf{z}}_1)$ to deduce conditional independence restrictions $\xi|(\mathbf{x}, \mathbf{z}) \sim \xi|(\mathbf{x}, \zeta) \sim \xi|\zeta$ to construct a non-parametric analogue to the parametric 2SLS estimator, where in this application $\tilde{\mathbf{z}}_1$ corresponds to the deprivation index and population variables. Essentially, this estimator corrects for the endogeneity of the number of firms, n, by including the residuals $\hat{\zeta}^{NP} = n - \hat{m}^{NP}(\tilde{\mathbf{z}}_1)$ from the first-stage non-parametric regression $\hat{m}$ of n onto $\tilde{z}_1$ into the main non-parametric regression of p on n and subsequently integrating with respect to a non-parametric estimate $\hat{F}_{\hat{\zeta}^{NP}}$ of the distribution of $\hat{\zeta}^N$, i.e. $\hat{m}^{CF}(n) = \int \hat{\pi}^{NP}(n, \hat{\zeta}^{NP}) d\hat{F}_{\hat{\zeta}^{NP}}$. Table 3 summarizes the non-parametric point estimates, presenting conventional non-parametric conditional mean estimates (NP) next to non-parametric estimates using the control function approach (CF), as well as the estimated percentage impact on price of an additional fascia.[20] The non-parametric estimates confirm the bias introduced from ignoring the endogeneity of n. In comparison with the parametric estimates, the non-parametric estimates also show that the impact of an additional firm on local

[20]There is only one area where $n = 5$; this observation is excluded from the analysis.

Non-parametric Log Price Regressions

fascias	$\hat{m}^{NP}(n)$	$\Delta\hat{m}^{NP}(n)$	$\hat{m}^{CF}(n)$	$\Delta\hat{m}^{CF}(n)$
$n = 1$	1.841393		1.843627	
$n = 2$	1.775552	0.065841	1.784242	0.059385
$n = 3$	1.745016	0.030536	1.730825	0.053417
$n = 4$	1.678534	0.066482	1.643298	0.087527

Table 3: Non-parametric point estimates, with and without correcting for endogeneity

price is likely to be nonlinear.[21]

Assuming uniformly distributed consumers and no location restrictions on the part of the firms, the number of competitors n and the average inter-firm distance $1/n$ are inversely related.[22] In this case, a dual analysis can be carried out using the minimum drive time to the closest multiplex (mtm) as an alternative measure that is indicative of competitors' proximity. This is essentially a consistency check on the estimated model. Table 4 provides the corresponding 2SLS and OLS estimates. The estimates from the dual model yield qualitative conclusions which are consistent with the ones obtained in the original model specification. In the original model, a higher number of firms, on average, is estimated to induce lower prices. Conversely, a higher number of firms would be expected to reduce the minimum drive time to the closest competitor, so that lower minimum drive times to competitors would be associated with lower prices. The dual analysis confirms this.

For the purpose of competition analysis, the 2SLS estimation results suggest that a 10-minute drive time area around a multiplex is the relevant antitrust market, because reducing the number of competing multiplexes or fascias by one is estimated to increase the adult ticket price on the order of 10 percent. From a methodological

[21]This is difficult to assess, however, because there are fewer localities with $n = 4$ than with $n < 4$, so that the estimates $\hat{m}^{NP}(4)$ and $\hat{m}^{CF}(4)$ are less precise than the other point estimates presented in the table. To the author's knowledge, there is as of yet no distribution theory for the non-parametric control function estimator. Blundell and Powell (2003), in an illustrative application, present bootstrapped standard errors.

[22]This is, for instance, the case in a Salop (1979) style model.

Log Price Regressions (2)

	2SLS	OLS
C	1.544	1.685
	(0.015)	(0.004)
MTM	0.015	0.004
	(0.003)	(0.001)

Table 4: OLS and 2SLS point estimates; standard errors in parenthesis

point of view, accounting for the endogeneity of number of competitors by using suitable instruments appears to be critical. While OLS in the specification using fascias still predicts a 5 percent price rise from increased market concentration, the OLS estimates using number of multiplexes as concentration measures yield predictions that are below the 5 percent threshold that is the typical benchmark used by competition authorities for market definition, so this effect may be considered economically insignificant.

3.4 Alternative Specification

An alternative view of the market might be cast in terms of firms' pricing decision, setting prices with an eye to local wealth. This view is essentially tacit about the possible endogeneity of the number of competitors. The deprivation index would appear in it as a covariate in the main estimating equation. Hence, while still being exogenous and, hence, an instrument for itself, the deprivation index can no longer act as instrument for the number of competitors, should this variable be deemed endogenous. In this case, the only available instrument in the data is the local population. It is worth noting that this model is inconsistent with the consumer choice model as part of the structural econometric model outlined in section 2 above, unless income is taken to influence consumers' generalized cost. There, consumers make choices by minimizing generalized costs, and income does not feature in this decision problem.

