ISSN 1518-3548

BANCO CENTRAL DO BRASIL

Working Paper Series 12

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March, 2001

ISSN 1518-3548
CGC 00.038.166/0001-05

Working Papers Series	Brasília	n. 12	Mar	2001	P. 1 – 24

Working Paper Series

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General Control of Subscription:

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SBS – Quadra 3 – Bloco B – Edifício-Sede – 2º subsolo
70074-900 - Brasília (DF)
Telefone (61) 414-1392
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Number printed: 450 copies

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A Test of Competition in Brazilian Banking

Márcio I. Nakane[1]

Abstract

This paper implements an empirical test of market power for Brazilian banking based on Bresnahan (1982) and Lau (1982). A dynamic version of the test is applied. The results show that the banking industry in Brazil is highly competitive, although the perfect competition hypothesis is rejected. The hypothesis that Brazilian banks behave like a cartel arrangement is also rejected.

1 Introduction

The Brazilian banking industry, like in many other countries, is highly concentrated. The five-bank concentration ratio ranges from 49.7% [for total assets] to 55.3% [for total loans], and to 57.9% [for total deposits] according to data for June 2000. Although concentration does not necessarily imply that the market is imperfect, it is small wonder that suspicions about cartel behavior amongst Brazilian bankers are frequently raised in the country.[2]

The aim of this paper is to perform an empirical test to evaluate the degree of competition in Brazilian banking. The main result reported here is that the available evidence suggests that the banking industry in Brazil can be described as highly competitive, although not perfect competitive. The hypothesis that Brazilian banks behave collusively is strongly rejected.

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[2] See, for example, "Acordo para não baixar juros", *Jornal do Brasil*, 24.06.2000.

The paper is organized as follows. The next section introduces the methodology to be applied in the paper. Section 3 makes the transition to the empirical implementation of the model. Section 4 describes the empirical data. Section 5 presents the main results. Section 6 concludes.

2 Methodology

The methodology followed in this paper borrows from the approach developed by Bresnahan (1982) and Lau (1982) to identify and empirically measure the degree of market competition in an industry [see Breshanan (1989) for a survey]. This methodology has been applied to bank studies for places such as Colombia [Barajas *et al*. (1999)], Uruguay [Spiller and Favaro (1984)], Canada [Shaffer (1993)], the United States [Shaffer (1989)], Finland [Vesala (1995)] and a group of some European countries [France, Denmark, Belgium, Germany, Netherlands, Spain and United Kingdom by Neven and Röller (1999)].

The empirical evidence is diverse. The above mentioned studies reach the conclusion that the banking industry is competitive in Canada, in the United States, and in Colombia. The Uruguayan banking can be represented as a Stackelberg oligopoly model of leaders and followers. The Finnish banking shows imperfect competition but no cartelization. Finally, there is evidence of cartel behavior for the European banks.

The model to be estimated below supposes that the long-run demand function for bank loans can be represented by the following expression:

$$\ln L = \alpha_1 r^L + \alpha_2 \ln Y + \alpha_3 \left(r^L \ln Y \right) \tag{1}$$

where L is the aggregate amount of bank loans in real terms, r^L is the market loan interest rate in real terms, Y is an index of economic activity, and α_1, α_2, and α_3 are coefficients to be estimated.

As emphasized by Bresnahan (1982) and Lau (1982), the interaction term between r^L and ln$Y$ rotates the demand curve for bank loans, which permits the identification of the market power parameter.

To complete the model, it is necessary to specify the bank behavior. Assume that a bank i is able to raise an amount D_i of deposits in real terms by promising to pay r^D of interest in real terms. Deposits are subject to required reserves. Let μ be the required reserve ratio. Apart from the required reverses, bank assets are composed of bank loans and public bonds. Let r^B be the real interest rate paid on public bonds. The balance sheet for bank i is then given by:

$$D_i = \mu D_i + L_i + B_i \tag{2}$$

where B_i is the amount of public bonds held by bank i in real terms.

The profit function in real terms for an individual bank i is expressed as:

$$\Pi_i = \left(r^L - r^B\right)L_i + [(1-\mu)r^B - r^D]D_i - C\left(L_i, D_i\right) \tag{3}$$

where $C(L_i;D_i)$ is the resource cost incurred by bank i to raise deposits D_i and grant loans L_i.

Assume that the cost function $C(L_i;D_i)$ is additive separable on its arguments. Assume further that bank i takes the interest rate on public bonds r^B as given. Under such assumptions, the bank activities in the loan and in the deposit markets can be treated as separable. Market power issues in the deposit market are not considered here.[3] The first order condition in the loan market is then given by:

$$\frac{\partial \Pi_i}{\partial L_i} = r^L - r^B + L_i \frac{\partial r^L}{\partial L_i} - \frac{\partial C_i}{\partial L_i} = 0 \tag{4}$$

Assume that the loan marginal cost can be described by the following expression:

$$\frac{\partial C_i}{\partial L_i} = \beta_1 \ln L_i + \beta_2 w + \beta_3 z_i \tag{5}$$

where w is the price of the inputs and z_i is a factor controlling for the quality of the output, measured here by loan losses, and β_1, β_2, and β_3 are coefficients to be estimated.

