Applied Financial Economics, 1998, **8**, 191–200

Market structure and performance in Spanish banking using a direct measure of efficiency

JOAQUÍN MAUDOS

Universidad de Valencia and Instituto Valenciano de Investigaciones Económicas (IVIE); Departamento de Análisis Económico, Edificio departamental oriental; Avda. de los Naranjos, s/n; 46022 Valencia, Spain

This paper analyses the relationship between market structure and performance within the Spanish banking industry. Three different stochastic measures of efficiency are used (based on three alternative distributional assumptions for inefficiency: half-normal, normal-truncated and exponential). The results obtained support the 'modified efficient structure' hypothesis since, even though efficiency is the main determinant of profitability, market power (as reflected in a market share variable), also affects profitability. The results obtained also show that market share is an inadequate proxy for efficiency.

I. INTRODUCTION

The relationship between performance and market structure has generated two competing hypotheses. On one hand, the traditional collusion hypothesis proposes that market concentration lowers the cost of collusion between firms and results in higher than normal profits. On the other hand, the efficient structure hypothesis postulates that the most efficient firms obtain greater profitability and market share and, as a consequence, the market becomes more concentrated.

Traditionally, various studies have tested these two alternative hypotheses using market share as a proxy for efficiency. These studies (Smirlock *et al*., 1984, 1986; Smirlock, 1985; Evanoff and Fortier, 1988; Molyneux *et al*., 1994; Molyneux and Forbes, 1995, for example), argue that the most efficient firms have lower costs and will consequently gain market share. Therefore, market share can be used as a proxy for efficiency.

Most recently, some authors (Shepherd, 1986; Timme and Yang, 1991; Berger, 1995) have questioned the use of market share as a proxy for efficiency in testing the efficient structure hypothesis versus the structure-conduct-performance paradigm. This is due to the fact that the market share variable may capture the effect of other variables rather than efficiency.

However, in spite of the criticisms of using market share as a proxy for efficiency, recent papers on Spanish banking (Molyneux *et al*., 1994) continue to use this approximation to test the efficient structure hypothesis against the traditional collusion hypothesis.

This paper analyses the relationship between profitability and market structure (concentration and/or market share) in the Spanish banking industry applying for the first time direct measures of productive efficiency. Using the stochastic frontier approach, a frontier cost function is estimated to obtain a direct measure of efficiency of Spanish banks. The main contribution of this paper is that it analyses the sensitivity of the results of testing the efficient structure hypothesis versus the collusion hypothesis using three alternative distributional assumptions for inefficiency: half-normal, normal-truncated and the exponential model.

The results obtained show that market share is an inadequate proxy for efficiency taking into account that the R^2 between the two variables is under 1%. The 'modified efficient structure hypothesis' is shown to be useful because even though efficiency is the main determinant of profitability, market power, reflected by the residual influence of market share, also positively affects profitability. These results contradict those recently obtained by Molyneux *et al.* (1994), due mainly to the fact that these authors use market share as a proxy for efficiency, not a direct efficiency measure as is used here.

The structure of the paper is as follows: Section II analyses the alternative hypotheses that explain the relation between performance and market structure; Section III describes the methodology used to obtain the efficiency measures; Section IV describes the variables used as well as their construction; and Section V presents the empirical results. Finally, Section VI contains the conclusions.

II. THE RELATION BETWEEN MARKET STRUCTURE AND PERFORMANCE

Studies of the relationship between performance and market structure have been divided between two alternative hypotheses. On one hand, the collusion hypothesis, also called the structure-conduct-performance hypothesis (Bain, 1951), postulates that greater benefits are the result of the concentration of the market since this facilitates the collusion between the firms of the industry. On the other hand, the efficient structure hypothesis (Demsetz, 1973, 1974; Peltzman, 1977) proposes an alternative explanation for the existing positive correlation between concentration and profitability, affirming that the most efficient firms obtain greater profitability and market share and, as a consequence, the market becomes more concentrated. In this case, the positive observed relationship between concentration and profits is spurious and simply proxies for the relationship between superior efficiency, market share, and concentration.

The studies directed to test these hypotheses are based on the estimate of the following model (Smirlock, 1985; Evanoff and Fortier, 1988; Molyneux *et al.*, 1994; Molyneux and Forbes, 1995, etc)

$$\pi = \beta_0 + \beta_1 CR + \beta_2 MS + \alpha' X \tag{1}$$

where π is a measure of a firm's performance (ROA, ROE, Tobin's q, etc), MS is the market share of the firm, CR is a measure of the concentration of the market, and α is a vector of additional control variables specific to the firm and the market that prior studies have found to affect bank profitability. In this context, Smirlock (1985) shows that if β_1 is statistically greater than zero and β_2 is zero, the collusion hypothesis holds, while if β_1 is zero and β_2 is statistically greater than zero the efficient structure hypothesis prevails.

The implicit assumption in testing the efficient structure hypothesis versus the collusion hypothesis is that market share is a proxy variable for efficiency. Under this assumption, the most efficient firms gain market share at the expense of the less efficient. However, as pointed by Shepherd (1986), the market share variable can capture the effect of unrelated variables to efficiency.

The studies based on the model shown in Equation 1 sometimes obtain similar results although they interpret them in a very different way. Some studies posit that a positive sign in the case of market share and null effect in the case of concentration shows the existence of market power, because market share is only the reflection of market power (Shepherd, 1986). Elsewhere, other authors attribute this same result as support of the efficient structure hypothesis in the sense that market share is a proxy variable for efficiency (Smirlock *et al.*, 1984, 1986; Smirlock, 1985; Evanoff and Fortier, 1988; Molyneux *et al.*, 1994; Molyneux and Forbes, 1995). However, and as pointed by Berger (1995), these last papers do not show a direct efficiency measure.

