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Abstract

This paper compares the estimates of the two most widely used non-structural models for market power measurement in banking, namely the conduct parameter method and the revenue test, as applied to a panel of Greek banks over the period 1993-2004. We also propose a dynamic reformulation of these models within a panel data context, in order to address possible statistical problems associated with the dynamic nature of bank-level data. The results suggest that both static methods provide lower estimates of market power relative to their dynamic counterparts. Therefore, the inclusion of some dynamics in the models, even though it increased estimation complexity, helped to reveal some collusive behavior of banks.

JEL Classification: G21; L10; P20

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1. Introduction

In the banking sector, prudential regulation and competition policy are in many respects intertwined, while the soundness and stability of the system may in various ways be influenced by the degree of competition and concentration. Enhanced competition may have a deleterious impact on stability if it causes banks' charter value to drop, thus reducing the incentives for prudent risk-taking behavior. A more concentrated system, inasmuch as it implies the presence of a few relatively large banks, is more likely to display a "too big to fail" problem, by which large banks increase their risk exposure anticipating the unwillingness of the regulator to let the bank default in the event of insolvency problems (Hughes and Mester, 1998). Any market failure, inefficiency, or anticompetitive conduct among banks is likely to impose more severe costs throughout the economy than would similar defects in many other industries; thus it becomes particularly important to understand the causes and consequences of competition in the banking industry. In this respect, the first and most important step is the robust estimation of the degree of imperfectly competitive conduct.

Recent trends in empirical industrial organization have popularized the use of non-structural approaches to measure market power. These approaches, that came to be known as the new empirical industrial organization (NEIO) models, provide empirically implementable techniques for the analysis of non-competitive behavior in production and cost structure. The way to these approaches has been paved by a number of forerunners, with the two most popular models being those of Bresnahan (1982) and Lau (1982), and Panzar and Rosse (1987). Bresnahan and Lau (BL), in a line of research termed conduct parameter method (CPM) thereafter, parameterize the extent to which firms perceive a distinction between marginal revenue and price. On the other hand, in the Panzar and Rosse (PR) model, market power is measured by the extent to which changes in factor prices are reflected in revenue. In other words, this revenue test (RT) involves estimating a reduced-form equation relating gross revenue to a vector of input prices and other control variables. While each of the two approaches nurtures its own theoretical discourse, they should not be viewed as mutually exclusive but, more eclectically, as complementary tests.

Whereas there is fairly extensive application of these models to banking, there is a limited body of work that compares their results. The purpose of this study is to fill this gap, using a panel dataset of Greek banks during the period 1993-2004. Furthermore, we propose a dynamic reformulation of the static versions of the CPM and the RT within a dynamic panel data context. The most common motivation for also using a dynamic approach is the statistical importance of accounting for short-run dynamics in the data. Further, the formulation solves the inference problem when using non-stationary data (Steen and Salvanes, 1999). Finally, the dynamic nature of the Greek banking industry and certain changes in the regulatory environment may bias the resulting implications if only static models are considered.

We present and describe some important dynamic factors of the Greek banking sector, such as the patterns of consolidation and concentration, the changes in the regulatory framework, and the liberalization process that occurred during the sample period. The presence of these factors may lead to different adjustment costs, which in turn make static models inadequate. Indeed, the results suggest that both static methods (CPM and RT) tend to provide lower estimates of market power compared to their dynamic counterparts. In light of recent critiques of the NEIO approach to measuring anticompetitive conduct (e.g. Corts, 1999), this has important implications for public and private policy-makers alike.

The rest of the paper is structured as follows. Section 2 presents the two theoretical non-structural models applied in the current study, and Section 3 is then devoted to the analysis of the empirical static and dynamic versions of these models. Section 4 outlines the institutional structure of the Greek banking system and offers a discussion on the dataset. Section 5 presents and discusses the empirical evidence of applying the models to the Greek banking sector, while some conclusions are offered in the final section.

2. Theoretical framework

The literature on the measurement of competition can be divided into two major streams: structural and non-structural. The structural approach embraces the structure-conduct-performance and the efficiency hypotheses. These two models investigate, respectively, whether a highly concentrated market causes collusive behavior among the larger banks, resulting in superior market performance, or whether it is the efficiency of larger banks that enhances their performance. Although these hypotheses lack formal theoretical support by traditional microeconomic theory, they have frequently been employed empirically in the banking industry (e.g. Evanoff and Fortier, 1988; Bourke, 1989).

However, owing to several deficiencies arising from the application of the structural approach,[1] developments in industrial organization, as well as the recognition of the need to endogenize the market structure, many empirical studies followed a new course. The

[1] Among these are the interpretation of the positive relationship between profitability and concentration and various other methodological issues (see Bresnahan, 1989).

novelty to competition evaluation has emerged under the impulse of the NEIO approach (Carlton and Perloff, 2005). This approach, pioneered by Iwata (1974) and strongly enhanced by the papers of Bresnahan (1982, 1989), Lau (1982), and Panzar and Rosse (1987), tests competition and the use of market power, and stresses the analysis of banks' competitive conduct in the absence of structural measures. Specifically, each of these techniques attempts to measure the competitive conduct of banks without explicitly using information on the structure of the market.

The first model, the Iwata model, allows the estimation of conjectural variation values for individual banks supplying a homogeneous product in an oligopolistic market (Iwata, 1974). This measure, to the best of our knowledge, has been applied to the banking industry only once, by Shaffer and DiSalvo (1994), for a duopolistic banking market (in South Central Pennsylvania).[2] They find that banks' conduct is imperfectly competitive, but closer to perfect competition than one would expect, given the very high degree of concentration in the market.

2.1. The conduct parameter method (CPM)

The second model, which has been applied to the banking sector in a number of studies, is based on the procedure first suggested by Bresnahan (1982) and Lau (1982) and further elucidated in Bresnahan's (1989) survey of the NEIO.[3] It requires the estimation of a simultaneous-equation model, where a parameter representing the degree of market power of firms is included. The basis of the test is the established principle that, in equilibrium, profit-maximizing firms will choose prices or quantities such that marginal

[2] Applying this model to the banking industry is difficult, particularly where micro data for the cost and production structure for homogeneous products are scarce.

cost equals their perceived marginal revenue, which coincides with the demand price under perfect competition and with the industry's marginal revenue under perfect collusion. As such, the key parameter in this test is interpreted as the extent to which the average firm's perceived marginal revenue schedule deviates from the demand schedule, thus representing the degree of market power actually exercised by the firms in the sample.

In this respect, consider a non-competitive industry in which N banks produce a homogeneous output Q, facing a market demand function of the following stylized form:

$$Q = D\,(P, Z, \alpha) + \varepsilon \tag{1}$$

where D is the demand function, P is the market price of industry output Q, Z is a vector of exogenous variables affecting demand (often including some variable measuring the general economic activity), α is the demand parameter vector, and ε is a stochastic disturbance.

On the supply side, the representative bank i, ($i \in \{1, 2, \ldots, N\}$) is assumed to maximize profits by solving the following one-shot game in output level:

$$max\ p_i\, q_i - C\,(q_i, \omega_i) \tag{2}$$

where (for the ith bank), q_i is the output level, p_i is the respective price imposed, C is the cost function (which for now is homogeneous across all banks in the industry), and w_i is the price vector of inputs. The optimality condition corresponding to this problem is given by the following inverse supply relation:

$$P = MC(q_i, w_i) - \left(\frac{\partial Q}{\partial q_i}\frac{q_i}{Q}\right)\left(\frac{\partial P}{\partial Q}Q\right) \tag{3}$$

where MC is the marginal cost function.

[3] For a review of these studies see Shaffer (2004).

As is evident from the last equation, an integral part of the solution is the elasticity concept $\lambda_i \equiv \dfrac{dQ/Q}{dq_i/q_i}$, that is the conjectural variation coefficient of the NEIO literature. As λ_i moves farther from zero, the conduct of bank i moves farther from that of a perfect competitor. Thus, the (average) conjectural variation coefficient will reveal what kind of imperfectly competitive behavior characterizes the market, and there is no need to impose any *a priori* restriction on it. In other words, it is not necessary to assume a certain conduct beforehand and test for its propriety.