After making use of equation (1) and replacing equation (5) in (4), one gets an expression for the bank interest margin as follows:

$$r^L - r^B = \beta_1 \ln L_i + \beta_2 w + \beta_3 z_i - \lambda \left(\frac{1}{\alpha_1 + \alpha_3 \ln Y} \right) \tag{6}$$

where $\lambda = \partial \ln L / \partial \ln L_i$. In words, λ measures the percent response of the market supply of loans to a one percent increase in the loan supply by bank *i*.

The λ coefficient is the crucial parameter that summarizes the degree of average market power in the industry. If the banking industry is characterized by perfect competition then λ equals zero. In the other extreme, in the case of monopoly or cartel behavior, the λ parameter equals one. In the intermediary case of a symmetric Cournot oligopoly λ is equal to the inverse of the number of banking firms in the industry.

Note that the λ coefficient is identified only when the α_3 coefficient on the interaction term is different from zero.

Shaffer (1993) argues that the λ coefficient can be estimated without bias as long as the sample data encompasses at least one complete market. If the analyzed industry is made up of multiple markets then the estimated λ represents the average degree of market power over the various markets. It is not necessary that all the individual banks in the

[3] See Suominen (1994) for a model where market power in both deposit and loan markets are jointly investigated.

sample have the same market power. The λ coefficient reflects the behavior of the average bank (representative bank) in the sample. If, for example, a reduced group of banks behaves in a cartel like arrangement while the remaining banks behave competitively, then the estimated λ would be a weighted average of the collusive and competitive values.

The model developed thus far is a static one. However, when applied to Brazilian data static models suffer from severe misspecification problems. Besides, given the high frequency data used in the empirical section, a dynamic model allowing for adjustments over time of the endogenous variables is more appropriate. Steen and Salvanes (1999) extend Bresnahan-Lau's approach to a dynamic context. Their approach is followed here and described in the next section.

3 Empirical Model

The starting point is to write the dynamic versions of the loan demand function (1), and of the bank first order condition (6). Assume that auto-regressive distributed lag (ADL) models are suitable to describe them. The error-correction (ECM) representation of such models are respectively given by:

$$\begin{aligned}\Delta \ln L_t &= \gamma_0 + \sum_{i=1}^{k-1} \gamma_{1,i} \Delta \ln L_{t-i} + \sum_{i=0}^{k-1} \gamma_{2,i} \Delta r^L_{t-i} + \sum_{i=0}^{k-1} \gamma_{3,i} \Delta \ln Y_{t-i} \\ &+ \sum_{i=0}^{k-1} \gamma_{4,i} \Delta (r^L \ln L)_{t-i} + \gamma^* \left[\ln L_{t-k} - \alpha_1 r^L_{t-k} - \alpha_2 \ln Y_{t-k} - \alpha_3 (r^L \ln L)_{t-k} \right] + \varepsilon_t \end{aligned} \tag{7}$$

and:

$$\begin{aligned}\Delta (r^L - r^B)_t &= \delta_0 + \sum_{i=1}^{k-1} \delta_{1,i} \Delta (r^L - r^B)_{t-i} + \sum_{i=0}^{k-1} \delta_{2,i} \Delta \ln L_{t-i} + \sum_{i=0}^{k-1} \delta_{3,i} \Delta w_{t-i} + \sum_{i=0}^{k-1} \delta_{4,i} \Delta z_{t-i} \\ &+ \sum_{i=0}^{k-1} \lambda_i \Delta Y^*_{t-i} + \delta^* \left[(r^L - r^B)_{t-k} - \beta_1 \ln L_{t-k} - \beta_2 w_{t-k} - \beta_3 z_{t-k} + \lambda Y^*_{t-k} \right] + \eta_t \end{aligned} \tag{8}$$

where Δ is the first-difference operator, ε_t and η_t are the statistical error terms, and $Y_t^* = (\alpha_1 + \alpha_3 \ln Y_t)^{-1}$.

Without loss of generality, equations (7) and (8) are written with the same lag order for all variables. In the empirical estimation, the lag orders can be different for different variables. The subscript identifying individual banks is dropped from (8) on the basis that aggregate data will be used in the empirical analysis.

The specification in first-differences account for the use of non-stationary data. Unit root tests applied to the variables do not reject the hypothesis that the variables are integrated of order one I(1).

The error-correction model is a synthetic way to combine the modeling of the short-run dynamics with the consideration of the adjustments of the dependent variable towards its equilibrium level.

The error-correction specification is consistent with the presence of co-integration for the variables. Co-integration tests performed for the two sets of variables are reported in the next section. They do not reject the existence of at least one co-integration vector for each of the systems.

