

External Liberalization of Banking and Industrial Concentration: The Evidence from Spain

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Abstract

This paper investigates the effect of international liberalization on industrial bank concentration in Spain. By employing time series modeling techniques, it is shown that, albeit a weak short run effect may exist, there has been no long run effect on industrial bank concentration from external liberalization. The data base for the study consists of monthly time series for the period 1974-1982.²

Introduction

In March 1979 the Spanish government, with the issuance of Ministerial Order 16101, embarked on a policy of international liberalization of its banking sector. It is of empirical interest to policymakers to examine the effect this action has on the domestic banking sector in terms of industrial concentration as well as if it is dampening this sector's industrial concentration ratio. This research seeks to provide answers to these issues.

In particular, this research investigates the causal *prima facie* temporal effect external liberalization has on the domestic bank concentration ratio. The null hypothesis is that there is no effect, short and/or long term, on domestic industrial bank concentration from external liberalization. If the null hypothesis is rejected, a corollary inquiry is if external liberalization has reduced the domestic banking sector's concentration ratio. One would expect *a priori* that external liberalization would increase the competitive structure of Spanish banking by, among other things, numerically increasing banking institutions as well as foster greater changes among Spanish banks so as to enable them to compete with their technologically advanced foreign brethren.

The empirical methods employed to test the hypothesis are from the applied econometric time series literature which differs from the obfuscating structural model approach. The empirical investigation is for the historical period 1974-1982. Two industrial concentration ratios are employed: a domestic bank concentration ratio, defined as the ratio of total assets of national banks to the total assets

of all banks; and a foreign bank concentration ratio, defined as the ratio of the total assets of all foreign banks to the total assets of all banks. The time series modeling techniques employed will ascertain if there is a short run relationship between these two parameters as well as relate the presence and direction of their long run association.

This study proceeds with a review of the literature in section two, then the data, empirical model, and testing procedure are described in section three, section four presents the empirical findings, and section five summarizes the conclusions of this research. The paper closes in section six by presenting some suggestions for further research.

Review of the Literature

There has been a great deal of research on the relationship between market structure and foreign investment. The focus of this avenue of research concerns the types of market structures that provide a climate for the growth of multinational firms in developed economies. In contrast, the question as to how multinational firms influence a particular market structure as a result of *de novo* entry is not extensively discussed in the literature. Steuer (1973) and Rosenbluth (1970), however, address this matter for the United Kingdom and Canada, respectively. Similarly, Newfarmer (1978) and Newfarmer and Mueller (1975) and Lall (1979) address this issue for Brazil and Malaysia. A significant shortcoming of this body of empirical literature is essentially its focus on the manufacturing sector.

There has been, in essence, very little research emphasis on the banking sector.

With regard to the competitive characteristics of banking markets, the research primarily concerns how interest rates for loans are influenced by deposit concentration ratios and the number of firms in the market. Davis and Verbrugge (1978), Heggsted (1977), and Kaufman (1966) are studies that represent this line of research. Marlow (1983) studied increased market entry into the commercial banking industry and its influence on the performance of the industry. Clark and Speaker (1992), Smirlock (1985), and Smirlock and Brown (1986) analyze the relationship between concentration and profitability in banking. Rivera (1993a) examines this relationship for Spanish banking to find that net interest margins declined as a result of a multinational bank presence.

However, these research findings do not address the issue of how the entry of multinational banks into a previously internally-closed banking market affects industrial concentration. Rivera (1985, 1993b) addresses this particular issue by analyzing the influence of multinational bank presence on Spanish financial structure. He finds that a multinational presence serves to decrease banking concentration. Cho (1990) examines the presence of foreign banks on Indonesian banking structure, he finds that the presence of foreign banks did not necessarily reduce competition in the domestic banking market. These studies, however, employ annual data and the statistical technique of ordinary least squares within contemporaneous model specification.

Vives (1990, p. 409) examines deregulation and competition in Spanish banking and notes that “[F]oreign banks have an edge in the wholesale business, international operations in particular, and also do well in the higher income segment....” and that foreign banks accommodate themselves to traditional Spanish banking practices while maintaining a high degree in innovation. This work is purely descriptive and does not delve into the issue of the effect of foreign banking on the domestic bank concentration ratio. It does point out, however, that the internal liberalization, i.e., deregulation, of the Spanish banking system was well under way prior to its external liberalization.