Moreover, in Eq. (3), $Q(\partial P/\partial Q)$ represents the semi-elasticity of market demand, that is $h(\cdot) \equiv \dfrac{Q}{\partial Q/\partial P}$, which is a function of aggregate output and other exogenous variables (Bresnahan, 1982). Shaffer (1993) describes the banking industry's marginal revenue function as industry price P plus $h(\cdot)$. Yet, a bank's perceived marginal revenue is generally not equal to the industry's marginal revenue. In fact, the perceived marginal revenue for bank i equals $P + \lambda_i h(\cdot)$. The range of possible values of the conjectural variation elasticity λ_i is given by (0,1). In the special case of the Cournot behavior, $\partial Q/\partial q_i = 1$, and λ_i is simply the output share of the ith bank. In the case of perfect competition, $\lambda_i = 0$; under pure monopoly, $\lambda_i = 1$; and, finally, $\lambda_i < 0$ would imply pricing below marginal cost and could result, for example, from a non-optimizing behavior of banks. Clearly, aggregation implies that the average value of λ_i across all banks equals the industry's conjectural elasticity, defined as L, the latter having the same properties as λ_i. Thus, this framework provides a benchmark, which can be used to identify the actual underlying market structure.

2.2. The revenue test (RT)

The PR (1987) approach (initially developed by Rosse and Panzar, 1977) for measuring market power relies on the premise that each bank will employ a different pricing strategy in response to a change in input costs, depending on the market structure in which this bank operates. In other words, market power is measured by the extent to which changes in factor prices (unit price of funds, capital, and labor) are reflected in revenue. The authors define a measure of competition, the *H*-statistic, as the sum of the elasticities of the reduced-form revenue function with respect to factor prices. Thus, the *H*-statistic represents the percentage variation of the equilibrium revenue derived from an infinitesimal percent increase in the price of all factors used by the firm.

Panzar and Rosse (1987) show that this statistic can reflect the structure and conduct of the market to which the firm belongs. They assert that the *H*-statistic is negative when the competitive structure is a monopoly, a perfectly colluding oligopoly, or a conjectural variations short-run oligopoly; an increase in input prices will increase marginal costs, reduce equilibrium output, and subsequently reduce revenue.[4] Under perfect competition, where banks' products are regarded as perfect substitutes of one another, the Chamberlinian model, based on free entry of banks and determining not only the output level but also the equilibrium number of banks, produces the perfectly competitive solution, as demand elasticity approaches infinity. Thus, in this case, the *H*-statistic is equal to unity. Shaffer (1982) shows that the *H*-statistic is also unity for a natural monopoly operating in a perfectly contestable market and also for a sales-maximizing firm that is subject to

[4] In the case where the monopolist faces a demand curve of constant price elasticity (i.e. $e > 1$) and where a constant returns to scale Cobb–Douglas technology is employed, Panzar and Rosse proved that H is equal to e-1. Hence, apart from the sign, the magnitude of H may also be of importance, as H yields an estimate of the Lerner index of monopoly power $L = (e-1)/e = H/(H-1)$ (see Bikker and Haaf, 2002).

breakeven constraints. Consequently, an increase in input prices raises both marginal and average costs without altering the optimal output of a bank. Exit from the market will evenly increase the demand faced by each of the remaining banks, thereby leading to an increase in prices and total revenue by the same amount as the rise in costs (i.e. demand is perfectly elastic). Finally, if the *H*-statistic is between zero (inclusive) and unity (exclusive), the market structure is characterized by monopolistic competition. Under monopolistic competition, potential entry leads to contestable market equilibrium, and income increases less than proportionally to the input prices, as the demand for banking products facing individual banks is inelastic.

When applying the RT to assess banks' market conduct, various assumptions about banks' production activity have to be made. First, the methodology requires assuming that banks are treated as single-product firms, producing intermediation services by using labor, physical capital, and financial capital as inputs. Second, one needs to assume that higher input prices are not associated with higher quality services that may generate higher revenue, since such a correlation may bias the computed *H*-statistic. Yet, if one rejects the hypothesis of a contestable competitive market, this bias cannot be too large (Molyneux et al., 1996). Among other underlying assumptions inherent in the PR model, we can mention that: (a) banks are profit-maximizing firms; (b) the performance of these banks needs to be influenced by the actions of other market participants; (c) the cost structure is homogeneous; and (d) the price elasticity of demand is greater than unity (see also De Bandt and Davis, 2000). Finally, a limitation of the RT is that any monopsony power or upward-sloping aggregate supply curve of any essential input (such as deposits) would render econometric identification an issue. Monopsony power, by tending to drive up input prices (and hence equilibrium revenues) as a function of scale, would tend to yield higher values of H and thereby mask any market power present on the output side, in contrast to the CPM (Shaffer, 2004).

Despite these assumptions, the RT is a valuable tool in assessing market conditions, mainly owing to its simplicity and transparency, without lacking efficiency. Moreover, data availability becomes much less of a constraint, since revenue is more likely to be observable compared to output prices. Also, by utilizing bank-level data, this approach allows for bank-specific differences in the production function. In addition, the non-necessity to define the location of the market *a priori* implies that the potential bias caused by the misspecification of market boundaries is avoided; hence for a bank that operates in more than one market, the H-statistic will reflect the average of the bank's conduct in each market.[5]

3. Empirical specification

This section aims to identify robust econometric procedures for applying both non-structural models of market power described above. As Bresnahan (1989) states, only econometric problems, not fundamental theoretical problems, could cloud inference on the empirical results of these models. To this end, the econometric methodology to be followed is afforded particular consideration.

3.1. CPM

[5] Owing to the reasons described above the revenue test has been extensively applied to the banking industry. For a thorough review see Mamatzakis et al. (2005).

The general empirical problem in studies relying on a CPM is the identification of the elasticity concept L. Using the results of Bresnahan (1982) and Lau (1982), Shaffer (1999) suggests a structural econometric model that combines separate demand and supply functions including cross-equation restrictions. A necessary and sufficient condition for the identification of L is that the demand function must not be separable in at least one exogenous variable that is excluded from the marginal cost function. Following this view, we specify the following linear demand function:

$$Q = a_0 + a_1 P + a_2 Y + a_3 Z + a_4 PY + a_5 PZ + a_6 YZ \tag{4}$$

where Y is an exogenous variable that affects demand. Such a demand function, which includes three cross-product terms to improve flexibility,[6] can be interpreted as a first order local approximation of the true aggregate demand function. The exogenous variable Z is critical for the solution of the identification problem and is generally chosen as the price of a substitute. Notice that the demand function is specified at the aggregate level. Since we are considering the single product case, $Q = \sum_i q_i$ is well defined.[7]

Turning to the supply side of the model, we follow Shaffer (1999) who specifies a marginal cost function out of a generalization of the minflex Laurent functional form (which in turn is a generalization of a standard translog cost function) as follows:

$$P_i = -L\left[Q/(\alpha_1 + \alpha_4 Y + a_5 Z)\right] + \frac{C_i}{q_i}[b_0 + b_1 \ln q_i + b_2 (\ln q_i)^{-2} + \sum_{j=3}^{5} b_j \ln \omega_{ij}] \tag{5}$$

[6] Applications of the Bresnahan-Lau method in the banking industry have favored linear demand functions with one or two cross-product terms (Shaffer, 1999; Toolsema, 2002; etc.). Sjoberg (2004) uses a log-linear demand function, with one cross-product term (namely price of output times the exogenous variable), while Uchida and Tsutsui (2005) use a log-linear demand function with a number of explanatory variables, mainly corresponding to the quality of loans. We too have undertaken a log-linear demand function; however the changes in the coefficients on L are negligible.

[7] Given that data on a substitute price of individual banks are really hard to find (one would require bank-level data on the price of securities) we opt for an aggregate demand function.

Regarding the functional form imposed on the supply relation, the differentiation in the literature is broader than in the respective one imposed on the demand relation. According to the purpose of each study, the cost function has been given a linear, minflex Laurent, translog, generalized Leontief, or Fourier functional form. The appropriate choice rests on the assumptions of each study, the number of outputs specified, and the level of flexibility required. We have relied on the generalization of the minflex Laurent since, as Barnett and Lee (1985) suggest, even though models produced from second order Taylor series expansions (such as the translog and the generalized Leontief) can be rendered both flexible and regular at the median data point, that regularity usually quickly disappears as the recent-periods boundary of time series datasets is approached. The minflex Laurent model also can be rendered both flexible and regular at the median data point. However, doing so with this function assures simultaneously that the model's region of regular behavior actually expands as the recent period boundary is approached. As a result, with time series (and hence with panel) data, the minflex Laurent possesses substantial advantages over the Taylor series specifications for conventional modeling purposes.[8]

Most of the usual properties of a cost function pose no specific requirements for the parameters of the above supply equation. Symmetry and concavity do not involve any of the coefficients,[9] while monotonicity involves but does not constrain the coefficients on the input prices. The property of linear homogeneity in factor prices implies $\sum_{j=1}^{3} b_j = 0$. We tested for linear homogeneity after performing the regressions; however, in all models the

[8] For the transformation of the minflex Laurent model to our partially restricted Laurent specification and its advantages, see Shaffer (1999).