The estimation of the long-run parameters in the model given by equations (7) and (8) is done through the transformation due to Bårdsen (1989). The Bårdsen transformation estimates the following equations:

$$\begin{aligned}\Delta \ln L_t &= \gamma_0 + \sum_{i=1}^{k-1} \gamma_{1,i} \Delta \ln L_{t-i} + \sum_{i=0}^{k-1} \gamma_{2,i} \Delta r_{t-i}^L + \sum_{i=0}^{k-1} \gamma_{3,i} \Delta \ln Y_{t-i} \\ &+ \sum_{i=0}^{k-1} \gamma_{4,i} \Delta (r^L \ln L)_{t-i} + \gamma^* \ln L_{t-k} + \alpha_1^* r_{t-k}^L + \alpha_2^* \ln Y_{t-k} + \alpha_3^* (r^L \ln L)_{t-k} + \varepsilon_t\end{aligned} \tag{9}$$

and:

$$\Delta(r^L - r^B)_t = \delta_0 + \sum_{i=1}^{k-1} \delta_{1,i} \Delta(r^L - r^B)_{t-i} + \sum_{i=0}^{k-1} \delta_{2,i} \Delta \ln L_{t-i} + \sum_{i=0}^{k-1} \delta_{3,i} \Delta w_{t-i} + \sum_{i=0}^{k-1} \delta_{4,i} \Delta z_{t-i}$$
$$+ \sum_{i=0}^{k-1} \lambda_i \Delta Y^*_{t-i} + \delta^* (r^L - r^B)_{t-k} + \beta_1^* \ln L_{t-k} + \beta_2^* w_{t-k} + \beta_3^* z_{t-k} + \lambda^* Y^*_{t-k} + \eta_t \quad (10)$$

The long-run parameters of the model (7)-(8) can be recovered from the model (9)-(10) through the application of the following identities:

$$\begin{aligned} \alpha_i &= -\frac{\alpha_i^*}{\gamma^*}, \qquad i = 1,2,3 \\ \beta_i &= -\frac{\beta_i^*}{\delta^*}, \qquad i = 1,2,3 \\ \lambda &= \frac{\lambda^*}{\delta^*} \end{aligned} \quad (11)$$

Statistical inference about the long-run parameters is possible through the computation of the variance of the estimated coefficients using a linear approximation to the variance of the ratios in (11):

$$\begin{aligned} \mathrm{var}(\hat{\alpha}_i) &\cong \left(-\frac{1}{\hat{\gamma}^*}\right)^2 \mathrm{var}(\hat{\alpha}_i^*) + \left[\frac{\hat{\alpha}_i^*}{(\hat{\gamma}^*)^2}\right]^2 \mathrm{var}(\hat{\gamma}^*) + 2\left(-\frac{1}{\hat{\gamma}^*}\right)\left[\frac{\hat{\alpha}_i^*}{(\hat{\gamma}^*)^2}\right] \mathrm{cov}(\hat{\alpha}_i^*, \hat{\gamma}^*), i = 1,2,3 \\ \mathrm{var}(\hat{\beta}_i) &\cong \left(-\frac{1}{\hat{\delta}^*}\right)^2 \mathrm{var}(\hat{\beta}_i^*) + \left[\frac{\hat{\beta}_i^*}{(\hat{\delta}^*)^2}\right]^2 \mathrm{var}(\hat{\delta}^*) + 2\left(-\frac{1}{\hat{\delta}^*}\right)\left[\frac{\hat{\beta}_i^*}{(\hat{\delta}^*)^2}\right] \mathrm{cov}(\hat{\beta}_i^*, \hat{\delta}^*), i = 1,2,3 \\ \mathrm{var}(\hat{\lambda}) &\cong \left(\frac{1}{\hat{\delta}^*}\right)^2 \mathrm{var}(\hat{\lambda}^*) + \left[-\frac{\hat{\lambda}^*}{(\hat{\delta}^*)^2}\right]^2 \mathrm{var}(\hat{\delta}^*) + 2\left(\frac{1}{\hat{\delta}^*}\right)\left[-\frac{\hat{\lambda}^*}{(\hat{\delta}^*)^2}\right] \mathrm{cov}(\hat{\lambda}^*, \hat{\delta}^*) \end{aligned} \quad (12)$$

where a hat over a variable indicates its estimated value, var(.) is the estimated variance of the respective estimated coefficient and cov(.,.) is the estimated covariance between the respective estimated coefficients.

The empirical model to be implemented in the next sections take equations (9) and (10) as the benchmark for the analysis.

4 Data

The sample refers to aggregate data for Brazilian banks observed on a monthly basis during the period from August 1994 to August 2000. The use of data for the period previous to July 1994 is avoided due to the regime shift introduced with the launching of the stabilization program in that month.

The volume of bank loans refer to the freely allocated fixed rate bank credit, which excludes from the total bank loans those operations related to credit extended at floating rates, transfers of external funds, credit extended at binding interest rate ceilings, and earmarked resources. The lending interest rate is the weighted average of the interest rate for each bank for the same operations.

The index of economic activity is the index of industrial production as compiled by IBGE. Both the amount of bank lending as well as the index of economic activity have been seasonally adjusted by the application of the multiplicative X-11 method.

The interest rate on public bonds is a monthly compounded overnight rate (Selic).

The price of inputs is approximated by the ratio of administrative costs due to the business of granting loans to the volume of bank loans. Overall administrative costs are allocated to the different bank business units according to the participation of each such unit in the overall bank gross income. The different business units considered in the paper are the following: freely allocated loan granting, other loan granting, foreign exchange activities, securities transactions, leasing, and bank services.