To test the efficient structure hypothesis versus the collusion hypothesis, we will estimate the following equation

$$\pi = \beta_0 + \beta_1 CR + \beta_2 MS + \beta_3 EF + \alpha' X \tag{2}$$

EF being a direct efficiency measure obtained after estimating a stochastic cost frontier.

Based on the estimation of Equation 2, the different explanatory hypotheses of the performance can be summarized as (Timme and Yang, 1991)

$$\frac{\partial \pi}{\partial CR} > 0; \quad \frac{\partial \pi}{\partial MS} = 0; \quad \frac{\partial \pi}{\partial EF} = 0 \tag{3}$$

$$\frac{\partial \pi}{\partial CR} = 0; \quad \frac{\partial \pi}{\partial MS} = 0; \quad \frac{\partial \pi}{\partial EF} > 0 \tag{4}$$

$$\frac{\partial \pi}{\partial CR} = 0; \quad \frac{\partial \pi}{\partial MS} > 0; \quad \frac{\partial \pi}{\partial EF} > 0 \tag{5}$$

$$\frac{\partial \pi}{\partial CR} > 0; \quad \frac{\partial \pi}{\partial MS} = 0; \quad \frac{\partial \pi}{\partial EF} > 0 \tag{6}$$

where Equations 3 and 4 represent the pure collusion hypothesis and efficient structure hypothesis, respectively, while Equations 5 and 6 represent the modified efficient structure hypothesis and the hybrid collusion/efficiency hypothesis, respectively.

The modified efficient structure hypothesis (Shepherd, 1986) establishes that the variance in performance is explained by efficiency as well as by the residual influence of the market share, because market share captures the influence of factors unrelated to the efficiency, such as the power of market and/or the product differentiation.[1] As in the pure efficient structure hypothesis, the modified efficient structure hypothesis postulates that market concentration does not directly affect business performance.

The hybrid efficient structure/collusion hypothesis (Schmalensee, 1987) establishes that concentration affects profitability as a result of market power. Also, this hypothesis affirms that the most efficient firms are more

[1] Banks with a large market share may have higher quality products, enabling them to charge higher prices and earn higher profits.

profitable, with the residual effect of market share being held as negligible.

III. THE MEASUREMENTS OF EFFICIENCY

Frontier functions can be estimated statistically or not according to whether we adopt certain assumptions related to the stochastic properties of the data. Furthermore, we can distinguish between a parametric and non-parametric approach depending on whether or not a specific functional form between the variables is assumed (data envelopment analysis (DEA) is the non-parametric approach more frequently used).

Another way to classify frontier functions distinguishes between a deterministic and a stochastic approach. In the first case, it is assume that all deviations from the frontier are due to inefficient behaviour while in the second case deviations can be due to inefficiency as well as to circumstances not under the control of the firm (random fluctuations).[2]

The main advantage of DEA is that is not necessary to make distributional assumptions to estimate efficiency. However, one disadvantage is the general assumption that the distance that separates the observed observation from the frontier is due exclusively to inefficiency (there is no random fluctuation), therefore estimates of inefficiency can be upwardly biased.

The stochastic frontier approach was introduced simultaneously by Aigner *et al.* (1977) and Meeusen and van den Broeck (1977). This approach modifies the standard production function (or costs) by assuming that inefficiency forms part of the error term. It also posits that the compound error term includes inefficiency as well as a purely random component that captures the effect of variables not under the control of the firm (economic climate, bad luck, etc).

Thus, the basic stochastic cost frontier assumes that the observed costs of a firm differ from the cost frontier as a consequence of random fluctuations (v_i) and inefficiency (u_i). That is, in the case of the costs frontier

$$\ln C_i = \ln C(Y_i, P_i) + \varepsilon_i; \qquad \varepsilon_i = u_i + v_i \quad i = 1, \dots, N \tag{7}$$

where C_i are the observed costs of the firm i, Y_i is the output vector, P_i is the vector of input prices, and $\ln C_i(Y_i, P_i)$ is the logarithm of the predicted costs of a firm that minimizes the costs of production. The random error term v_i is assumed independent and identically distributed, and inefficiency term u_i is assumed independently distributed of v_i.

To separate the effect of both components, it is necessary to specify a distributional assumption for both components of the error term. Since inefficiency can only increase costs above the frontier, it is necessary to specify asymmetric distributions for the inefficiency term. Commonly, it is assumed that v_i is drawn from a normal distribution with mean zero and variance σ_v^2, and u_i from a half-normal distribution (u_i is the absolute value of a variable that is distributed as a normal with mean zero and variance σ_u^2).

Under the assumption that both components of the composed error term are distributed independently, the frontier function can be estimated by maximum likelihood, with inefficiency derived from the residuals of the regression. Individual inefficiency estimates can be calculated by using the distribution of the inefficiency term conditional on the estimate of the composed error term. Thus, Jondrow *et al.* (1982) shows that in the case of the half-normal distribution, the mean of this conditional distribution adopts the following expression

$$E[u_i|\varepsilon_i] = \frac{\sigma\lambda}{(1+\lambda^2)}\left[\frac{\phi(\varepsilon_i\lambda/\sigma)}{\Phi(-\varepsilon_i\lambda/\sigma)} - \frac{\varepsilon_i\lambda}{\sigma}\right] \tag{8}$$

where $\lambda = \sigma_u/\sigma_v$, $\sigma^2 = \sigma_u^2 + \sigma_v^2$, ϕ and Φ are the standard normal distribution and the standard normal density function, respectively.

As noted above, the half-normal distribution assumes that inefficiency is distributed according to a normal distribution truncated with zero mean. This restrictive assumption has been criticized by Stevenson (1980) who proposes as an alternative specification the truncated normal distribution with the mean (μ) different from zero ($|N(\mu, \sigma_u^2)|$). In this case, individual inefficiencies are calculated as in Equation 8 substituting the term $[\varepsilon_i\lambda/\sigma]$ for

$$\mu_i^* = \frac{\varepsilon_i\lambda}{\sigma} + \frac{\mu}{\sigma\lambda} \tag{9}$$

since in this case the mean of the distribution (μ) is different from zero.