[9] Symmetry in the coefficients of output is irrelevant as we consider the one-output case. Symmetry in the coefficients of inputs would be necessary if we estimated different parameters for a *pair* of cross-product terms of the flexible functional form. Here we rely on estimation of the marginal cost function.

above hypotheses are not rejected at the 5 per cent level. Therefore, we do not impose any initial restrictions. Finally, note that the measurement of the term C_i / q_i, as well as the dependent variable at the bank level, while not consistent with the profit-maximizing solution (since a quantity game was considered), allows for heterogeneity in marginal costs and in the price setting policy, respectively (on this issue see Tsutsui and Uchida, 2002 and Sjoberg, 2004).

For L to be correctly specified in Eq. (5) it is necessary to treat the input prices as exogenously given. This assumption seems reasonable, at least as far as the markets for labor and physical capital are concerned, because banks face intense competition for these inputs from other banks as well as non-bank firms. The market for funds will also be treated as competitive based on the argument that depositors today have many other tempting saving options (such as government and corporate bonds and the stock markets), which exert a competitive pressure upon banks' deposit rate policy. To the extent that this is not true, i.e. that bank have some degree of market power in the deposit market, it has been shown (Shaffer, 1999) that this will not escape identification and that it will simply be misattributed to the asset side (consequently L will be strengthened).

For empirical implementation purposes, the CPM has to be embedded within a stochastic framework. Thus, we assume that Eqs. (4) and (5) are stochastic owing to errors in optimization. Introducing the additive disturbance terms in System {(4), (5)}, the latter is specified as:

$$\begin{aligned} Q_t &= a_0 + a_1 P_t + a_2 Y_t + a_3 Z_t + a_4 P_t Y_t + a_5 P_t Z_t + a_6 Y_t Z_t + \varepsilon_t \\ P_{it} &= -L\left[Q_t /(\alpha_1 + \alpha_4 Y_t + a_5 Z_t)\right] + \frac{C_{it}}{q_{it}}[b_0 + b_1 \ln q_{it} + b_2 (\ln q_{it})^{-2} + \sum_{j=3}^{5} b_j \ln \omega_{jit}] + u_{it} \end{aligned} \tag{6}$$

The disturbance vectors ε and u are assumed to be iid as multivariate $N \sim (0, \Sigma)$, where Σ is a positive definite matrix.

As discussed previously, equations in the latter system are interrelated, because of the underlying CPM, through the endogeneity of the semi-elasticity of market demand, which appears in the supply equation. Thus, we should employ a system estimator such as nonlinear three-stage least squares (3SLS), generalized method of moments (GMM)[10] or full information maximum likelihood (FIML). We have applied all three methodologies, with the results being similar, yet we have resorted to the 3SLS method (as most of the relevant literature) for two main reasons. First, the FIML estimator is the asymptotically efficient estimator only under the assumption of normally distributed residuals (see Amemiya, 1977). We tested for normality using the Jarque-Bera statistic, which is significant at the 1 per cent level, thus ruling against normality. Second, use of different instruments does not significantly alter the results in the 3SLS case, while in the GMM case the variability is larger. Therefore, we proceed with the estimation of System (6) using a 3SLS procedure.

A crucial feature of the CPM is that the variables involved are usually characterized by short-run dynamics, which are not accounted for in the static equations. Further, a reformulation of the static model to a dynamic one may help with the inference problem when using non-stationary data or when severe autocorrelation in the demand equation is present (Toolsema, 2002). Steen and Salvanes (1999) developed a dynamic version of the Bresnahan and Lau model, based on an error-correction (ECM) framework. They applied it to the French market for fresh salmon, using time series data. Toolsema (2002) opted for applying this model to the Dutch consumer credit market. However, multicollinearity in the demand equation and difficulty in identifying a purely exogenous Z variable (so as the demand function will not be separable in Z) are problems that she could not overcome, thus failing to estimate the model. We have closely followed her approach and indeed multicollinearity seems to be a major problem when using time series data.

Given the above, we opt for the following dynamic representation of the demand equation (autoregressive-distributed lag model):

$$\begin{aligned} Q_t &= a_0 + a_l Q_{t-1} + a_1 P_t + a_{1l} P_{t-1} + a_2 Y_t + a_{2l} Y_{t-1} + a_3 Z_t + a_{3l} Z_{t-1} \\ &\quad + a_4 (PY)_t + a_{4l} (PY)_{t-1} + a_5 (PZ)_t + a_{5l} (PZ)_{t-1} + a_6 (YZ)_t + a_{6l} (YZ)_{t-1} + \varepsilon_t \end{aligned} \tag{7}$$

which is derived as a static conditional demand equation, assuming the same technology as in the static case and with dynamics resulting from an AR(1) disturbance.[11] As we do not impose the implied common factor restrictions, the dynamics may be thought of as an empirical approximation to some more general adjustment process (Blundell and Bond, 1998). This is equivalent to identifying the short-run relationships, which are of main interest in this study.[12]

Similarly, we specify the dynamic supply equation of the CPM as:

$$\begin{aligned} P_{it} &= -LQ^*_{it} - L_l Q^*_{i,t-1} + \lambda_0 P_{i,t-1} + c_{it}[b_0 + b_1 \ln q_{it} + b_2 (\ln q_{it})^{-2} + \sum_{j=3}^{5} b_j \ln \omega_{itj}] + \\ &\quad c_{i,t-1}[b_{0l} + b_{1l} \ln q_{i,t-1} + b_{2l} (\ln q_{i,t-1})^{-2} + \sum_{jl=3}^{5} b_{jl} \ln \omega_{ji,t-1}] + u_{it} \end{aligned} \tag{8}$$

[10] Reference here is made to the GMM estimator developed by Hansen (1982).

[11] To decide that lag length is equal to 1, we used the Ljung-Box Q-statistic (Ljung and Box, 1979). The results ensured that no serial correlation was present in the residuals (the maximum lag length used was 3). Indeed, we did not expect a higher order autocorrelation due to the use of annual data.

[12] Alternatively, we may rely on an ECM specification (as in Steen and Salvanes, 1999), which could allow distinction between the long- and short-run effects. However, this (*i*) would preferably require panel

where $Q_t^* = Q_t / (\alpha_1 + \alpha_4 Y_t + a_5 Z_t)$, $Q_{t-1}^* = Q_{t-1} / (\alpha_1 + \alpha_4 Y_{t-1} + a_5 Z_{t-1})$, and $c_{it} = C_{it} / q_{it}$.[13] Eq. (8) is non-linear in its parameters and therefore requires a non-linear estimation procedure. Since we cannot simultaneously estimate the System {(7), (8)} (a non-linear system estimator for dynamic panels is not available) we proceed in two steps. First, to account for the simultaneity problem, Eq. (7) is estimated using two-stage least squares (2SLS), as in Steen and Salvanes (1999), with input prices and a time trend as instruments. In order to test the validity of these instruments we used a Sargan test of over-identifying restrictions, whose statistic has an approximately chi-square distribution with degrees of freedom equal to the number of instruments minus the number of regressors. The statistic rejects the hypothesis of over-identifying restrictions at the 10 per cent level of significance (p-value = 0.203).

Then, after having calculated Q^* and its lag, we estimate Eq. (8). As discussed above, the supply equation is estimated using panel data. Estimation of a dynamic panel data model in the supply equation may still present considerable robustness problems due to the non-stationarity of the variables involved and the possible existence of cointegrating relationships between them.[14] We cannot robustly test for stationarity here, nor use panel cointegration techniques, owing to the small time dimension of the panel. For consistent and efficient estimates of the supply equation we apply the system GMM approach proposed by Arellano and Bover (1995) and Blundell and Bond (1998). This estimator

cointegration techniques, which are difficult to apply here because of the relatively small time dimension of the panel, (*ii*) is more demanding in terms of degrees of freedom, and (*iii*) is not directly comparable to the dynamic Panzar-Rosse model described below.

[13] Again the Ljung-Box Q-statistic ensures serial correlation of order not higher than 1.