The research presented in this paper differs from other studies in that it employs several time series techniques to determine the dynamic association between domestic bank concentration and the liberalization of the domestic banking market to foreign banks with special emphasis on Spain. Indeed contemporaneously specified regression models do not always lead to an appropriate picture of the true relationship between parameters, albeit results may be

statistically significant. This research, by employing non-contemporaneous equation specifications, seeks to overcome this shortcoming and hence to find the dynamic association, if any, of this relationship.

Data Description, Empirical Model, and Testing Procedure

This section discusses the data and the empirical model used to investigate the relationships under study. The two ratios constructed from the data are: the domestic bank concentration ratio (DBCR), defined as total assets of national banks to the total assets of all banks, and the foreign bank concentration ratio (FBCR), defined as total foreign bank assets to total assets of all banks.³ The empirical model investigates the dynamic association between these two banking concentration ratios by sequentially combining several applied econometric time series procedures.

The Data

The historical monthly data are for the period 1974-1982 and are aggregated from two sources: *Anuario Estadístico de la Banca Privada* and *Balances y Estadísticas de la Banca*; both tomes are published by Spain’s Consejo Superior Bancario (Higher Banking Board). These sources provide complete balance sheet information for all banks operating in Spain. The data sample is limited to the period 1974-1982 given the limitation of data availability and, more importantly, to changes in the recording of assets and liabilities after 1982. These changes result in the non-compatibility of post-1982 data with prior years. The available data, however, does provide an asymptotically robust sample size of 108 observations.

Table 1 concisely presents the structure of Spanish banking for the historical period under review. In general, as part A of Table 1 indicates, banks fall into one of five categories: national banks, which encompass the nation’s seven largest banks, regional banks, industrial banks, local banks, and foreign banks. For the period under review the number of national banks declined from eighteen in 1974 to thirteen in 1982. In contrast, during this same period, the number of foreign banks increased from only four in 1974 to thirty in 1982. It is clear that the major increase in foreign bank presence occurs after 1978, i.e., subsequent to the official external opening of the Spanish banking system.

Part B of Table 1 shows that during this period the domestic bank concentration ratio declined from sixty-nine percent to sixty-four percent and the foreign bank concentration ratio increased from two percent to eight percent. This pattern suggests that the decline in one concentration ratio was offset by an increase in the other ratio. Indeed,

as part C of the table shows, the simple coefficient of correlation between these variables is roughly a statistically significant negative eighty percent.⁴ It is well documented in the econometric literature, however, that contemporaneous correlations may not reflect the appropriate relationship between parameters. This is particularly true when

there is a theoretical basis to expect that a set of variables interact intertemporally. In such a scenario it is appropriate to apply noncontemporaneous econometric modeling procedures to investigate the hypothesis under study. These dynamic econometric methods, as applied to the issues of this research, are presented below.

Table 1

Structure of Spanish Banking: 1974-1982

A. Institutional Structure

Year	Bank Structure				
	National	Regional	Industrial	Local	Foreign
1974	18	13	18	55	4
1975	18	12	22	55	4
1976	18	12	22	56	4
1977	16	12	23	55	4
1978	14	12	24	55	4
1979	14	12	26	53	14
1980	14	12	26	51	22
1981	13	12	25	49	28
1982	13	12	25	49	30

B. Domestic and Foreign Bank Concentration Ratios

Year	DBCR	FBCR
1974	.686	.017
1975	.684	.015
1976	.652	.022
1977	.653	.019
1978	.654	.019
1979	.638	.028
1980	.632	.052
1981	.622	.077
1982	.638	.077

C. Simple Coefficients of Correlations

	DBCR	LDBCR
FBCR	-.787*	
LFBCR		-.808*

*Statistically significant at the one percent level, $T = 108$, and the t-statistics are -10.902 and -14.132 , respectively. For parts A and B of the table, numbers are year-end figures. DBCR and FBCR are the domestic and foreign bank concentration ratios, respectively; LDBCR and LFBCR are their natural logarithms.