Finally, assuming that inefficiency is drawn from an exponential distribution, individual inefficiency at the firm level can be estimated according to the following expression (Greene, 1993)

$$E[u_i/\varepsilon_i] = (\varepsilon_i - \theta\sigma_v^2) + \frac{\sigma_v\phi[(\varepsilon_i - \theta\sigma_v^2)/\sigma_v]}{\Phi[(\varepsilon_i - \theta\sigma_v^2)/\sigma_v]} \tag{10}$$

There has been a substantial amount of work calculating X-inefficiencies in banking markets,[3] but only three published papers exist that use direct efficiency measures to test the efficient structure hypothesis versus the collusion hypothesis. Berger (1995) estimates the efficiency measures using the distribution-free approach (Berger, 1993), which assumes that the differences of efficiency between firms are stable over time while random error tends to average out. The advantage of this approach in measuring efficiency is

[2] A review of the different approaches to the efficiency measurement can be found in Bauer (1990), Greene (1993) and Lovell (1993).
[3] See Berger *et al.* (1993) and Berger and Humphrey (1997) for a review of studies on the efficiency of financial institutions.

that it does not impose arbitrary assumptions on the distribution of efficiency.[4]

Timme and Yang (1991) use the stochastic frontier approach to obtain individual efficiency measures assuming a half-normal distribution. Goldberg and Rai (1996) also applied a stochastic cost frontier for European banking under the assumption that the errors are distributed half-normal.

We have preferred the stochastic frontier approach as compared to the distribution free-approach and to DEA. Even though the first approach has the advantage that it is not necessary to assume distributional assumptions for the inefficiency term (as the standard fixed and random effects models), it has the disadvantage that it assumes that inefficiency is constant over time. Concerning the DEA, the disadvantage is that this method generally assumes that all deviations from the frontier are due to inefficiencies. Thus, inefficiency could be upwardly biased.

Obviously, the different approaches used can affect inefficiency measurement, which in turn affects the evaluation of the efficient structure versus collusion hypothesis. For this reason, we have opted to use the stochastic frontier approach, although we analyse the robustness of the results using three different distributional assumptions for estimating the efficiency scores.

IV. VARIABLES USED

To measure the efficiency of Spanish banks, we assume a translog frontier cost function as a consequence of its greater flexibility in relation to other specifications. The translog function is a quadratic function obtained by a Taylor series expansion in logarithms around the point of approximation. Among the principal advantages we note the following: (1) no restriction is imposed *a priori* on the substitution elasticity between inputs; (2) the cost function can be U-shaped; and (3) potential complementarities in cost through multiproduct specification can be permitted as well.

In our case, the translog cost function adopts the following specification

$$\begin{aligned}\ln TC_{it} = \alpha_0 &+ \sum_{k=1}^{2} \alpha_k \ln Y_{kit} + 1/2 \sum_{k=1}^{2}\sum_{j=1}^{2} \alpha_{kj} \ln Y_{kit} \ln Y_{jit} \\ &+ \sum_{k=1}^{3} \beta_k \ln P_{kit} + 1/2 \sum_{k=1}^{3}\sum_{j=1}^{3} \beta_{kj} \ln P_{kit} \ln P_{jit} \\ &+ \sum_{k=1}^{2}\sum_{j=1}^{3} \lambda_{kj} \ln Y_{kit} \ln P_{jit} + \sum_{t} \sigma_t DT_t + \varepsilon_{it} \qquad (11)\end{aligned}$$

where TC_{it} = total costs (operating plus financial) of the firm i in the year t, Y_{1it} = total deposits in real terms of the firm i in the year t, Y_{2it} = total loans in real terms of the firm i in the year t, P_{1it} = price of the labour input of the firm i in year t calculated as the ratio of labour expenses to the average number of employees, P_{2it} = price of the deposits[5] of the firm i in the year t calculated as the ratio of financial expenses to the average deposits, and P_{3it} = price of the physical capital of the firm i in the year t calculated as the ratio of total capital expenses to fixed assets, and

$$\varepsilon_{it} = u_{it} + v_{it} \qquad (12)$$

these two elements being inefficiency and the random term, respectively.[6] Also, time dummies (DT) are introduced to capture the influence of technical progress.[7]

Some banks were dropped from the sample for two reasons: (1) as a consequence of the lack of information in some of the necessary variables to estimate the cost function; (2) because of questions about the reliability of the reported information especially after mergers. For this reason, the final used sample is made up of 353 observations over the period 1990–93.[8] The sample size varies from 72 in 1990 to 94, 94 and 93 in 1991, 1992 and 1993, respectively. However,

[4] Another reason for using the distribution-free approach in panel estimation of stochastic cost frontiers is that it averages out cyclical/luck factors. The standard approaches differ mainly in the distributional assumptions used to disentangle X-efficiency differences from random errors that temporarily give decision-making units high or low costs. Most of these methods were designed for application to a single period of data, where random fluctuations in costs owing to luck and measurement error, as well as changes in regulation and macroeconomic conditions, can make inferences about underlying efficiency differences across firms difficult to divine. This is why Berger uses the distribution free approach.

[5] Following the valued added approach of Berger and Humphrey (1992), we consider deposits as an input (since input costs are affected by changes in interest paid on deposits) and output (since the production of deposit services account for the majority of capital and labour expenses) simultaneously.

[6] The use of total costs (operating + financial) and the output metric used is consistent with Berger *et al.*'s (1987) approach. The authors show that when outputs are defined in terms of the value of loans and/or deposits, the modelled costs should include both operating and interest expenses. The problem with this approach is that, if market power exists, the effect of a lower remuneration of the deposits (less interest expenses) can be shown as efficiency. However, auxiliary regressions do not show any relation between average financial cost and market share. See Timme and Yang (1991) for a more detailed discussion of this issue.