[14] Binder et al. (2003) and Baltagi (2005), among others, suggest that in panels with a small time dimension and a larger cross-sectional dimension (which is usually the case in the relevant literature), instrumental variables and GMM estimators based only on standard orthogonality conditions break down if the underlying time series contain unit roots.

combines the T-2 equations in differences with the T-2 equations in levels into a single system. It uses the lagged levels of dependent and independent variables as instruments for the difference equation and the lagged differences of dependent and independent variables as instruments for the level equation. This estimation procedure is especially appropriate when: (*i*) the cross sectional dimension is large compared to the time dimension of the panel; (*ii*) some explanatory variables are endogenous; and (*iii*) unobserved bank-specific effects are correlated with other regressors; all three criteria are relevant in the present analysis. Also, this estimator does not break down in the presence of unit roots (for a proof see Binder et al., 2003). We choose the two-step estimator, since it is asymptotically more efficient than the respective one-step estimator, and we account for its downward bias by using the finite-sample correction to the two-step covariance matrix derived by Windmeijer (2000). We use Z and a linear time trend as "GMM-style" instruments (for a discussion see Arellano and Bond, 1991).

To determine whether our instruments are valid in the system GMM approach, we use the specification tests proposed by Arellano and Bond (1991) and Arellano and Bover (1995). First, we apply the Sargan test, a test of over-identifying restrictions, to determine any correlation between instruments and errors. For an instrument to be valid, there should be no correlation between the instrument and the error terms. The null hypothesis is that the instruments and the error terms are independent. Thus, failure to reject the null hypothesis could provide evidence that valid instruments are used. Second, we test whether there is a second order serial correlation with the first differenced errors. The GMM estimator is consistent if there is no second order serial correlation in the error term of the first-differenced equation. The null hypothesis in this case is that the errors are serially uncorrelated. Thus, failure to reject the null hypothesis could supply evidence that valid orthogonality conditions and instruments are used.

3.2. RT

As in Shaffer (2004) we derive the *H*-statistic using the following specification of the reduced-form revenue equation for a panel dataset:

$$\ln TR_{it} = \beta_0 + \beta_1 \ln w_{1,it} + \beta_2 \ln w_{2,it} + \beta_3 \ln w_{3,it} + \beta_4 \ln TA_{it} + \varepsilon_{it} \tag{9}$$

where *it* is the subscript indicating bank *i* at time *t*, *TR* stands for a bank's real total revenue, w_1, w_2 and w_3 are the three input prices, and *TA* stands for real total assets. The log specification is used to improve the regression's goodness of fit and to reduce possible simultaneity bias (De Bandt and Davis, 2000). Molyneux et al. (1996) found that a log-linear revenue equation gives results similar to those of a more flexible translog equation. The revenue equation is interpreted as a reduced form rather than as a structural equation (Shaffer, 2004).

As discussed above, the *H*-statistic is equal to the sum of the elasticities of total revenue with respect to the three input prices, i.e. $H = \beta_1 + \beta_2 + \beta_3$. The *H*-statistic is interpreted here as a continuous measure of the level of competition, in particular ranging between 0 and 1, with higher values of the statistic indicating stronger competition. This does not follow automatically from the Panzar and Rosse (1987) study, which concentrates only on testing the hypotheses $H = 0$ (or more precisely $H \leq 0$) and $H = 1$. However, it can be shown that under stronger assumptions (in particular under the assumption of a constant price elasticity of demand across banks) our interpretation of the *H*-statistic is correct. The Chamberlinian equilibrium model provides a simple link between the *H*-statistic and the number of banks, and thus between market behavior and market structure. Vesala (1995) proves that the *H*-statistic is an increasing function of the demand elasticity *e*, that is, the less market power is exercised, the higher the *H*-statistic becomes. This implies that the *H*-statistic is not used solely to reject certain types of market behavior, but that its magnitude serves as a measure of competition. One of the general assumptions underlying the Chamberlinian equilibrium model mentioned above is that *e* is a non-decreasing function of the number of rival banks. Vesala's result, together with this latter assumption, provides a positive (theoretical) relationship between *H* and the number of banks, or – in a looser interpretation – an inverse relationship between *H* and banking concentration. Hence, the more negative the *H*-statistic is, the larger is the monopoly markup, while the closer the *H*-statistic is to unity, the more competitive is the market (Vesala, 1995; Barajas et al., 2000).

Input prices are measured as in the CPM. As regards *TA*, a positive coefficient is expected, as a higher volume of output envisages greater revenue. Furthermore, causation may run from *TR* to assets if bank managers tend to retain marginal changes in earnings rather than distributing them to shareholders, thus raising assets, or if banks that expect to have better performance credibly transmit this information through expansion of the asset base. Therefore, *TA* should be treated as an endogenous variable and, consequently, an instrumental variable panel data estimation method is required. We resort to a two-stage least squares (2SLS) procedure under a random effects (RE) model.[15]

The dynamic extension of the Panzar-Rosse model is less demanding than the equivalent CPM, since it is linear in the parameters and does not require system estimation. Following Bond (2002), we specify an autoregressive-distributed lag model of the form:

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[15] The suitability of a RE model was tested against a fixed effects (FE) model, using a Hausman test. The results showed that the difference in coefficients is not systematic, thus providing evidence against FE.

$$\ln TR_{it} = \beta_0^{'} + \beta_l \ \ln TR_{i,t-1} + \beta_1^{'} \ln w_{1,it} + \beta_2^{'} \ln w_{2,it} + \beta_3^{'} \ln w_{3,it} + \beta_4^{'} \ln TA_{it} + \varepsilon_{it} \tag{10}$$

where $\ln TR_{i,t-1}$ is the lagged dependent variable in a logarithmic form. Once again, the *H*-statistic is obtained as $H^{'} = \beta_1^{'} + \beta_2^{'} + \beta_3^{'}$. Persistence in revenue may reflect impediments to product market competition, generating market power in output markets as well as better forecasting of industry and/or macroeconomic developments (for a related discussion on the persistence of bank profits, see Berger et al., 2000).

Eq. (10) is estimated using the dynamic panel data estimation method proposed by Blundell and Bond (1998). The instruments used are $\ln TR_{i,t-2}, \ln TR_{i,t-3}, ..., \ln TR_{i,1}; \ln TA_{i,t-2}, \ln TA_{i,t-3}, ..., \ln TA_{i,1}; \ln w_{ji,t-2}, \quad \ln w_{ji,t-3}, ..., \ln w_{ji,1}$ and $\Delta \ln TR_{i,t-2}, \Delta \ln TA_{i,t-2}$. Note that ln*TA* is specified as an endogenous variable, being instrumented in "GMM style" and symmetrically to the dependent variable ln*TR* (see Bond, 2002).

A critical feature of the *H*-statistic is that the test must be undertaken on observations that are in long-run equilibrium. The empirical test for equilibrium is justified on the grounds that competitive capital markets will equalize the risk-adjusted rate of returns across banks, so that, in equilibrium, rates of return should not be correlated statistically with input prices. Therefore, to test for equilibrium, one can calculate the *H*-statistic (H_n) using the rates of return, instead of total revenue, as the dependent variable in the regression equation. All authors use a regression relating return on assets to input prices. However, the argument also holds if the return on equity is used as the dependent variable instead (Molyneux et al., 1996; Yildirim and Philippatos, 2002; Bikker and Haaf, 2002). A value of $H_n < 0$ would show non-equilibrium, whereas $H_n = 0$ would prove equilibrium. However, if the sample is not in long-run equilibrium, it is true that $H < 0$ no longer proves monopoly, but it remains also true that $H > 0$ disproves monopoly or conjectural variation short-run oligopoly (Shaffer, 2004).

4. Data description and analysis

4.1. An overview of the Greek banking industry

Since the mid-1990s, the Greek financial and banking landscape has changed rapidly as a result of the new regulatory framework characterizing the market. In 2004, there were 62 credit institutions operating in Greece, a figure much higher than that observed in 1990, when only 39 credit institutions were in operation (Bank of Greece, 2006). The environment that emerged after 1993 gave impetus to the establishment and operation of new credit institutions, either domestic ones or branches of foreign banks. Foreign presence concentrated mainly in niche markets, specializing in areas such as shipping and corporate finance, private and personal banking, asset management, and capital market activities. In addition, in 1993 the Bank of Greece set the operational and supervisory framework concerning cooperative credit institutions, resulting in banks of this type getting established.