The Empirical Model

The sequential modeling procedures employed in this research are given by equations (1)-(5):

$$\Delta DBCR_t = \alpha + \sum_{j=1}^T \beta_j \Delta DBCR_{t-j} + \sum_{j=1}^T \gamma_j DV_{t-j} + \delta LT + \epsilon_t. \tag{1}$$

$$\begin{aligned} \Delta DBCR_t = \alpha^* + \sum_{j=1}^T \beta_j^* \Delta DBCR_{t-j} + \sum_{j=1}^T \psi_j^* \Delta FBCR_{t-j} \\ + \sum_{j=1}^T \gamma_j^* DV_{t-j} + \delta^* LT + \epsilon_t^*. \end{aligned} \tag{2}$$

$$FPE(r, s) = \left[\frac{T+r+s+1}{T-r-s-1} \right] \left[\frac{\sum_{t=1}^T (\Delta DBCR_t - \Delta \hat{D}BCR_t)}{T} \right]. \tag{3}$$

$$DBC R_t = \alpha_0 + \beta_0 FBC R_t + v_t. \tag{4}$$

$$\Delta \hat{v}_t = \rho \hat{v}_t + \omega_t. \tag{5}$$

Where: $DBC R_t$ is the domestic banking concentration ratio, defined as the ratio of total assets of national banks to the total assets of all banks; $FBC R_t$ is the foreign banking concentration ratio, defined as the ratio of the total assets of all foreign banks to the total assets of all banks; Dv_t is the government policy dummy variable, it takes the value of zero prior to March 1979 and thereafter has a value of unity indicating the international liberalization of domestic banking in Spain; LT is the time trend variable; and Δ is the first difference operator. The innovations in equations (1) and (2) are two independently distributed white noise uncorrelated error terms. All regressions are estimated using the Regression Analysis of Time Series (RATS) software.

Equation (3) is the Akaike final prediction (FPE) error statistic, its parameters are: T is the sample size; r is the lag structure of $DBC R_t$, the Granger-caused variable; s is the lag structure of $FBC R_t$, the hypothesized causal variable; $DBC R_t$ and $\hat{D}BC R_t$ with a “^” are the actual and predicted values, respectively, of the Granger-caused parameter. The Akaike statistic balances a model’s estimation error with its average modeling error -- it is an overall fit statistic; its value is independent of the existence of multicollinearity among an equation’s independent variables.

Finally, it is important to note, that for all relevant equations $j = 1, 2, \dots, 36$, representing the extensive lag specification search of one to thirty-six months.

The empirical modeling procedure uses an econometric definition of *prima facie* intertemporal association or econometric causality between variables suggested by Granger (1969) concomitant with Akaike’s (1969) final prediction error statistic. Granger (1988) suggests, as a necessary first step, for two parameters to achieve a long run equilibrium there must exist a dynamic short run causal ordering of the parameters.⁵ Equations (1)-(3) describes this short run relationship. The long run association between the parameters is investigated, as suggested by Engle and Granger (1987), with a cointegration model. Equation (4) is this cointegration model between the parameters under study. The coefficient for the $FBC R_t$ parameter, referred to in the literature as the cointegrating coefficient, relates the long run relationship between the parameters; provided, of course, that the residuals from the estimated cointegration equation are stationary. Equation (5) represents the test of the residuals of the cointegration equation for stationarity.

The short run relationship is investigated in terms of first differences of natural logarithms for the domestic and foreign bank concentration ratios so as to obtain covariance stationarity, this is necessary for testing for econometric causality. For the existence of a dynamic association between the variables must be independent of time. Indeed, Augmented Dickey-Fuller (1981) tests for the null hypothesis of a unit root in the data result in the acceptance of the null hypothesis for the log levels of the data, but its rejection for the first difference of the log levels of these variables.⁶ Alternatively stated, the log levels of both the domestic and foreign bank concentration ratios are integrated of order one, $I(1)$.