[7] We impose the usual symmetry and homogeneity constraints.

[8] In the period under analysis, the Spanish bank sector has seen many mergers. Appropriate sample selection becomes a concern. In this paper we have preferred to work with an unbalanced panel. Therefore, when two banks merge, they singly disappear and a new entity is shown. This strategy allows us to use the maximum available information, unlike in the two alternative strategies we discuss. To deal with mergers, the authors of previous papers have either completely eliminated the merged banks or have, in effect, gone back in time to create new fictitious banks.

Table 1. *Summary statistics (1990–93)*

	Mean	Standard deviation
ROA	0.0161	0.0220
ROE	0.2278	0.2790
Concentration (CR)	0.0962	0.0204
Market share (MS)	0.0417	0.0525
ASSETS*	450 450	1 213 769
Loans/assets (LOASS)	0.4358	0.2042
Growth in market deposits (GMD)	0.0943	0.0587
Market deposits (MAKDEP)*	2352	2042
Inefficiency (Half-normal)	0.2178	0.1759
Inefficiency (Normal-truncated)	0.2029	0.3168
Inefficiency (Exponential)	0.2010	0.3631

*Millions of pesetas.

the sample does contain 99% of all bank assets, so the missing banks are very small.

Performance measures used (see Table 1) are return of assets (ROA) and return on equity (ROE), as proxies for gross profits. We used these measures because they represent the benefits obtained by the banks before taxes, provision for insolvency and extraordinary items, and reflect the difference between earnings and costs derived from lending and from bank services. We have chosen to specify profits this way since net profit after taxes captures the effects of random factors that are sometimes beyond the firm's control (provision for insolvency, for example).[9]

It is also important to define carefully what we mean by the firm's market. In this paper, competition among banks takes place at a regional level because, in fact, many banks only operate in one province of Spain.

One problem associated with defining the Spanish banks' market as a regional one is that no information currently exists concerning the regional distribution of the representative variables of banking output (deposits, loans). Only regional branch distribution data are available. We assume that the regional distribution of the deposits of a bank is proportional to the number of branches.[10] Therefore, deposit market shares and concentration levels are calculated using regional branch distribution data which proxies for deposit distribution. We use a Henfindahl index of branches to determine concentration.

The control variables that we used to estimate Equations 1 and 2 are firm and market specific variables. More precisely, firm variables include the size of each bank (ASSETS) to show the influence of factors related to the size of production (for example, economies of scale), and the ratio loans/assets (LOASS) to show the risk assumed by banks.[11] We assume the latter to be positive because loans are riskier than other primary assets.

Market specific variables include the size of the deposit market (MAKDEP), and market growth (GMD). In the first case, we assume a negative sign for this variable since the largest markets tend be markets where there is more competition; easier market entry, and greater awareness among customers for bank services. Relative to market growth, we assume a positive sign since expanding markets can generate higher profits. We weighted the relative importance of each regional market in terms of the provincial distribution of the branches of each bank. The size of each provincial market is approximated by the value of deposits since this is the only available information at the province level.

V. EMPIRICAL RESULTS

Table 2 shows the results of the estimation of Equation 2 using ROA as the dependent variable. We also show the results of progressively introducing the variables *CR*, *MS* and *EF*. Thus, the results of the first row (1) are directly comparable with previous studies of the collusion hypothesis (control variable plus *CR*). In our specification, we reject the collusion hypothesis since the *CR* variable is not statistically significant.

In row (2), market share is used as representative of the market structure, and has a statistically significant positive effect on profitability.

The simultaneous introduction of *CR* and *MS* as explanatory variables of ROA (row (3)), shows how *MS* has a positive effect on profitability. These results are consistent with those obtained in other studies (Smirlock, 1985; Smirlock *et al.*, 1984; Evanoff and Fortier, 1988) in which it is shown that when concentration and market share are introduced simultaneously in the regression, market share has a positive effect, while the effect of the concentration is not significant. Also, these results have often been interpreted as support of the efficient structure hypothesis as a result of using the market share as a proxy for efficiency.[12]

Rows (4) to (6) show the results of additionally introducing a direct measure of efficiency. Irrespective of the assumed distributional assumption for the inefficiency term, the results show that efficiency is highly significant and positive, adding substantial explanatory power in the

[9] In Section V we check the robustness of the results using net profits instead of gross profits.
[10] What this assumption implies is that for a bank *i* the ratio of deposits per branch is equal in every province where it operates. The ratio varies for individual banks.
[11] Since the profit measure is not risk-adjusted, the loan-to-asset ratio is included to account for differing risk levels between banks.
[12] See for example Smirlock (1985) and Evanoff and Fortier (1988).

Table 2. *Collusion versus efficient structure hypothesis, 1990–93 (353 observations); dependent variable: ROA (gross profits)*