Thus, as of 2004, the Greek banking system comprises 21 Greek commercial banks, 23 foreign-owned banks (which constitute a subgroup of commercial banks), 16 cooperative banks, and 2 specialized credit institutions. Commercial banks incorporated in Greece have been the dominant group in the banking system. Indeed, these credit institutions hold a high market share, both in terms of assets (81 per cent), as well as in terms of loans (85 per cent) and deposits (82 per cent). On the other hand, the market share of foreign-owned banks stands at 10 per cent in terms of assets (9 per cent and 8 per cent for loans and deposits respectively), while the market share of the cooperative credit institutions remains very low (less than 1 per cent of aggregate balance sheet figures). The dominance of commercial banking can also be confirmed by the number of branches and employees. As of end 2004, Greek commercial banks have 2,953 branches in operation (out of 3,403 for all credit institutions), while the number of their employees stands at 51,741 (out of 59,337 employed in all credit institutions) (Bank of Greece, 2006).

A specific structural feature of the *status quo ante* of the Greek financial system, characterizing in particular the old banking regime, was the significant level of state intervention, which for a long time hindered competition and created a distorted market environment. Indeed, in the early 1990s, the state commercial banks controlled around 85 per cent of total commercial banking operations. Since then, a notable trend observed in the Greek banking sector was the privatization of several banks controlled by the Greek state, contributing to the enhancement of competition in the market. In the process, the number of directly or indirectly state-controlled banks[16] was reduced significantly, from 10 in 1993 to only 2 in 2004.[17]

While the number of commercial banks operating in the Greek banking system remained almost unchanged since 1993, the number of those banks' branches and employees has increased significantly (see Table 1). During this period, a number of new, mainly small, commercial banks opened and a series of mergers and acquisitions were

[16] The indirect control comes from the majority equity participation of public pension funds, municipalities and other funds, or from equity holdings of other state-owned or state-controlled banks.
[17] The largest credit institution, the National Bank of Greece, has come to a large degree into non-state ownership, and may be considered to operate largely on private economy criteria, while the fifth largest bank, Emporiki Bank (also known as Commercial Bank of Greece), is in the process of disentanglement of the Greek state.

undertaken, altering the level of bank concentration and substantially changing the structure of the Greek banking system. Specifically, especially during the first half of the 1990s, new private-owned foreign commercial banks were established, taking advantage of new products and services that were not available in the Greek market just a decade ago. Later on, other Greek commercial banks were established, primarily focusing on retail banking. Moreover, since the mid 1990s, several Greek banks have been involved into mergers and acquisitions, in order to become more efficient and obtain a size that would enable them to increase or, at least, maintain their domestic market shares, facilitate their access to international financial markets, and exploit any possible economies of scale. Most of them concerned the domestic market, including not only banks but also non-bank financial enterprises. Some large credit institutions opted to merge with their subsidiaries with a view to restructuring their activities and cutting back on their operating expenses.

These mergers and acquisitions have reversed the downward trend observed in bank concentration during the previous decade. The market share of the 5 largest credit institutions reached 65 per cent in terms of balance sheet aggregates in 2004 (being in similar and higher levels in the period 2001-2003), up from 57 per cent in 1997 (Bank of Greece, 2006). Similar are the conclusions if we use the Herfindahl-Hirschman Index (HHI).[18] Its value stands at 1,069 in 2004, up from 885 in 1997. In any case, both concentration ratios characterizing the Greek banking sector are much higher than the average European levels (the share of the 5 largest credit institutions stands, on average, at 40 per cent, while the HHI index reaches 569 in 2004) (European Central Bank, 2004, 2005).

[18] The Herfindahl-Hirschman index (HHI) is defined as the sum of squares of the market shares of all banks in the particular banking market.

Even though the Greek banking system is characterized by a relatively high degree of market concentration, the five larger Greek commercial banks would be classified as mid-size by European standards; only the first two banks are included among the top 100 European banks, while none in the top 150 credit institutions at a global level (Bank of Greece, 2005). This is principally due to the limited size of the domestic credit market and the absence of significant presence abroad.

To compete in the new financial landscape and strengthen their position in the market, Greek commercial banks are transforming themselves into financial groups, (*i*) adding subsidiaries such as insurance companies, brokerages, credit card companies, mutual fund firms, factoring companies and finance houses, so as to offer additional services, and (*ii*) expanding their activities abroad, principally to the South Eastern European region (Albania, Bulgaria, Former Yugoslav Republic of Macedonia, Romania, Serbia and Montenegro, and recently Turkey), via subsidiaries or through the establishment of branches. This latter trend signifies that Greek banks in the region have some comparative advantage, in the form either of access to capital markets, or of superior organization, know-how, and good understanding of local conditions.

The above-mentioned developments in the structure of the Greek banking system resulted in significant modifications in the balance sheet and profit and loss accounts. More notably, the ratio of net interest income to average total assets (i.e. the net interest margin) of Greek commercial banks increased considerably during the examined period, raising from 1.57 per cent in 1993 to 2.80 per cent in 2004.[19] Also, the proportion of

[19] Initially, net interest income was low, compared to other EU countries, mainly owing to the portfolio structure of the credit institutions (high proportion of public securities and cash and balance with central banks). Since then, banks' balance sheets have been restructured to make their financial positions sounder.

loans to total assets reached 63 per cent at 2004 (compared to 24 per cent in 1993), catching up rapidly with the average European levels.[20]

The high proportion of operating expenses is related to the specific features of the Greek banking system, such as the high number of branches of large banks and the fact that the products offered are relatively limited (Hondroyiannis et al., 1999). However, although Greek banks' operating expenses relative to their average total assets remain above the average European figures, they fell from 2.9 per cent to 2.3 per cent between 1996 and 2004. During this period, Greek credit institutions took important steps towards improving their efficiency by deploying modern information technology systems, cutting down on their operating costs and improving their organizational structure, while extending their scope of business by offering new products and services. Finally, Greek banks have increased their levels of loan loss provisions, mainly by reason of the significant credit expansion. Taking into account these developments, indications of a long-term downward trend in profitability will be observed, evident from the beginning of liberalization (towards the end of the 1980s) onwards (Gibson, 2005).

4.2. Data

The models specified above are used in order to examine the level of competition in the Greek banking industry during the period 1993-2004. All bank-level data are taken from Balance Sheet Accounts and Income Statements published annually by the Greek

[20] Several factors have been responsible for the high rates of growth of bank lending, including the relatively-high rate of growth of the Greek economy, the convergence of Greek lending rates to those in the rest of the euro area, the enhancement of competition among credit institutions, especially with regard to extending credit to households, and the release of commercial bank funds from the Bank of Greece due to the harmonization of reserve requirements in the Eurosystem.

banks included in the sample. The macroeconomic data (national income and Greek Treasury bond rates) were drawn from Eurostat and the Bank of Greece.

Our sample covers all Greek commercial banks, plus a foreign-owned credit institution, namely Bank of Cyprus. The institutions that do not publish profit and loss statements, i.e. branches of foreign banks and certain specialized credit institutions, are not included in the sample.[21] Specialized credit institutions and smaller cooperative banks are also excluded from the analysis, since (*i*) their operations differ substantially from those of the mainstream commercial banks and (*ii*) sometimes they have a different legal form. Last, we omit investment banks and banks focusing in corporate banking, since they fail to meet the criterion of being a well-rounded commercial banking institution (universal credit institution). The number of credit institutions included in the sample ranges from 13 to 23 commercial banks in each year of the examined period (see Table 2). In all examined years, the banks included in the sample accounted for a significant proportion of total banking assets (around 80 per cent).

The output variable in the CPM (namely Q when referring to the industry's total output and q when referring to the individual bank's output) is defined as the value of total bank assets in real terms (in million euros). The unit price of output (namely P at the market level and p at the bank level) is measured as interest income over total assets. The choice of cost, price and output variables follows either the intermediation or the production approach. According to the intermediation approach, banks are considered as financial intermediaries that combine deposits together with purchased inputs to produce bank assets. Total cost includes interest expenses and operating expenses, i.e. staff costs

[21] According to the Greek law governing the operation of corporations, foreign banks operating branches in Greece are not required to publish full-blown annual financial statements for their branch operations in Greece

and administrative expenses. In the alternative production approach, banks utilize capital and labor inputs to produce outputs of loans and deposit accounts. In this paper we follow the intermediation approach. The three input variables are defined as follows:

- *Labor* (L): Defined as total number of employees.
- *Physical capital* (K): Defined as fixed assets, including tangible fixed assets (land, lots, buildings and installations, furniture, office equipment, etc., less depreciation), as well as intangible fixed assets (goodwill, software, restructuring expenses, research and development expenses, minority interests, formation expenses, underwriting expenses, etc).
- *Total intermediated funds* (F): Include current accounts, savings accounts, time deposits, repurchase agreements, as well as alternative funding sources (e.g. retail bonds).