Testing Procedure

The empirical investigation to test the hypothesis proceeds as follows: equation (1) is estimated to obtain the optimal univariate model for the domestic banking concentration ratio. In essence, the equation is estimated with distributed lags of one to thirty-six months. The equation with the minimum Akaike statistic is selected. Given this results, the foreign bank concentration ratio is introduced into the equation (this gives equation two) which is estimated to obtain the optimal bivariate model. This equation is also estimated with lags of one to thirty-six months. It is clear the distributed lag space must be extensively searched, a thirty-six month lag space should be sufficient to ascertain the presence of a dynamic association between the variables.

The investigation into the short run effect on the domestic banking concentration ratio from international liberalization requires a comparison of the Akaike statistics of the optimal univariate and bivariate models. On the one hand, if the Akaike statistic of the former model is less than that of the latter model the null hypothesis is accepted. There is no short run relationship, i.e., Granger-econometric causality, from the foreign bank concentration ratio to the domestic bank concentration ratio. On the other hand, if the optimal bivariate model has a lower Akaike statistic, then the domestic bank concentration ratio is optimally modeled by including the informational content of both its own past values as well as the past values of the foreign bank concentration ratio. The null hypothesis is rejected, $FBCR_t$ affects $DBCR_t$ in the short run. Hence, a cumbersome seventy-two equation specification route must be taken to validate the existence of a short-run intertemporal association between the parameters under study.

The testing of the long run relationship -- which would indicate if external liberalization has a dampening impact on the bank concentration ratio -- can only proceed if and only if a short run relationship exists. If this is so, the cointegration model, i.e., equation (4), is estimated, and its residuals retrieved so as to estimate equation (5); if the t-statistic for the regression coefficient of the lagged error term is less than its non-standard t-critical value then the null hypothesis of no cointegration is rejected. In short, the cointegrating beta coefficient of equation (4) would relate the long run equilibrium relationship between the foreign and domestic bank concentration ratios. A negative sign for the cointegrating coefficient would indicate a dampening effect of $FBCR_t$ on $DBCR_t$.

Economic Empirical Findings

Table 2 summarizes the econometric findings for the search for the optimal univariate and bivariate models. In part A of this table the regression results for the optimal univariate equation are presented. Indeed, the domestic bank concentration ratio, the ratio of national banks' assets to the total assets of all banks, is optimally modeled with two monthly lags and has a minimum FPE statistic of .3194E-04. All other equation specifications have Akaike statistics that exceed this value. The optimal univariate model is estimated with 105 observations, 100 degrees of freedom, and has an $R^2 = .07$ and an adjusted R^2 . The F-statistic, a test of the null hypothesis that the one and two-monthly- $DBCR$ -distributed lags of the optimal model are jointly zero, is rejected at the seven percent level of significance.

Turning to part B of Table 2, the final result for the bivariate model search is given: the domestic bank concentration ratio is optimally modeled with an own-one to two monthly lags and a foreign bank concentration ratio of one monthly lag. The optimal bivariate model is estimated with 105 observations, 99 degrees of freedom, and has an $R^2 = .09$ and an adjusted $R^2 = .04$. The F-statistic, a test of the null hypothesis that the coefficients of the lagged foreign bank concentration ratio is zero, is rejected at the eleven percent level of significance. The Akaike statistic for this equation is $FPE(2, 1) = .3171E-04$.

Comparing the Akaike statistics for the optimal univariate and bivariate models, one concludes that $FPE(r = 2, s = 1) < FPE(r = 2)$. The foreign bank concentration ratio does have a *prima facie* Granger-causal effect on the domestic bank concentration ratio. Econometrically stated, there is a short run relationship between these two concentration ratios.

Albeit the proportion of the total variation of the domestic bank concentration ratio explained by the two regressions is low, these statistical results are not considered unusual. For the econometric techniques employed in this research formulate a test of the dynamic association of the variables and does not test for a structural model relationship. Moreover, the transformation of the data to covariance stationarity requires the extraction of the time dependent informational content of the variables. All these procedures reduce the magnitude of a model's R^2 . Similarly, the statistical insignificance of the regression coefficients is not surprising, for distributed lags of a variable are correlated; indeed, the F-statistics for the univariate and bivariate models supports this notion. For these reasons it is the magnitude of the Akaike statistic, an overall fit statistic, that is instrumental in determining the optimal model and thus the presence of a short run relationship.