Loan losses are calculated as the ratio of net provisions for non-performing loans to the volume of bank loans. Both loan losses and the price of inputs were calculated using weighted average data for a sample of seventeen large private banks.

All variables are transformed in real terms by deflating them according to the increase in the General Price Index (IGP-DI) calculated by FGV. January 2000 is taken as the basis period.

Figure 1 shows the behavior of the monthly real interest rate on bank loans and on overnight operations. The overall trend of both rates is to decrease with the passage of time. The inflation-adjusted loan interest rate went down from a high of 8.9% p.m. in October 1995 to a low of 1.6% p.m. in July 2000. By the same token, the inflation-adjusted market interest rate reduced from a peak of 3.5% p.m. in October 1995 to -0.7% p.m. in July 2000.

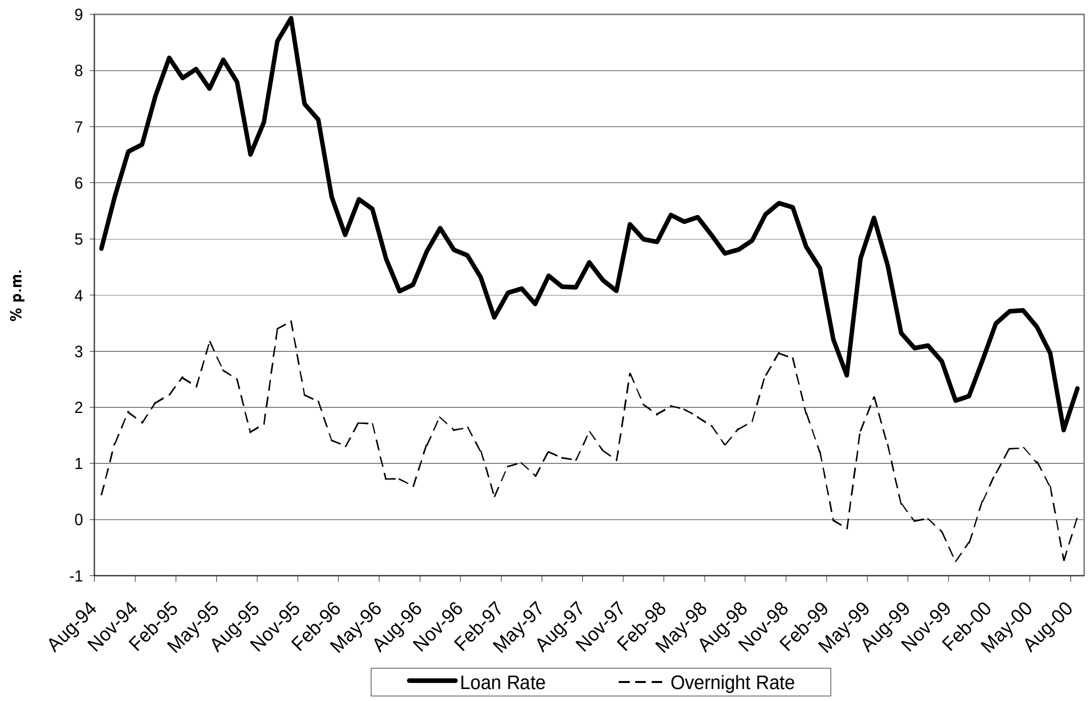

Figure 1: Monthly Loan and Overnight Inflation-Adjusted Interest Rates

Figure 1 may suggest that the similar behavior of both interest rates left the bank interest spread, as measured by the difference between them, relatively constant throughout the sample period. However, the observation of Figure 2 indicates otherwise.