The unit prices of the three respective inputs (w_1, w_2, and w_3) are defined as follows:

- *Unit price of labor* (w_1): Ratio of personnel expenses to total labor. Personnel expenses include wages and salaries, social security contributions, contributions to pension funds, and other staff-related expenses.
- *Unit price of funds* (w_2): Ratio of interest expenses to total funds. Interest expenses include interest paid on deposits and other sources of funds.
- *Unit price of physical capital* (w_3): Ratio of administrative expenses to fixed assets. Administrative expenses include rents, service charges, security, information systems and communications, other office and insurance expenses, professional charges, publicity and advertising, and depreciation.

Finally, regarding the RT, *TR* stands for real total revenue (this refers to real total operating income, which includes interest income, dividend income, fee and commission income, gains less losses from securities, and other operating income). In Table 3 we report banking indicators of the variables described above for the period 1993-2004.

5. Empirical results

In Tables 4-7 we present the empirical results of the static and dynamic CPM and RT. There are four pairs of columns in each table. In the first, we provide the results of the basic models, as described in Section 4. In the second, the real variables are replaced by nominal ones (in order to be consistent with the part of the literature that uses nominal variables), while in the third we include a quadratic time trend, to capture any trends in output prices and revenues in the CPM and RT, respectively. Finally, in the fourth pair of columns, we add time dummies for all years (time fixed effects), thus modifying the model to a three-way error component (see Baltagi, 2005). The time dummy specification is more general than the linear trend specification, as it will pick up potential trends of the variables used and more complex bank-level patterns.[22] All these modifications are applied on a one-by-one basis and not in a cumulative manner; so, for example, the modified model containing time dummies for every year does not also contain a trend, and the amounts it uses are real. The lag length in the dynamic models is set to one, which rejects autocorrelation; hence higher order autocorrelation is not accounted for in the regressions (this is expected given the fact that the data are annual).

5.1. Static CPM

The results obtained by running the 3SLS estimation procedure on System (6) are presented in Table 4. The R-square statistics of both demand and supply relations, ranging from 0.89 to 0.96, indicate fine goodness of fit. The Durbin-Watson (DW) statistic reveals possible autocorrelation in the supply equation, which is a common problem in the literature (see Toolsema, 2002), whereas the DW statistic of the demand equation surprisingly rejects the hypothesis of serial correlation. All parameters of the demand relation were found to be statistically significant, which is a crucial factor in the identification of *L* (except the intercorrelation variable *YZ* when nominal amounts are used). In particular, the coefficient on *P* is negative and statistically significant, meaning that the demand function is decreasing in its own price, as expected. The negative sign on *Y* suggests that income refers to the ability to pay for the goods bought with consumer credit, in the sense that high income may imply less need for such credit (Toolsema, 2002). On the other hand the coefficient of the Greek government bond yield (*Z*) is positive and statistically significant, implying that our choice of *Z* is well suited as the price of a substitute.

In the supply equation, the parameters were not all statistically significant, which may be due to the dynamic nature of output and inputs in banking (especially when we include time dummies most *t*-statistics decrease). The coefficient on the unit price of labor contrasts our expectations, a fact that may be due to the labor surplus in the beginning of the period examined and the gradual reduction in labor expenses by Greek banks in order to improve their operating efficiency. In contrast, the coefficients of the other two inputs have the expected sign.

[22] The coefficients of the time dummies are not reported in the tables owing to space considerations.

The estimate of market power, L, is very close to zero in all cases and the corresponding t-statistics are small, indicating that L is not significantly different from zero. Thus, the static CPM analysis specifies that the Greek banking market is characterized by perfect competition. This means that the high degree of concentration characterizing the Greek banking industry reflects the efforts by the most efficient banks to take advantage of economies of scale and scope and does not necessarily influence competition in a negative manner. This conclusion is different from the results of studies of market power in banking sectors of several European countries, which generally find evidence of monopolistic competition or collusive conduct.

5.2. Dynamic CPM

The results obtained by examining the dynamic CPM are presented in Table 5. Once again, as in the static model, the parameters in the demand equation are strongly statistically significant at the 5 per cent level of significance, the only exception being the coefficient of the cross-term PZ when nominal values are applied. The R-square statistic of the supply relation attains values slightly higher compared to that of the static CPM, which indicates that the dynamic model contains somewhat more information, especially for input prices. The Durbin-Watson statistic is also improved, verifying that the autocorrelated errors can be made to disappear by incorporating additional dynamics; as Kennedy (2003) states, modern econometric models typically have a rich dynamic structure and only seldom involve autocorrelated errors. The signs of the estimated coefficients are unchanged; however the t-statistics of both bank output and inputs (especially for w_2) are strengthened.

The estimate of market power, *L*, is positive and close to zero (yet with a significantly higher *t*-statistic), except the one from the specification that includes time dummies. In the latter case, *L* is statistically significant at the 5 per cent level of significance. This is a striking result since we cannot accept the interpretation of perfect competition as in the static CPM, meaning that some form of collusive behavior characterizes the Greek banking sector.[23] Such an outcome may imply that the dynamics inherent in the use of bank-level data mask – to some extent – the market power exercised in the industry, a result effectively towards the same direction with the theoretical considerations of Corts (1999). To this end a dynamic CPM model may be a more appropriate specification.

5.3. Static RT

The competitive position tests of the static RT are presented in Table 6. In all models we include time dummies for the years 1999 and 2000 to account for the exceptional developments in the Greek stock market that took place during this period and led to a boom in bank revenues. The R-square statistic in all four estimated equations, ranging from 0.97 to 0.98, indicates fine goodness of fit. The coefficients on w_1 and w_3 are reported with the expected positive sign and they are statistically significant only when the trend or the time dummies are included in the specifications, a fact that provides evidence of improved stability of the equations. The coefficients on w_2 and *TA* are always positive and statistically significant.

[23] We run various robustness checks to verify this result, including a different estimation method and an ECM reformulation of the dynamic model as in Steen and Salvanes (1999).

The value of the H-statistic ranges from 0.214 (basic specification) to 0.605 (when time dummies are incorporated), and in all cases is statistically significant. The Wald test reveals that the H-statistic differs significantly from both zero and unity and, therefore, the hypotheses of both monopoly and perfect competition are rejected. Given the discussion in Sections 2 and 3, the dominant market form suggested by the static RT is monopolistic competition. Finally, we test for long-run equilibrium using the return on assets as the dependent variable. The Wald test performed does not reject the hypothesis of equilibrium (H_n = 0) at conventional statistical levels (x^2 (4) = 52.93, p-value = 0.000), which implies that our analysis is well specified.

5.4. Dynamic RT

Table 7 presents the results of the dynamic RT model (Eq. 10). In this specification, the significance of the input prices falls compared to the static RT (only the coefficient on w_2 is positive and statistically significant in the basic model and when nominal values are applied and the coefficient on w_3 when time dummies are included in the specification). The effect of TA remains positive and statistically significant, TA attaining however lower values compared with the static RT.

More importantly, we note that the value of the H-statistic is practically indistinguishable from zero in the cases of the basic model and its variant with nominal values, whereas in the models where time-related terms are added, either in the form of a time trend or in the form of time dummies, H departs from zero, attaining values around 0.16. The hypothesis tests H = 1 and H = 0 show that perfect competition is rejected for every variant model, and that some form of collusive behavior cannot be rejected, respectively (at the 5 per cent level of significance). The test for examining the existence of long-run equilibrium once again confirms the hypothesis of equilibrium at the 5 per cent level, even though the Wald test attains a lower value compared to the static case (x^2 (4) = 16.08, p-value = 0.003). [24]

The signs found are consistent with the view that the CPM and RT offer similar conclusions regarding the structure of the Greek banking sector. Yet, both dynamic models indicate at least some anticompetitive behavior of banks, while their static counterparts point towards competitive conduct. These results have noteworthy implications for researchers and policy makers, as they challenge the dominant arguments regarding the structure of the Greek banking sector. One could elicit further information if these models were compared to a variety of differently structured banking systems or if they were extended to account for the critiques of the NEIO literature (e.g. Corts, 1999). Yet, before we move on to another issue we had better bring this entry to a close.