Given the econometric support for the existence of a short run association between the two variables under study, the long run effect of external liberalization on the domestic bank concentration ratio can now be investigated. Table 3 presents these findings. Part A of this table indicates that the cointegrating coefficient is $-.4$, i.e., a one percent increase in the foreign bank concentration ratio will result in roughly half a percentage point decrease in the domestic bank concentration ratio.⁷ However, part B of this same table reports that the residuals of the cointegration model are not stationary. This implies that the null hypothesis of no cointegration between the variables must be accepted. There is no long run association between the foreign and domestic bank concentration ratios.⁸ The policy implication is that the change in the Spanish government's stance towards an external opening of the

Table 2

**Optimal Univariate and Bivariate Causal
Short Run Domestic Bank Concentration Ratio Models:
Final Results of Seventy-Two Equation Lag Space Search**

A. Optimal Univariate Equation: 1974.4-1982.12

$$\begin{aligned} \Delta DBCR_t = & - .0023 + .1646 \Delta DBCR_{t-1} - .1868 \Delta DBCR_{t-2} \\ & (-1.631) \quad (1.659^{**}) \quad (-1.875^{**}) \\ & - .0021 DV_t + .47E-04 LT. \\ & (-.965) \quad (1.288) \end{aligned}$$

$$R^2 = .07, \bar{R}^2 = .03, T = 105, d.f. = 100, SSR = .003, \\ DW = 2.00, F(2, 100) = 2.74^a, FPE(2) = .3194E-04.$$

B. Optimal Bivariate Equation: 1974.4-1982.12

$$\begin{aligned} \Delta DBCR_t = & - .0022 + .1019 \Delta DBCR_{t-1} - .2466 \Delta DBCR_{t-2} \\ & (-1.545) \quad (.964) \quad (-2.338^*) \\ & - .0248 \Delta FBCR_{t-1} - .0012 DV_t + .43E-04 LT. \\ & (-1.620) \quad (-.532) \quad (1.183) \end{aligned}$$

$$R^2 = .09, \bar{R}^2 = .04, T = 105, d.f. = 99, SSR = .003, \\ DW = 1.97, F(1, 99) = 2.62^b, FPE(2, 1) = .3171E-04.$$

*Significant at the one percent level. **Significant at the ten percent level. Values in parenthesis are t-statistics, all tests are two-tail tests. DBCR_t and FBCR_t are the domestic and foreign bank concentration ratios. ^aSignificant at the seven percent level. ^bSignificant at the eleven percent level. See text for the appropriate interpretation of the econometric results and for definitions of variables.

banking sector did not have any long term relevance in affecting the domestic bank concentration ratio. Perhaps, the statistical insignificance of the government policy dummy variable -- albeit it has the appropriate sign, in both the optimal univariate and bivariate models, was ominous of this finding.

Conclusions

This paper presents the results of applied econometric investigations into the intertemporal association between the domestic and foreign bank concentration ratios in light

of the Spanish government's decision to externally open this sector in March 1979. The null hypothesis being that this event had no effect on Spanish industrial banking concentration. This hypothesis is rejected in the short run, but accepted for the long run. This may reflect the evolving strength and momentum this sector developed from the government's earlier, pre-external-opening, measures of internally liberalizing banking.

The main empirical result of this research differs from the cited findings of other research in which a foreign multinational bank presence reduces industrial concentra

Table 3

**The Long Run Association Between External Liberalization'
and the Domestic Bank Concentration Ratio:
Cointegration Model Results**

A. The Cointegration Equation: 1974.1-1982.12

$$DBCR_t = \alpha_0 + \beta FBCR_t + v$$

$$DBCR_t = - .5879 - .0436 FBCR_t$$

(-52.285) (-14.132*)

$$R^2 = .65, \quad DW = .078, \quad T = 108, \quad d.f. = 106, \quad SSR = .038.$$