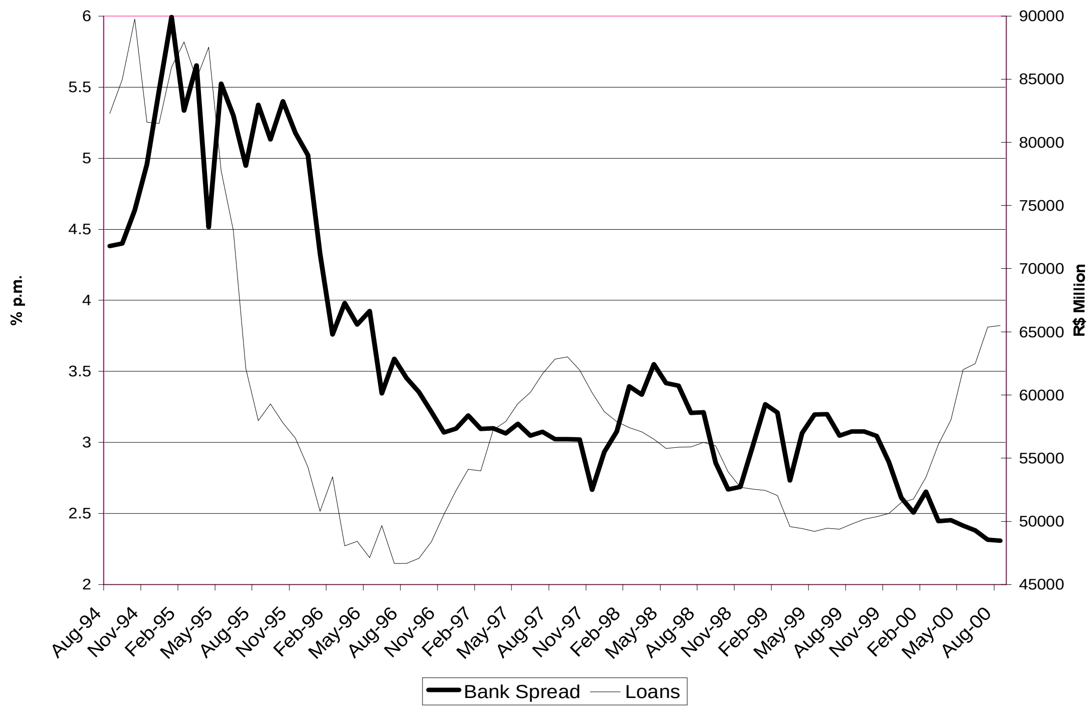

Figure 2: Monthly Bank Spread and Bank Loans

Bank interest spread also witnessed an overall decreasing trend over the period, having reached a peak of 6.0% p.m. in January 1995. By the end of the sample period, bank interest spread was at its lowest level, bottoming at 2.3% p.m.

After reaching a high of around R$ 95,000 million in the immediate aftermath of the stabilization program, bank credit suffered a dent following the restrictive monetary policy pursued in 1995 and the subsequent banking crisis. The recovery that started in late 1996 was interrupted one year later with the eruption of the Asian crisis. The upsurge in the granting of banking loans was resumed with the currency flotation and the shift in monetary policy regime towards an inflation target in 1999.

Figure 3 shows roughly that the same events underlying the upswings of the bank loans were also responsible for the fluctuations observed in the index of economic activity. The marked difference between the two series is the faster recovery of economic activity by the end of the sample period, when this variable recorded its highest level.

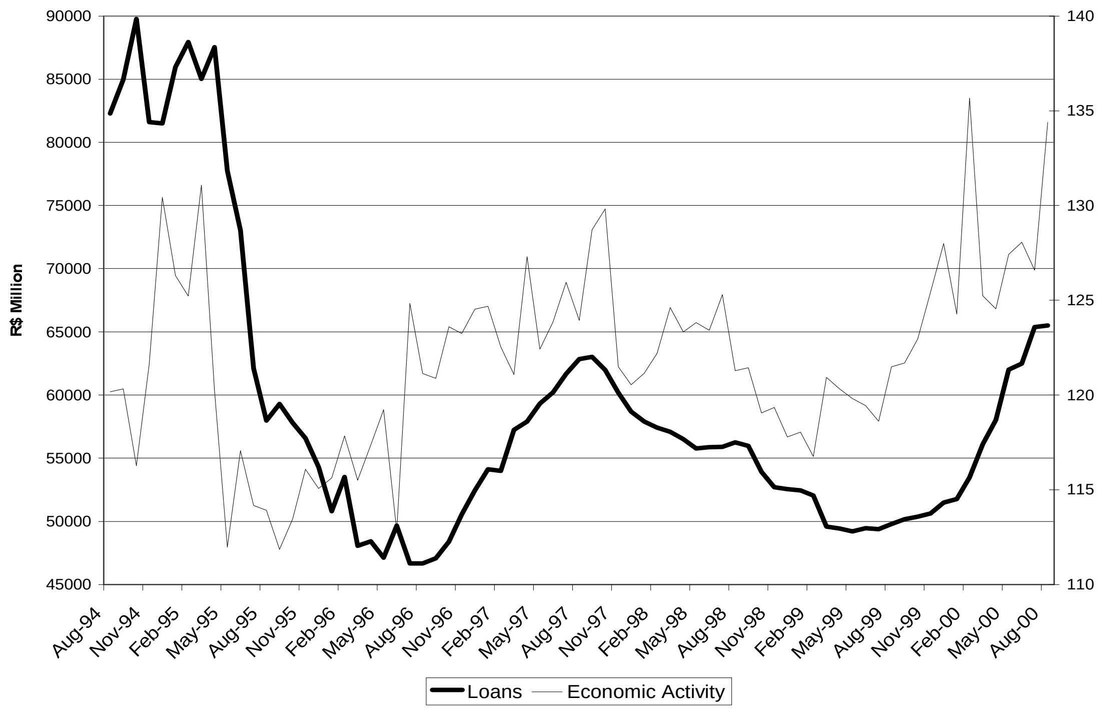

Figure 3: Bank Loans and Economic Activity

5 Results

This section presents the results of the empirical estimation of expressions (9) and (10) using monthly data for Brazil. Each of these equations was estimated via two stage least squares (2SLS) to account for endogeneity problems. All contemporaneous right side variables are treated as endogenous in the estimations.