6. Concluding remarks

Contrary to standard accounts, we have used both a CPM and a RT to assess competitive conditions in a specific banking industry. The analysis further distinguished between static and dynamic versions of these models in order to substantiate whether predictions regarding the market structure remained unchanged. We tested the four

[24] We also included the lags of all independent variables (i.e. lagged input prices and lagged *TA*) in the model. Since an upward-sloping supply curve for bank inputs will have the effect of driving up input prices as a function of contemporaneous quantities of bank outputs, whereas the revenue levels predicted by the PR model may respond to input prices only with a lag, this may imply that inclusion of lagged input prices could help overcome the mask of monopoly power (see Shaffer, 2004). Yet, not only the estimation results were similar, but also the annual character of the data probably suggests that the preferred model is given by Eq. (10).

resulting specifications, using panel data from the Greek banking industry over the period 1993-2004.

We contend that our results indicate that both static models tend to underestimate the level of market power. In particular, while the static CPM and RT indicate no anticompetitive conduct and monopolistic competition respectively, their dynamic counterparts signal some anticompetitive behavior of banks. This is especially true for the dynamic RT, and for the dynamic CPM when time dummy variables are included in its empirical specification. We may partially attribute the mask of market power by static models to the important dynamics that characterized the Greek banking sector during the examined period, which were not reflected in the empirical specifications of either the static CPM or the static RT. These results hold consistently across a number of econometric specifications and estimation methods, as applied separately to the static and dynamic models, enhancing some recent critiques regarding the suitability of static NEIO models to robustly estimate market power.

At a broader level of analysis the conclusions of the present article underline the crucial relevance of the special features of the examined banking industry and they highlight the need to develop more appropriate empirical methodologies to characterize the level of collusive behavior in banking.

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Table 1
An overview of Greek commercial banks

	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004
NCB	20	19	18	20	19	19	16	17	20	19	20	21
NE	38	40	40	43	44	46	47	53	52	54	54	52
NB	1,200	1,244	1,469	1,599	1,788	2,048	2,070	2,670	2,766	2,854	2,876	2,953
NIM	1.57	1.36	2.09	1.98	2.25	2.42	2.70	2.69	2.67	2.42	2.72	2.80
NoIM	2.18	2.85	2.15	2.22	2.21	1.92	3.74	2.16	1.50	0.91	0.96	0.90
OEA	2.35	2.51	2.73	2.87	2.82	2.57	2.68	2.58	2.44	2.27	2.31	2.30
LLP	0.34	0.38	0.23	0.52	0.62	0.52	0.64	0.37	0.33	0.39	0.49	0.60
ROA	1.06	1.31	1.26	0.79	0.99	1.20	3.04	1.86	1.39	0.66	0.87	0.70
LA	23.78	25.17	28.06	31.53	31.90	36.27	36.58	43.79	47.67	52.46	56.98	62.70
EA	4.55	4.87	4.84	4.47	5.10	5.98	9.89	8.94	9.28	6.61	6.84	6.70

This table reports some figures for Greek commercial banks over the period 1993-2004. NCB: number of Greek commercial banks; NE: number of employees (x1000); NB: number of branches; NIM: net interest margin i.e. net interest income to average assets (in percentage terms); NoIM: non-interest income to average assets (in percentage terms); OEA: operating expenses to average assets (in percentage terms); LLP: loan loss provisions to average assets (in percentage terms); ROA: return on assets (in percentage terms); LA: loan to assets (in percentage terms); EA: equity to assets (in percentage terms).

Source: Organization for Economic Cooperation and Development, OECD, Bank Profitability – Financial Statements of Banks: 1994 – 2003, OECD, 2004 and Bank of Greece, Annual Report 2005, BoG, 2006.

Table 2
The dataset

Bank Name	Years
National Bank of Greece	1993-2004
Alpha Bank	1993-2004
Eurobank	1993-2004
Agricultural Bank of Greece	1993-2004
Emporiki Bank (aka Commercial Bank of Greece)	1993-2004
Piraeus Bank	1993-2004
Geniki Bank (aka General Bank of Greece)	1993-2004
Egnatia Bank	1993-2004
Bank of Attica	1993-2004
Laiki Bank (formerly European Popular Bank)	1993-2004
Aspis Bank	1993-2004
NovaBank	2001-2004
Probank	2002-2004
Omega Bank	2001-2004
Marfin Bank (formerly Crédit Lyonnais, Piraeus Prime)	1993-2004
Panellinia Bank	2002-2004
Telesis Bank (formerly Dorian Bank)	1993-2000
Ergobank	1993-1999
Ionian Popular Bank	1993-1999
Bank of Macedonia-Thrace	1993-1999
Xiosbank	1993-1999
Cretabank	1993-1998
Bank of Central Greece	1993-1998
National Mortgage Bank	1993-1997
Bank of Athens	1993-1997
Interbank	1993-1996
National Housing Bank	1993-1996
Bank of Cyprus	1993-2004

This table reports the Greek commercial banks constituting the sample (only one foreign-owned bank is included, namely Bank of Cyrpus).

Source: Annual Balance Sheet and Income Statements of Greek Commercial Banks, 1993-2004.

Table 3
Banking indicators

Year	Q	P	Y	Z	C	w_1	w_2	w_3	TR
1993	2358.23	12.73	79771.30	22.75	295.84	19.312	12.634	5.671	317.06
1994	2316.54	13.37	76048.87	19.00	338.02	20.445	14.132	6.021	330.70
1995	2506.28	11.61	74386.21	15.60	330.27	22.304	15.475	5.771	298.25
1996	2545.69	10.77	74937.20	14.40	351.62	24.850	14.967	6.097	296.72
1997	2964.64	9.40	77622.13	9.90	379.47	28.460	11.982	6.622	316.09
1998	3351.22	9.39	75385.52	8.50	470.25	27.723	9.792	6.448	372.79
1999	4487.44	7.40	79429.47	6.30	572.75	30.253	5.390	7.179	495.64
2000	6504.77	7.38	80468.60	6.10	811.09	33.978	5.345	5.589	657.63
2001	6499.75	6.18	83255.77	5.30	620.04	36.878	4.220	5.571	493.34
2002	5840.03	4.69	86847.19	5.12	489.30	33.994	3.534	5.212	360.12
2003	6042.43	4.70	90822.54	4.27	458.04	35.376	3.001	5.220	352.57
2004	6318.65	4.64	95722.17	4.25	490.93	39.274	3.288	5.406	366.63
Average	4311.31	8.52	81224.75	10.12	467.30	29.404	8.647	5.901	388.13

Q represents total assets at the end of each year (in millions of euros); *P* is the ratio of annual interest income to total assets; *Y* is the gross domestic product (in billions of euros); *Z* is the 10-year Greek government bond yield (used as a substitute for bank deposits, in percentage units); *C* is total cost (in millions of euros); w_1 is the quotient of personnel expenses per employee (in thousand of euros); w_2 is the ratio of interest expense per total funds (in percentage units); w_3 is the ratio of administrative expenses per fixed assets (in percentage units); *TR* is total revenue (in millions of euros). Each row represents the average bank in our set of banks for a particular year, with the exception of attributes *Y* and *Z*, which correspond to macroeconomic data. All variables expressed in monetary units are in real terms.
Source: Annual Balance Sheet and Income Statements of Greek Commercial Banks, 1993-2004 and own estimations.