Table 5 provides the corresponding estimation results for this model. Across all

Log Price Regressions (3)

	IV			OLS		
C	1.882	1.879	1.980	1.876	1.876	1.992
	(0.021)	(0.020)	(0.024)	(0.019)	(0.019)	(0.020)
DI	-0.006	-0.006	-0.006	-0.006	-0.006	-0.005
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
MC1	0.016			-0.001		
	(0.0318)			(0.010)		
MC2		0.005			7.18e-06	
		(0.010)			(0.003)	
FASCIA			-0.027			-0.044
			(0.022)			(0.010)

Table 5: IV and 2SLS point estimates; standard errors in parenthesis

estimated model specifications, there appears to be little evidence of any competitive constraint on prices by number of local competitors, once deprivation has been controlled for. One structural model that might support this specification is one in which the various cinema operators do not compete. The plausibility of this model, therefore, critically hinges on other evidence that might support the pricing strategy that the model stipulates. Such considerations are discussed in the following section.

4 Policy Implications

The use of price concentration analyses for industries with differentiated products has been critizised in the past. Baker and Bresnahan (1992) point out that in differentiated product industries the concentration measure (based on a given market boundary) does not take into account substitutes just outside the market. This criticism is based on the same reasoning that suggests that diversion ratios may be more informative than market shares in such industries. While this is an important point it is not enough to dismiss the relevance of price concentration studies in such contexts. The illustration showed that in horizontally differentiated markets

such as the cinema industry a price concentration analysis can inform both market definition and the competitive assessment of a case. This is a case where differentiation is (only) geographical and therefore the model can be extended to test different hypotheses (different distances from a multiplex) with relative ease, albeit under symmetry assumptions the plausibility of which will hinge on the specific application. When there is more than one dimension of differentiation (for example in terms of quality as well as location), the procedure may be more complex but the relevance of these elements of differentiation, subject to data availability, can be tested and therefore inform both market definition and competitive assessment.[23]

The importance of a fully developed functional form for the econometric model is well recognized as it allows the decision maker to assess the implications of different regression specifications. In this case the structural form presented in section 2.1. highlights the simultaneity problem. Bishop and Walker (2002) argue that, where this problem is evidenced by positive tests of simultaneity, instrumental variables should be used. They suggest that this is not a serious issue as the downward bias of the estimates reflects the constraints placed by entry.[24] This view has an intuitive appeal. However, typically there is no information in this context on the time that it takes for entry to restore a competitive outcome, and therefore an antitrust authority assessing adverse effects to competition within a limited (e.g. 2-5 years) time horizon may want to try and gain unbiased estimates.

Fully developing a structural model also facilitates the choice between alternative specifications. The model in section 2.1 posits that there is some competition between the horizontally differentiated firms which depends on their location. The econometric model built on that specification can then be used to test different assumptions about the extent of the geographic (and possibly the product) market in the specific case at hand. The alternative specification in the illustration argues for

[23]For example, the 2000 Competition Commission Supermarket inquiry conducted a price concentration analysis that included dummy variables for different types of stores to capture differences in the product range or other quality measures.

[24]The direction and magnitude of the bias depends on its source and generally is not obvious. For example, if it is due to factors affecting marginal costs, then, as an increase in marginal costs shifts up both $\pi(\cdot)$ and $\eta(\cdot)$, it could even induce a positive OLS estimate for the coefficient on n in $\pi(\cdot)$.

the inclusion of a variable as a regressor and the contextual rejection of its validity as an instrument. How can the antitrust authority choose in this context? The choice can arguably be based on theory and evidence. The implication of a different specification should be traced back to the underlying structural model. In the illustration the inclusion of a variable linked to the level of wealth is incompatible with the model in section 2.1[25] and may be compatible for example with a model where each firm is a local monopoly and prices according to the demand it faces. Once this distinction is clarified other sources of evidence can be used to inform the decision. For example internal strategy documents could show some evidence of monitoring of competitors' prices, or alternative of pricing according to local demand. Using internal documents jointly with econometric evidence is not new (see e.g. Baker (1999)), however specifying a structural model expands the set of decision relevant evidence, broadly defined, that can inform a competition authority.

5 Conclusions

This paper presents a semi-nonparametric structural econometric model for price-concentration analyses, which play a critical role in most competition inquiries. This model goes well beyond conventional reduced form approaches. The paper illustrates this approach in an application under data conditions typical in competition inquiries. The are clear advantages to the advocated econometric methodology. The approach rests of relatively mild functional form assumptions and hence enjoys considerable robustness. Moreover, the estimation results can be given a sound structural interpretation, enabling the authority to reach more informed conclusions. And, furthermore, the estimation methodology is immune to biases that arise as a consequence of jointly endogenous covariates. These benefits should provide assurance to practitioners and enhance the evidentiary value of price concentration analyses in competition inquiries.

[25]This implicitly assumes that income or wealth does not affect consumer's conditional indirect utilities, e.g. through the "choosiness" parameter t.

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