The instruments for equation (9) include the first difference of the inflation rate, the first difference of the terms of trade, the first difference of the real overnight interest rate, and the first two lags of the latter variable.[4] The instrument set for equation (10) is composed of the contemporaneous and the first two lags of the first difference of the inflation rate, the first difference of the logarithm of the index of economic activity, and the first difference of the required reserve ratio. The validity of each instrument set is tested via the Sargan over-identification test.

The lag order in the unrestricted version of the models is seven for every variable (i.e., k is taken to be equal to seven). This lag order was chosen on the basis of two criteria, namely: a) the elimination of residual serial correlation, and b) the statistical significance of the explanatory variables at the chosen lag order.

A general-to-specific approach was then pursued through the elimination of the statistically insignificant variables. In the face of the reduced degrees of freedom for the available sample, only those variables for which the estimated coefficients implied probabilities for the t statistics below 20% were dropped from the models. The validity of the imposed restrictions was tested via the F test.

The results for the restricted model for the demand for bank loans is reported in Table 1. The estimates of the long run parameters and their standard errors are calculated from (11) and (12), respectively. The bottom of the table reports a test of the overall significance of the model, Sargan's over-identification test, an F test for residual serial auto-correlation, an F test for ARCH residuals, a normality test for the residuals, and an F test for the validity of the imposed zero restrictions on the insignificant coefficients.[5]

[4] Terms of trade is defined as the ratio of exports to imports price. This variable is calculated by FUNCEX.
[5] The same block of information is also reported at the bottom of Table 2.

VARIABLE	PARAMETER	ESTIMATE	STANDARD ERROR	t-VALUE
Constant	γ_0	-1.7386	2.3648	-0.735
$\Delta\ln L_{t-1}$	$\gamma_{1,1}$	-0.3745	0.1744	-2.148
$\Delta\ln L_{t-2}$	$\gamma_{1,2}$	0.3658	0.1762	2.076
Δr^L_t	$\gamma_{2,0}$	186.70	95.469	1.956
Δr^L_{t-2}	$\gamma_{2,2}$	121.19	67.593	1.793
Δr^L_{t-3}	$\gamma_{2,3}$	-27.932	18.251	-1.530
Δr^L_{t-4}	$\gamma_{2,4}$	118.83	53.412	2.225
Δr^L_{t-5}	$\gamma_{2,5}$	138.15	46.058	2.999
Δr^L_{t-6}	$\gamma_{2,6}$	150.71	52.204	2.887
$\Delta\ln Y_t$	$\gamma_{3,0}$	1.4198	1.0015	1.418
$\Delta\ln Y_{t-2}$	$\gamma_{3,2}$	1.1786	0.6294	1.872
$\Delta\ln Y_{t-4}$	$\gamma_{3,4}$	0.9911	0.5307	1.868
$\Delta\ln Y_{t-5}$	$\gamma_{3,5}$	1.1665	0.5518	2.114
$\Delta\ln Y_{t-6}$	$\gamma_{3,6}$	1.1847	0.5964	1.987
$\Delta(r^L\ln Y)_t$	$\gamma_{4,0}$	-38.826	19.909	-1.950
$\Delta(r^L\ln Y)_{t-2}$	$\gamma_{4,2}$	-24.865	14.034	-1.772
$\Delta(r^L\ln Y)_{t-3}$	$\gamma_{4,3}$	6.0600	3.8369	1.579
$\Delta(r^L\ln Y)_{t-4}$	$\gamma_{4,4}$	-24.729	11.151	-2.218
$\Delta(r^L\ln Y)_{t-5}$	$\gamma_{4,5}$	-28.742	9.5504	-3.009
$\Delta(r^L\ln Y)_{t-6}$	$\gamma_{4,6}$	-31.593	10.914	-2.895
Long run parameters				
$\ln L_{t-7}$	γ^*	-0.16056	0.0669	-2.400
r^L_{t-7}	α_1	598.58	311.85	1.919
$\ln Y_{t-7}$	α_2	4.5509	3.1463	1.446
$(r^L\ln Y)_{t-7}$	α_3	-125.16	65.610	-1.908

$\sigma_\varepsilon = 0.0306$ Model test: $\chi^2(23) = 74.411$ Sargan: $\chi^2(2) = 1.1125$

AR 1-5: $F(5,37) = 1.151$ ARCH 5: $F(5,32) = 0.534$ Normality: $\chi^2(2) = 0.886$

Number of observations: 66 Validity of imposed restrictions: $F(8,34) = 1.419$

Table 1: Estimates of the Demand for Bank Loans

The model shows no sign of misspecification. The set of instruments is a valid one given the non-significance of the Sargan statistics. Finally, the restricted model seems to be a valid simplification of the unrestricted one.

The signs of the estimated coefficients cannot be directly interpreted due to the interaction terms. The long run price elasticity can be calculated as $r^L(\alpha_1 + \alpha_3 \ln Y)$. Evaluated at the sample means of r^L and Y the own price elasticity is equal to -0.128. Thus the demand curve for bank loans is downward sloping. However, the inelastic long-run demand curve suggests that a perfect collusive market structure is not consistent with the available data. The evidence from the direct estimation of the λ coefficient, to be presented below, confirms such findings.