Table 4
Results (Static CPM)

	Basic model		Nominal values		Trend		Time dummies	
	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic
α_0	40566.1	16.70	32280.8	9.57	40524.5	13.91	40511.4	16.66
P	-10325.3	-14.01	-4983.8	-9.93	-10330.8	-8.88	-10303.5	-12.96
Y	-0.361	-13.28	-0.117	-5.17	-0.360	-10.65	-0.360	-13.14
Z	4293.1	9.42	-1231.1	-3.70	4290.3	4.85	4270.0	8.75
PY	0.122	13.33	0.033	8.95	0.122	8.10	0.122	12.00
PZ	47.477	10.08	147.497	8.57	48.231	6.15	47.849	9.61
YZ	-0.063	-11.16	-0.003	-1.36	-0.063	-5.79	-0.062	-10.41
R-sq		0.940		0.953		0.940		0.940
DW		1.947		1.793		1.950		1.947
L	0.012	1.32	0.001	0.42	0.006	0.81	0.107	0.83
β_0	199.88	3.01	129.04	3.49	116.38	3.98	97.52	2.07
lnq	-16.604	-1.48	-3.391	-0.75	-3.440	-0.77	3.649	0.66
$\ln q^{-2}$	1.354	1.71	0.396	1.26	0.339	1.05	-0.020	-0.05
$\ln w_1$	-38.509	-6.55	-14.274	-4.77	-14.453	-4.36	-4.328	-1.41
$\ln w_2$	3.008	2.72	4.717	4.58	2.948	2.71	2.708	2.17
$\ln w_3$	0.928	1.90	1.461	2.28	0.791	1.62	0.333	0.82
t					0.250	2.57		
t^2					-0.017	-2.26		
R-sq		0.901		0.906		0.898		0.893
DW		1.628		1.503		1.497		1.487
Obs		228		228		228		228

The table reports the results arising from the estimation of the static CPM (System 6). *q:* total assets at the end of each year (in millions of euros); *P:* the ratio of annual interest income to total assets; *Y:* the gross domestic product (in billions of euros); *Z:* the 10-year Greek government bond yield (used as a substitute for bank deposits, in percentage units); w_1: the ratio of personnel expenses per employee (in thousand of euros); w_2: the ratio of interest expense per total funds (in percentage units); w_3: the ratio of administrative expenses per fixed assets (in percentage units). Coefficient estimates, with corresponding *t*-statistics for four variants of the model; R-sq: the R-squared value of the equation; DW: Durbin-Watson statistic; Obs: number of observations.

Table 5
Results (Dynamic CPM)

	Basic model		Nominal values		Trend		Time dummies	
	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic
α_0	21872.6	11.17	17309.1	7.75	39442.2	18.49	20354.7	6.73
α_1	0.576	14.91	0.523	14.62	0.233	5.56	0.982	16.44
P	-9240.9	-16.73	-7626.7	-17.51	-10991.7	-23.35	-12820.2	-15.02
Y	-0.221	-11.06	-0.107	-8.09	-0.562	-17.56	-0.184	-5.96
Z	5750.8	13.90	4709.6	12.51	5135.9	15.11	8773.9	13.72
PY	0.119	17.44	0.068	19.98	0.141	24.21	0.167	15.81
PZ	23.904	5.31	-3.612	-0.30	22.317	6.09	29.818	4.28
YZ	-0.081	-14.75	-0.044	-16.24	-0.068	-14.73	-0.124	-14.57
R-sq		0.96		0.92		0.93		0.93
DW		2.26		2.20		2.19		2.05
L	0.008	1.65	0.012	1.92	0.004	1.68	0.143	2.70
β_0	307.19	7.66	135.32	3.54	245.12	5.64	240.10	5.51
Lnq	-18.387	-3.86	-3.702	-0.80	-5.224	-0.97	-7.551	-1.47
Lnq^{-2}	1.314	3.90	0.424	1.32	0.583	1.59	0.673	1.90
lnw_1	-20.374	-6.81	-20.696	-6.64	-14.268	-4.30	-14.498	-4.48
lnw_2	6.262	5.65	6.470	5.38	3.852	3.10	3.506	2.92
lnw_3	1.133	0.95	1.856	1.47	0.773	0.63	0.649	0.54
t					-0.834	-0.88		
t^2					-0.219	-0.65		
R-sq		0.931		0.908		0.918		0.930
DW		1.618		1.500		1.529		1.703
Sargan		0.315		0.223		0.329		0.416
AR (1)		0.092		0.007		0.083		0.102
AR (2)		0.179		0.123		0.214		0.233
Obs		202		202		202		202

The table reports the results arising from the estimation of the dynamic CPM (System 7 and 8). *q:* total assets at the end of each year (in millions of euros); *P:* the ratio of annual interest income to total assets; *Y:* the gross domestic product (in billions of euros); *Z:* the 10-year Greek government bond yield (used as a substitute for bank deposits, in percentage units); w_1: the ratio of personnel expenses per employee (in thousand of euros); w_2: the ratio of interest expense per total funds (in percentage units); w_3: the ratio of administrative expenses per fixed assets (in percentage units). Coefficient estimates, with corresponding *t*-statistics for four variants of the model; R-sq: the R-squared value of the equation; DW: Durbin-Watson statistic; Sargan: Sargan test for overidentifying restrictions (p-value); AR (1): test for first order serial correlation (p-value); AR (2): test for second order serial correlation (p-value); Obs: number of observations.

Table 6
Results (Static RT)

	Basic model		Nominal values		Trend		Time dummies	
	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic
β_0	-1.788	-2.67	-2.031	-2.82	-7.803	-8.70	-9.002	-9.15
$\ln w_1$	-0.076	-1.13	-0.006	-0.08	0.440	5.29	0.456	5.52
$\ln w_2$	0.262	12.32	0.255	12.50	0.103	3.76	0.095	3.37
$\ln w_3$	0.028	1.19	0.024	1.07	0.045	1.93	0.053	2.31
lnTA	0.861	37.54	0.852	43.05	1.095	29.04	1.098	28.93
D_{99}	0.197	3.92	0.203	4.21	0.109	2.12		
D_{00}	0.200	3.58	0.207	3.86	0.136	2.45		
t					0.627	3.83		
t^2					-0.397	-6.46		
H	0.214	2.74	0.274	3.58	0.588	6.85	0.605	7.05
H = 0	7.50	0.01	12.84	0.00	46.93	0.00	49.70	0.00
H = 1	101.30	0.00	89.99	0.00	23.09	0.00	21.17	0.00
FE	8.79	0.00	9.64	0.00	6.99	0.00	7.12	0.00
R-sq		0.970		0.971		0.978		0.978
Obs		228		228		228		228

The table reports the results arising from the estimation of the static RT (Equation 9). The dependent variable is the logarithm of total revenue scaled by total assets. w_1: the ratio of personnel expenses per employee (in thousand of euros); w_2: the ratio of interest expense per total funds (in percentage units); w_3: the ratio of administrative expenses per fixed assets (in percentage units); TA: total assets. Coefficient estimates, with corresponding *t*-statistics for four variants of the model. The H-statistic is equal to the sum of the elasticities of total revenue with respect to three input prices. The Wald test is used to test the H=0 and H=1 hypotheses and follows an F-distribution. R-sq: the R-squared value of the equation; FE: test for fixed effects; Obs: number of observations.

Table 7
Results (Dynamic RT)

	Basic model		Nominal values		Trend		Time dummies	
	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic
β_0	0.023	0.02	0.147	0.12	-2.574	-2.22	-2.916	-2.34
β_l	0.469	7.99	0.459	7.64	0.283	4.33	0.307	4.59
$\ln w_1$	-0.152	-1.61	-0.152	-1.57	0.078	0.79	0.069	0.70
$\ln w_2$	0.158	3.29	0.152	3.23	0.046	0.94	0.028	0.56
$\ln w_3$	0.005	0.15	0.001	0.03	0.044	1.45	0.060	2.00
lnTA	0.480	8.30	0.481	8.19	0.702	10.09	0.681	9.72
D_{99}	0.229	4.42	0.225	4.38	0.142	2.74		
D_{00}	0.232	4.37	0.226	4.30	0.180	3.69		
t					0.359	1.40		
t^2					-0.215	-2.72		
H	0.011	0.08	0.001	0.01	0.168	1.35	0.157	1.24
H = 0	0.010	0.94	0.000	0.99	1.830	0.18	1.550	0.21
H = 1	55.40	0.00	54.81	0.00	44.57	0.00	44.48	0.00
F-test	609.58	0.00	645.15	0.00	617.82	0.00	402.17	0.00
Sargan		0.307		0.326		0.343		0.514
AR (1)		0.023		0.017		0.004		0.014
AR (2)		0.743		0.699		0.475		0.423
Obs		202		202		202		202

The table reports the results arising from the estimation of the dynamic RT (Equation 10). The dependent variable is the logarithm of total revenue scaled by total assets. w_1: the ratio of personnel expenses per employee (in thousand of euros); w_2: the ratio of interest expense per total funds (in percentage units); w_3: the ratio of administrative expenses per fixed assets (in percentage units); TA: total assets. Coefficient estimates, with corresponding *t*-statistics for four variants of the model. The H-statistic is equal to the sum of the elasticities of total revenue with respect to three input prices. The Wald test is used to test the H=0 and H=1 hypotheses and follows an F-distribution. R-sq: the R-squared value of the equation; Sargan: Sargan test for overidentifying restrictions (p-value); AR (1): test for first order serial correlation (p-value); AR (2): test for second order serial correlation (p-value); FE: test for fixed effects; Obs: number of observations.