B. Test of the Null Hypothesis of No Cointegration

$$\Delta \hat{v}_t = \rho + \hat{v}_{t-1}$$

$$\Delta \hat{v}_t = -.0418 \hat{v}_{t-1}$$

(-1.541)

$$R^2 = .02, \quad DW = 1.869, \quad T = 107, \quad d.f. = 106, \quad SSR = .003.$$

*Significant at the one percent level. Values in parenthesis are t-statistics. DBCR_t and FBCR_t are the domestic and foreign bank concentration ratios. See text for the appropriate interpretation of the econometric results and for definitions of variables. It must be emphasized that since both variables are I(1), the OLS estimates, while super consistent, result in a lack of precision for the t-statistic. With regard to part B of the table, the non-standard t-critical value is less than the estimated t-statistic; hence, the null hypothesis of no cointegration is accepted, see the Appendix to Charemza and Deadman (1992).

tion. This body of literature, however, employs contemporaneously specified models. It is intuitively clear, therefore, that these models may have only captured a short run dampening effect while ignoring the intertemporal, non-contemporaneous, long run effect; which in this paper is found to be insignificant.

In summary, the international liberalization of Spanish banking did not have a statistically significance long run dampening affect on domestic bank concentration, defined as the ratio of total assets of national banks to the total assets of all banks. Ministerial Order 16101 did not inherently enhance the long run competitiveness of the banking

sector in Spain, given that a decrease in the domestic bank concentration ratio is viewed as augmenting the long run competitiveness of the banking system.

Suggestions for Future Research

The results reported in this research can be extended in several directions. In particular, although the *prima facie* nominal magnitude of the Spanish domestic bank concentration ratio has been affected by an increased international presence, it is difficult to conclude that Spanish banking has become more competitive. It follows that an investigation into the behavior of net interest margins

could shed more light on this issue. Similarly, if specific data for individual banking firms are made available in the future, then the findings of this paper can be compared with one using the Herfindahl Hirschman Index as a concentration ratio. In addition, a cross-sectional empirical study for a set of countries that have externally liberalized their banking systems would indicate the general empirical robustness of this paper's findings. □

*** Notes ***

1. This research was completed while on sabbatical leave as a Visiting Scholar at Columbia University during academic year 1993-1994.
2. Earlier drafts of this study were presented at the 1993 annual conferences of the Eastern Economic Association and the Northeast Business & Economics Association. The authors are grateful for the comments of the session participants.
3. These are commonly used simple bank concentration ratios. Ideally, one would like to use the Herfindahl Hirschman Index (HHI). In the present context HHI is defined as the sum of squares of market shares of all firms in the Spanish banking industry, this gives greater weights to larger banking firms relative to smaller ones. However, this concentration index cannot be calculated given the lack of firm specific data.
4. The structural changes in these bank concentration ratios do not necessarily imply an improvement in banking efficiency. Gibson and Tsakalotos (1993) discuss several reasons for such a scenario.
5. The term causality is used in its econometric sense, i.e., short run forecastability. For example, if $FBCR_t$ Granger-causes $DBCR_t$, then forecasts of $DBCR_t$ would have a smaller forecast error if the model includes past values of both parameters than only the past values of $DBCR_t$.
6. The ADF unit root tests for nonstationarity is given by the following equation:

$$\Delta z_t = \alpha + \beta t + \gamma z_{t-1} + \sum_{j=1}^k \Theta (z_{t-j} - z_{t-j-1}) + \eta_t$$

The null hypothesis is $(\alpha, \beta, \gamma) = (\alpha, 0, 0)$. The calculated F-statistics are: for the log levels, 2.21 for DBCR and 1.53 for FBCR; for the log first difference in levels, 30.46 for DBCR and 12.23 for FBCR, both are statistically significant at the one percent level. The relevant critical F-values are non-standard and are from Dickey and Fuller (1981), Table VI.

7. Although the DW-statistic indicates serial correlation in the cointegration model, no correction is needed. For such a correction would result in inconsistent estimates, see Engle and Granger (1987).

8. Hence, there is no need to construct an error correction model that would embody into one equation the short and long run relationships between the parameters.

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