The estimate of the adjustment parameter γ^* is reasonable, indicating the operation of the error-correction mechanism. In words, when the demand for bank loans deviates from its long-run path, bank loans are adjusted in the short-term to bring them back on course.

The significance on the coefficient on the interaction term α_3 indicates that the market power parameter λ is identified.

Table 2 records the estimates of the restricted model for the equation for bank interest spread. The estimated model is well specified and the imposed restrictions are not rejected. The set of instruments do not reject Sargan's over-identification test.

VARIABLE	PARAMETER	ESTIMATE	STD ERROR	t-VALUE
Constant	δ_0	-0.2229	0.0689	-3.238
$\Delta(r^L - r^B)_{t-1}$	$\delta_{1,1}$	-0.9196	0.2039	-4.511
$\Delta(r^L - r^B)_{t-2}$	$\delta_{1,2}$	-1.0057	0.3030	-3.319
$\Delta(r^L - r^B)_{t-3}$	$\delta_{1,3}$	-0.9549	0.2097	-4.555
$\Delta(r^L - r^B)_{t-4}$	$\delta_{1,4}$	-0.8715	0.2365	-3.685
$\Delta(r^L - r^B)_{t-5}$	$\delta_{1,5}$	-0.6626	0.2498	-2.653
$\Delta(r^L - r^B)_{t-6}$	$\delta_{1,6}$	-0.7009	0.1972	-3.554
Δw_{t-1}	$\delta_{3,1}$	-0.4067	0.3046	-1.335
Δz_{t-2}	$\delta_{3,2}$	0.2773	0.0894	3.101
Δz_{t-3}	$\delta_{3,3}$	0.5092	0.1621	3.142
Δz_{t-4}	$\delta_{3,4}$	0.8531	0.2605	3.274
Δz_{t-5}	$\delta_{3,5}$	1.0699	0.3253	3.289
Δz_{t-6}	$\delta_{3,6}$	1.1566	0.3660	3.160
ΔY^*_t	λ_0	-3.9011×10^{-4}	2.9261×10^{-4}	-1.333
ΔY^*_{t-1}	λ_1	-5.0467×10^{-4}	3.3209×10^{-4}	-1.520
ΔY^*_{t-2}	λ_2	-5.2845×10^{-4}	3.3503×10^{-4}	-1.577
ΔY^*_{t-3}	λ_3	-5.3782×10^{-4}	3.4324×10^{-4}	-1.567
ΔY^*_{t-4}	λ_4	-7.988×10^{-4}	3.8259×10^{-4}	-2.088
ΔY^*_{t-5}	λ_5	-8.43×10^{-4}	4.1853×10^{-4}	-2.014
ΔY^*_{t-6}	λ_6	-8.1484×10^{-4}	4.5632×10^{-4}	-1.786
Long run parameters				
$(r^L - r^B)_{t-7}$	δ^*	-0.5043	0.1200	-4.202
$\ln L_{t-7}$	β_1	0.0422	4.6878×10^{-3}	9.007
w_{t-7}	β_2	1.1863	0.3471	3.4173
z_{t-7}	β_3	2.2486	0.8764	2.8168
Y^*_{t-7}	λ	1.7482×10^{-3}	8.563×10^{-4}	2.0416

$\sigma_\eta = 0.0023$ Model test: $\chi^2(24) = 64.937$ Sargan: $\chi^2(8) = 10.515$

AR 1-5: $F(5,35) = 2.009$ ARCH 5: $F(5,30) = 0.229$ Normality: $\chi^2(2) = 1.857$

Number of observations: 65 Validity of imposed restrictions: $F(15,25) = 0.1325$

Table 2: Estimates for the Bank Interest Spread

The adjustment parameter δ^* indicates that bank interest spread adjusts faster to deviations from the long-run equilibrium than the demand for bank loans does.

The long-term equation for the bank interest spread has the expected signs for all the coefficients. The bank interest spread increases when the amount of extended loans increases, when the price of inputs is higher, and when default losses increase.

The market power parameter λ is equal to 0.0017. Despite its small numerical value it is statistically significant at 5% level. This result hints that Brazilian banks do not behave competitively. However, this result does not imply that the Brazilian banking industry works as a cartel. The cartel arrangement hypothesis implies that the market power parameter λ should be equal to one, which is strongly rejected by the data: the *t*-value to test this hypothesis is equal to 1165.77.

Even in the short-term, there is evidence that Brazilian banks do not behave competitively. The estimated short-term market power coefficients are statistically significant and their numerical values are one order of magnitude smaller than the long-term market power parameter. This result means that Brazilian banks exercise more market power in the long rather than in the short-term.

The inference concerning the long-term parameters was based on the approximate formulas for the variances of the estimated coefficients given by (12). The possible biases resulting from such approximation are unknown. The robustness of the results can be checked by means of the co-integration analysis.

The co-integration tests are those due to Johansen (1988) and they were applied to the systems ($\ln L$; r^L; $\ln Y$; $r^L \times \ln Y$) and ($r^L - r^B$; $\ln L$; w; z; Y^*). The first system represents the demand for bank loans while the latter encompasses the equation for the bank spread.