	Constant	CR	MS	EF	ASSETS	LOASS	GMD	MAKDEP	R^2
(1)	0.0214	0.0388			− 0.646E − 09	− 0.510E − 03	0.0179	− 0.436E − 05	0.17
	(3.100)	(0.643)			(− 0.806)	(− 0.010)	(0.907)	(− 6.838)	
(2)	0.0216		0.0898		− 0.191E − 08	− 0.407E − 02	0.0173	− 0.349E − 05	0.19
	(5.571)		(3.321)		(− 2.177)	(− 0.716)	(0.944)	(− 5.376)	
(3)	0.0213	0.0030	0.0895		− 0.191E − 08	− 0.0040	0.0170	− 0.348E − 05	0.19
	(3.130)	(0.050)	(3.252)		(− 2.173)	(− 0.713)	(0.873)	(− 5.097)	
(4a)	− 0.0108	− 0.0118	0.0788	0.0462	− 0.215E − 08	− 0.0068	0.0185	− 0.384E − 05	0.33
	(− 1.474)	(− 0.214)	(3.215)	(8.259)	(− 2.663)	(− 1.310)	(1.041)	(− 6.123)	
(4b)	0.0114	− 0.0130	0.0855	0.0164	− 0.21E − 08	− 0.0074	0.0174	− 0.347E − 05	0.24
	(1.653)	(− 0.223)	(3.209)	(4.975)	(− 2.536)	(− 1.306)	(0.923)	(− 5.254)	
(4c)	0.0144	− 0.0101	0.0863	0.0121	− 0.215 − 08	− 0.0069	0.0169	− 0.343E − 05	0.23
	(2.091)	(− 0.172)	(3.204)	(4.143)	(− 2.492)	(− 1.245)	(0.865)	(− 5.134)	

(4a) Half-normal model
(4b) Truncated-normal model
(4c) Exponential model
t-statistics in parentheses.

Table 3. *Collusion versus efficient structure hypothesis, 1990–93 (353 observations); dependent variable: ROE (gross profits)*

	Constant	CR	MS	EF	ASSETS	LOASS	GMD	MAKDEP	R^2
(1)	0.0573	1.7720			− 0.140E − 07	0.1352	0.5591	− 0.504E − 04	0.23
	(0.684)	(2.420)			(− 0.806)	(1.970)	(2.331)	(− 6.531)	
(2)	0.1483		2.0097		− 0.134E − 07	0.0551	0.6465	− 0.351E − 04	0.30
	(5.571)		(6.324)		(− 1.304)	(0.826)	(2.992)	(− 4.597)	
(3)	0.0555	1.0017	1.9260		− 0.132E − 07	0.0589	0.5396	− 0.316E − 04	0.31
	(0.694)	(1.410)	(5.972)		(− 1.284)	(0.884)	(2.359)	(− 3.953)	
(4a)	− 0.2210	0.8739	1.8351	0.3963	− 0.152E − 07	0.0351	0.5529	− 0.347E − 04	0.36
	(− 2.449)	(1.286)	(5.940)	(5.789)	(− 1.545)	(0.550)	(2.528)	(− 4.522)	
(4b)	− 0.0371	0.8520	0.8890	0.1533	− 0.155E − 07	0.0306	0.5344	− 0.311E − 04	0.33
	(− 0.543)	(1.222)	(5.971)	(3.394)	(− 1.524)	(0.463)	(2.372)	(− 3.947)	
(4c)	0.0144	− 0.0101	0.0863	0.0121	− 0.215 − 08	− 0.0069	0.0169	− 0.343E − 05	0.33
	(2.091)	(− 0.172)	(3.204)	(4.143)	(− 2.492)	(− 1.245)	(0.865)	(− 5.134)	

(4a) Half-normal model
(4b) Truncated-normal model
(4c) Exponential model
t-statistics in parentheses.

regression (the R^2 of the regression increases to 33%). Nevertheless, the explanatory power is greater in the half-normal (the R^2 of the regression raises to 73%) which may indicate that this distributional assumption is more adequate according to the data used.

Of the control variables, only size (ASSETS) and market size (MAKDEP) are statistically significant. In the case of size, its negative influence shows the effect of diseconomies of scale, while the negative effect of MAKDEP may be due to the fact that competition is greater in large markets.[13]

Using ROE as the dependent variable gives similar results (Table 3) with the only difference that the variable *CR* is significantly greater than zero when neither *MS* nor efficiency is introduced in the regression, although it is not significant once *MS* is included in the regression.

As pointed out by Berger (1995), the fact that the parameter that accompanies the market share variable is statistically significant and the coefficient is not altered when the effect of the efficiency is introduced in the estimation suggests that in the prior regressions in which efficiency was not

[13] As a consequence of the high explanatory power of the market size variable (MAKDEP) according to its t-ratio, and the possible negative correlation between concentration and market size, we have rerun the regressions and eliminated this variable. In this case, the influence of *CR* is statistically significant only when we do not take into account the influence of *MS* and/or *EF*.

included, market share cannot be interpreted as a proxy variable for efficiency. In other words, since the effect of efficiency is controlled in the regression, the positive effect of the market share indicates the existence of market power. Consequently, these results allow us to accept what is called the modified efficient structure hypothesis.

The results also are very similar if we estimate Equation 2 for the yearly data, 1990 through to 1993 (Table 4). In all the regressions, market share and efficiency have a positive and statistically significant coefficient allowing us once again to accept the modified efficient structure hypothesis.[14]

Table 5 shows the results using net profits (after-tax, loan loss reserve provisions and other extraordinary items) instead of gross profits. In this case, and as expected, the proportion of profits explained by the regressors is lower since net profits are sometimes influenced by volatile changes such as loan–loss provisioning. The results also show how efficiency affects profitability positively, the effect of market share and concentration being insignificant.

As pointed out by Berger (1995), one of the implications of the pure efficient structure hypothesis is that efficiency should be positively related to market share and/or concentration. For this reason, Table 6 shows how, when market share and concentration are regressed against efficiency and the control variables, efficiency is positively correlated with both, although not in a statistically significant way.[15,16] This result reinforces the acceptance of the modified efficient structure hypothesis, since market share captures the effect of variables unrelated to efficiency. Also, the weak correlation between market share and efficiency (R^2 below 1%), shows that it is inadequate to use the former as a proxy for the latter, as has been used in other studies.[17]

The results are contrary to those obtained by Molyneux *et al.* (1994) who test the collusion hypothesis versus the efficient structure hypothesis in the Spanish banking system over the period 1986–89 through the estimation of Equation 1 using market share as a proxy for efficiency. There are several possible reasons that may explain our results.