The results of the co-integration tests are reported in Table 3. The first row of the table refers to the maximal eigenvalue and to the trace statistics, respectively. The critical values for both statistics are those from Osterwald-Lenum (1992). In the first column of the table, r refers to the rank of the long-run matrix of coefficients, which coincides with the number of co-integration vectors detected in the system. Seven lags of each variable are introduced for both systems.

	Maximal	**Trace**
	Loan demand equation	
$r = 0$	33.07^{**}	59.88^{**}
$r = 1$	16.18	26.81
$r = 2$	9.16	10.63
$r = 3$	1.469	1.469
	Bank spread equation	
$r = 0$	43.96^{**}	116.9^{**}
$r = 1$	30.92^{*}	72.95^{**}
$r = 2$	22.63^{*}	42.03^{**}
$r = 3$	17.27^{*}	19.4^{*}
$r = 4$	2.123	2.123

* Significant at 5% ** Significant at 1%

Table 3: Johansen Co-integration Tests

The results show that there is evidence of one co-integration vector for the first system and of three or four co-integration vectors for the second. Such results imply that the error-correction representation is suitable for the data.

Steen and Salvanes (1999) suggest to test the significance of the coefficient of the interaction term in the loan demand equation by imposing a zero restriction in the co-integration vector. Under the restriction that there is one co-integration vector in the loan demand system, the null hypothesis that r^L and $\ln Y$ are separable is rejected by the data. The test statistics is equal to 12.724 with a $\chi^2(1)$ asymptotic distribution under the null hypothesis. This result confirms the previous findings based on the single equation framework with the implication that the market power parameter λ is indeed identified.

The significance of the market power parameter λ can also be alternatively tested by means of a zero restriction imposed on the coefficient of the Y^* variable in the co-integration vector. Under the assumption that there is one co-integration vector in the bank spread system, the null hypothesis that Brazilian banks behave competitively is again rejected. The test statistics is equal to 12.951 with a $\chi^2(1)$ asymptotic distribution under the null hypothesis. This result is in agreement with the ones reported above for the bank spread equation.

Summing up, the available evidence indicates that Brazilian banking is not competitive. Although the exact form of market structure is not identified it is possible to reject that Brazilian banks form a cartel.

6 Concluding Remarks

This paper adopts Bresnahan-Lau's approach to test the significance of market power in Brazilian banking. A dynamic specification is pursued. Error-correction models for both the demand for bank loans and the bank interest spread are developed.

The results are consistent with the view that Brazilian banks have some market power. Such market power is more pronounced in the long rather than in the short-term. The precise market structure is not known but the data strongly rejects that perfect collusion is practiced by Brazilian banks.

The conclusions reached in this paper are negative rather than positive in the sense that all that it was possible to infer was that Brazilian banking cannot be described by any of the two polar market structures, namely perfect competition and monopoly/cartel. It is left to future research an attempt to refine and better understand the exact nature of market imperfection characterizing the banking industry in Brazil.

References

Barajas, A., Steiner, R., and Salazar, N. (1999). "Interest spreads in banking in Colombia,1974-96". *IMF Staff Papers*, **46**, 196–224.

Bårdsen, G. (1989). "The estimation of long-run coefficients from error-correction models". *Oxford Bulletin of Economics and Statistics*, **51**, 345–350.

Bresnahan, T. F. (1982). "The oligopoly solution is identified". *Economics Letters*, **10**, 87-92.

Bresnahan, T. F. (1989). "Empirical studies of industries with market power". In Schmalensee, R. and Willig, R. D., editors, *Handbook of Industrial Organization*, Volume II. Elsevier Science Publishers, Amsterdam, NL.

Johansen, S. (1988). "Statistical analysis of cointegration vectors". *Journal of Economic Dynamics and Control*, **12**, 231–254.

Lau, L. J. (1982). "On identifying the degree of competitiveness from industry price and output data". *Economics Letters*, **10**, 93-99.

Neven, D. and Röller, L.-H. (1999). "An aggregate structural model of competition in the European banking industry". *International Journal of Industrial Organization*, **17**, 1059–1074.

Osterwald-Lenum, M. (1992). "A note with fractiles of the asymptotic distribution of the maximum likelihood cointegration rank test statistics: four cases". *Oxford Bulletim of Economics and Statistics*, **54**, 461–472.

Shaffer, S. (1989). "Competition in the U.S. banking industry". *Economics Letters*, **29**, 321–323.

Shaffer, S. (1993). "A test of competition in Canadian banking". *Journal of Money, Credit, and Banking*, **25**, 49–61.

Spiller, P. T. and Favaro, E. (1984). "The effects of entry regulation on oligopolistic interaction: the Uruguayan banking sector". *Rand Journal of Economics*, **15**, 244–254.

Steen, F. and Salvanes, K. G. (1999). "Testing for market power using a dynamic oligopoly model". *International Journal of Industrial Organization*, **17**, 147–177.

Suominen, M. (1994). "Measuring competition in banking: a two-product model". *Scandinavian Journal of Economics*, **96**, 95–110.

Vesala, J. (1995). *Testing for Competition in Banking: Behavioral Evidence from Finland*. Bank of Finland Studies E:1. Bank of Finland.

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