First, the Spanish banking sector has seen much deregulation since the middle and the end of the 1980s: branching restrictions for private banks were removed in 1985; interest rate ceilings disappeared in 1987; investment coefficients that froze a very significant share of total assets in regulated loans and public debt were gradually eliminated, and the ban on branch expansion for savings banks beyond regional markets was lifted in 1989.

Specifically, the previous stronger regulations and reduced pressure from external competition were more convenient for the establishment of collusive agreements among banks. Now, however, the greater pressure for competition as a consequence of the European Union, as well as the almost complete deregulation of the Spanish banking system at the beginning of the 1990s, are less likely to yield positive results for the collusion hypothesis.[18]

A second reason that can justify the different results obtained is the narrower definition of geographical area. Thus, while in Molyneux *et al.* (1994) market share, concentration, market size, etc. assume a national market, we have considered that the competition takes place at the regional level due to the fact that many banks only have branches in one province.[19]

To check the robustness of the results against different levels of disaggregation, Table 7 shows the results obtained when all variables are checked against the national market.[20] Once again, the results indicate that efficiency is the more significant variable in the regression, allowing us to accept the modified efficient structure hypothesis.[21]

Finally, the representative variable of business performance used in Molyneux *et al.* (1994) is the net income/assets ratio (ROA). However, such a profitability measure can be affected by a randomness component because it incorporates the effect of more discretionary items like the provision for insolvency and other extraordinary items.

Obviously, the rejection of the structure-conduct-performance paradigm, does not support the defence of measures taken to prevent the growth of market concentration (mergers, absorptions, etc), since greater market concentration does not imply reductions in competition and/or in efficiency, nor monopoly profits. Nevertheless, the acceptance of the modified efficient structure hypothesis, because it recognizes the influence of market power in addition to efficiency, implies that the measures directed to increase bank size, may have an ambiguous effect on social benefit,

[14] Only in 1992 is the market share not significant, leading in that case to the acceptance of the pure efficient structure hypothesis.
[15] These results are consistent with those reported by Berger (1995).
[16] The efficiency measure used in Table 5 corresponds to the half-normal model, being that the results are similar in the truncated-normal and exponential models.
[17] In addition, the results in Table 6 indicate that an increase in cost efficiency of 100 basis points would be associated with a 1.3 basis points increase in market share, suggesting this result is a relatively weak economic linkage between efficiency and market share.
[18] Vives (1991) has suggested that the main consequence of the deregulatory process leading to European monetary integration 'will be to change the focal point of the strategies of banks from collusion to competition.
[19] In the case of market concentration, the different geographical market chosen can affect the results obtained because, in Molyneux *et al.* (1994), this variable has a constant value for all banks in each year.
[20] Concentration, market share, market size and market growth are computed on the basis of the deposits.
[21] The results shown in Table 6 correspond to the half-normal model for the inefficiency term. The results in the truncated-normal and exponential models are very similar. If we use net profits instead of gross profits, R^2 of the regressions are lower, being statistically significant only in the effect of efficiency.

Table 4. *Collusion versus efficient structure hypothesis by years (half-normal model)*

Year	Constant	CR	MS	EF	ASSETS	LOASS	GMD	MAKDEP	R^2	Obs
1990	− 0.723E − 04	− 0.1550	0.0782	0.0550	− 0.337E − 08	− 0.773E − 03	− 0.0027	− 0.418E − 05	0.25	72
ROA	(− 0.003)	(− 1.132)	(1.656)	(2.507)	(− 1.799)	(− 0.063)	(− 0.081)	(− 2.031)		
1990	− 0.2983	− 1.3983	1.8302	0.9558	− 0.578E − 07	− 0.0501	0.7188	− 0.652E − 04	0.23	72
ROE	(− 0.568	(− 0.506)	(1.9181)	(2.155)	(− 1.527)	(− 0.202)	(1.052)	(− 1.567)		
1991	0.0132	− 0.1295	0.0969	0.0373	− 0.272E − 08	− 0.872E − 02	0.053	− 0.576E − 05	0.32	94
ROA	(0.684)	(− 0.914)	(1.916)	(3.168)	(− 1.565)	(− 0.926)	(0.908)	(− 2.419)		
1991	− 0.4296	2.5974	2.1930	0.3973	− 0.388E − 07	0.1148	0.0057	− 0.210E − 04	0.49	94
ROE	(− 1.964)	(1.620)	(3.832)	(2.979)	(− 1.972)	(1.355)	(0.0091)	(− 0.781)		
1992	0.0012	− 0.0429	0.0515	0.0554	− 0.182E − 08	− 0.0257	− 0.0210	− 0.617E − 05	0.35	94
ROA	(0.066)	(− 0.312)	(0.867)	(3.919)	(− 0.993)	(− 2.290)	(− 0.546)	(− 4.609)		
1992	− 0.3937	2.1414	1.5591	0.4613	− 0.449E − 08	− 0.0176	0.1112	− 0.291E − 04	0.47	94
ROE	(− 0.317)	(1.769)	(3.063)	(3.7081)	(− 0.279)	(− 0.182)	(0.328)	(− 0.727)		
1993	− 0.0047	− 0.2452	0.1396	0.0437	− 0.168E − 08	0.0070	− 0.1020	− 0.161E − 0·6	0.38	93
ROA	(− 0.240)	(− 1.043)	(2.561)	(5.626)	(− 1.328)	(0.718)	(− 1.410)	(− 0.116)		
1993	− 0.0680	− 0.6044	2.0808	0.2789	− 0.123E − 08	0.0250	− 0.0786	− 0.108E − 04	0.37	93
ROE	(− 0.317)	(− 0.239)	(3.553)	(3.340)	(− 0.091)	(0.238)	(− 0.101)	(− 0.727)		

t-values in parentheses.

Table 5. *Collusion versus efficient structure hypothesis, 1990–93 (353 observations); net profits*

	Constant	CR	MS	EF	ASSETS	LOASS	GMD	MAKDEP	R^2
ROA	0.0419	− 0.1076			− 0.127E − 08	− 0.0249	0.0388	− 0.402E − 05	0.05
	(3.867)	(− 1.138)			(− 1.019)	(− 2.784)	(1.252)	(− 4.028)	
	0.0308		0.0203		− 0.168E − 08	− 0.0254	0.0257	− 0.331E − 05	0.05
	(4.978)		(0.473)		(− 1.199)	(− 2.812)	(0.878)	(− 3.201)	
	0.0419	− 0.1197	0.0302		− 0.170E − 08	− 0.0258	0.0384	− 0.372E − 05	0.06
	(3.862)	(− 1.244)	(0.692)		(− 1.219)	(− 2.863)	(1.241)	(− 3.430)	
	0.0191	− 0.1302	0.0226	0.0326	− 0.187E − 08	− 0.0278	0.0395	− 0.397E − 05	0.08
	(1.518)	(− 1.373)	(0.526)	(3.412)	(− 1.357)	(− 3.121)	(1.296)	(− 3.708)	
ROE	0.6804	− 2.4330			− 0.345E − 08	− 0.3630	0.8896	− 0.653E − 04	0.015
	(2.118)	(− 0.868)			(− 0.093)	(− 1.384)	(0.969)	(− 2.210)	
	0.4251		0.5650		− 0.140E − 07	− 0.3843	0.5880	− 0.481E − 04	0.014
	(2.320)		(1.274)		(− 0.339)	(− 1.436)	(0.678)	(− 1.572)	
	0.6797	− 2.7499	0.7911		− 0.146E − 07	− 0.3940	0.8816	− 0.3576 − 04	0.017
	(2.114)	(− 0.964)	(0.610)		(− 0.354)	(− 1.474)	(0.959)	(− 1.791)	
	0.3243	− 2.9140	0.6731	0.5092	− 0.172E − 07	− 0.4255	0.8987	− 0.615E − 04	0.025
	(0.858)	(− 1.024)	(0.520)	(1.775)	(− 0.417)	(− 1.590)	(0.981)	(− 1.914)	

t-values in parentheses.

Table 6. *Efficiency-market share/concentration relationship (half-normal model)*

	Constant	EF	MAKDEP	ASSETS	LOASS	GMD	R^2
Dep var = MS	0.0243	0.0223					0.005
	(1.902)	(1.399)					
	0.0292	0.0131	− 0.115E − 04	0.144E − 07	0.0384	0.0544	0.44
	(2.556)	(1.086)	(− 10.209)	(9.249)	(3.468)	(1.498)	
Dep var = CR	0.0927	0.0044					0.0014
	(18.579)	(0.710)					
	0.0926	0.0043	− 0.441E − 05	0.942E − 09	− 0.900E − 03	0.1110	0.24
	(17.578)	(0.792)	(− 8.567)	(1.325)	(− 0.179)	(6.733)	

t-statistics in parentheses.

Table 7. *Collusion versus efficient structure hypothesis at national level, 1990–93 (353 observations); gross profits*

	Constant	CR	MS	EF	ASSETS	LOASS	GMD	MAKDEP	R^2
Dep. var	0.1236	− 0.0067			0.760E − 10	0.0135	− 0.1451	− 0.424E − 04	0.07
ROA	(0.924)	(− 0.363)			(0.091)	(2.423)	(− 0.410)	(− 0.967)	
	0.1354	− 0.1028	0.4343		− 0.789E − 08	0.0120	− 0.1863	− 0.4469E − 05	0.08
	(1.014)	(− 0.549)	(1.817)		(− 1.769)	(2.151)	(− 0.527)	(− 1.020)	
	0.1056	− 0.1483	0.4534	0.0439	− 0.870E − 08	0.0097	− 0.2126	− 0.437E − 05	0.20
	(0.847)	(0.175)	(2.032)	(7.235)	(− 2.065)	(1.848)	(− 0.645)	(− 1.069)	
Dep. var	− 0.4169	− 1.8199			0.0237E − 07	0.2959	3.4173	0.133E − 04	0.13
	(− 0.255)	(− 0.796)			(2.319)	(4.340)	(0.790)	(0.249)	
ROE	− 0.2828	− 2.2172	4.9397		− 0.679E − 07	0.2795	2.9483	0.108E − 04	0.14
	(− 0.173)	(− 0.968)	(1.690)		(− 1.231)	(4.069)	(0.682)	(0.203)	
	− 0.5419	− 2.6129	5.1057	0.3821	− 0.742E − 07	0.2588	2.7191	0.117E − 04	0.20
	(− 0.343)	(− 1.178)	(1.805)	(4.966)	(− 1.389)	(3.887)	(0.650)	(0.226)	

t-statistics in parentheses.

since mergers can lead to more efficient banks but with greater market power.

VI. CONCLUSIONS

This paper has tested the efficient structure hypothesis versus the collusion hypothesis in the Spanish banking industry. We use, for the first time, a direct measure of efficiency obtained through the estimate of a stochastic cost frontier. The study also determines the sensitivity of the results using three different procedures for measuring efficiency.

The results obtained for Spanish banks over the period 1990–93 allow us to accept the so-called 'modified efficient structure hypothesis' since efficiency positively affects profitability, although market power, reflected in market share, does so as well. Also, because market concentration is shown to be insignificant in bank performance, we reject the traditional collusion hypothesis.

These findings suggest that bank regulatory decisions based on concerns for their impact on changes in concentration may be inappropriate and should focus instead on bank efficiency. Thus, and according to the results obtained in this paper, the recent mergers encouraged by the government and the Bank of Spain might be justified on efficiency grounds.

Our results are contrary to those of Molyneux *et al.* (1994), where the structure-conduct-performance paradigm was accepted. Although there are several possible reasons that may explain our different results (different period of analysis, different performance measure, different geographical market measure, etc), the main reason appears to be the fact that Molyneux *et al.* (1994) use market share as a proxy for efficiency while we use a direct measure (not proxy). Two other studies that have used a direct measure of efficiency in testing these hypotheses (Timme and Yang, 1991; Berger, 1995) find results similar to our own: they reject the traditional collusion hypothesis and find that efficiency is a more important determinant of profitability than is either market concentration or market share.

ACKNOWLEDGEMENTS

This paper was written while the author was a Visiting Researcher at the Finance Department of Florida State University. The comments provided by David B. Humphrey and an anonymous referee are gratefully acknowledged. The author would also like to thank the financial support of the Fundación Caja de Madrid and the DGICYT PB94-1523. A preliminary version of this article was published as Working Paper in the EC Series (WP-EC 96-12) of the *Instituto Valenciano de Investigaciones Económicas* (IVIE).

REFERENCES

Aigner, A., Lovell, C. A. K. and Schmidt, P. (1977) Formulation and estimation of stochastics frontier production function models, *Journal of Econometrics*, **86**, 21–37.

Bain, J. S. (1951) Relation of profit rate of industry concentration, *Quarterly Journal of Economics*, **65**, 293–324.

Bauer, P. W. (1990) Recent developments in the econometric estimation of frontiers, *Journal of Econometrics*, **46**, 39–56.

Berger, A. N. (1993) Distribution-free estimates of efficiency in the U.S. banking industry and test of the standard distribution assumptions, *Journal of Productivity Analysis*, **4**, 261–92.

Berger, A. N. (1995) The profit-relationship in banking – tests of market-power and efficient-structure hypotheses, *Journal of Money, Credit and Banking*, **27** (2), 405–31.

Berger, A. N. and Humphrey, D. B. (1992) Measurement and efficiency issues in commercial banking. In *Output Measurement in the Service Sectors*, Zvi Griliches (ed) Chicago. National Bureau of Economic Research, University of Chicago Press, Chicago, pp 245–79.

Berger, A. N. and Humphrey, D. B. (1997) Efficiency of financial institutions: international survey and directions for future research, *European Journal of Operational Research*, **98** (2), 175–212.

Berger, A. N., Hanweck, G. A. and Humphrey, D. B. (1987) Competitive viability in banking. Scale, scope and product mix economies, *Journal of Monetary Economics*, **20**, 501–20.

Berger, A. N., Hunter, W. C. and Timme, S. G. (1993) The efficiency of financial institutions, *Journal of Banking and Finance*, **17**, 219–49.

Demsetz, H. (1973) Industry structure, market rivalry and public policy, *Journal of Law and Economics*, **16**, 1–9.

Demsetz, H. (1974) Two systems of belief about monopoly. In *Industrial Competition: The New Learning*, H. Goldschmid, H. M. Mann and J. F. Weston (eds), 164–84. Little, Brown, and Company, Boston.

Evanoff, D. D. and Fortier, D. L. (1988) Reevaluation of the structure-conduct-performance paradigm in banking, *Journal of Financial Services Research*, **1**, 277–94.

Goldberg, L. G. and Rai, A. (1996) The structure-performance relationship for European banking, *Journal of Banking and Finance*, **20**, 745–71.

Greene, W. M. (1993) The econometric approach to efficiency-analysis. In H. O. Fried, C. A. K. Lovell, and S. S. Schmidt, (eds), *The Measurement of Productive Efficiency: Techniques and Applications*, 68–119, Oxford University Press, Oxford.

Jondrow, J., Lovell, C. A. K. Materov I. S. and Schmidt, P. (1982) On the estimation of technical inefficiency in the stochastics frontier production models, *Journal of Econometrics*, **19**, 233–38.

Lovell, C. A. K. (1993) Production frontiers and productive efficiency. In: H. O. Fried, C. A. K. Lovell and S. S. Schmidt (eds), *The Measurement of Productive Efficiency: Techniques and Applications*, 3–67, Oxford University Press, Oxford.

Meeusen, W. and van den Broeck, J. (1977) Efficiency estimation from Cobb–Douglas production function with composed error, *International Economic Review*, **18**, 435–44.

Molyneux, P. and Forbes, W. (1995) Market structure and performance in European Banking, *Applied Economics*, **27**, 155–59.

Molyneux, P., Lloyd-Willimas, D. M. and Thornton, J. (1994) Market structure and performance in Spanish banking, *Journal of Banking and Finance*, **18**, 433–44.

Peltzman, S. (1977) The gains and losses from industrial concentration, *Journal of Law and Economics*, **20**, 229–63.

Schmalensee, R. (1987) Collusion versus differential efficiency: testing alternatives hypotheses, *The Journal of Industrial Economics*, **35**, 399–425.

Shepherd, W. G. (1986) Tobin's *q* and the structure performance relationship: reply, *American Economic Review*, **76**, 1205–10.

Smirlock, M. (1985) Evidence on the (non)relationship between concentration and profitability in banking, *Journal of Money, Credit and Banking*, **17**, 69–83.

Smirlock, M., Gilligan, T. and Marshall, W. (1984) Tobin's *q* and the structure–performance relationship, *American Economic Review*, **74**, 1050–60.

Smirlock, M., Gilligan, T. and Marshall, W. (1986) Tobin's *q* and the structure–performance relationship: reply, *American Economic Review*, **76**, 1211–13.

Stevenson, R. (1980) Likelihood functions for generalized stochastic frontier estimation, *Journal of Econometrics*, **13**, 58–66.

Timme, S. G. and Yang, W. K. (1991) On the use of a direct measure of efficiency in testing structure–performance relationships, Working Paper, Georgia State University.

Vives, X. (1991) Regulatory reform in European banking, *European Economic Review*, **35**, 505–15.