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Assessing market power in the U.S. commercial banking industry under deregulation

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ABSTRACT

This study attempts to investigate market power in the U.S. commercial banking industry since the U.S. government began to deregulate the banking sector in the early 1990s using the static Bresnahan–Lau model (SBLM) and dynamic Bresnahan–Lau model with error corrections (DBLEC). In particular, panel unit root and panel cointegration techniques are utilized to examine the dynamic model. The empirical results of the SBLM show that the banking industry is highly competitive. The empirical results of DBLEC also suggest that the commercial banking industry is close to being perfectly competitive in the short run. By contrast, the adjustment speeds of the supply and demand sides towards the long-run equilibrium are quite slow in that market, which implies that the U.S. commercial banks enjoy a certain degree of long-run market power.

1. Introduction

The U.S. financial sector has been highly innovative since the 1960s. For example, the widespread issuance of large-value, fixed-term negotiable certificates of deposit, and new types of futures and options contracts met the rapidly growing demand for liquid securities. In the early 1990s, the U.S. government began to relax its stringent regulations by eliminating legal barriers to mergers and acquisitions. The Federal Deposit Insurance Corporation Improvement Act (FDICIA) of 1991 allowed for unrestricted merger and acquisition activity (in compliance with the Bank Merger Act) between national banks and all types of credit institutions.[2] In the late 1990s, 46 state governments permitted out-of-state banks to acquire banks based on their states. In 1999, the Gramm-Leach-Bliley Financial Services Modernization Act repealed Glass-Steagall restrictions on banks that had lasted for over 60 years. While banks continued to grow larger through merger and acquisition activities, new types of financial instruments, such as commercial bills, junk bonds and financial derivatives, came to play an increasingly important role in banks' operations.

In addition to the number of banks and the concentration ratio of the industry (Berger et al., 2004), the government regulations (Canhoto, 2004; Cetorelli, 2004; Demirgüç-Kunt et al., 2004; Ho, 2010), the soundness of the financial system as a whole, the shape taken by banking networks and the level of financial innovation have also had substantial impacts on the degree of competition within the banking industry (Cetorelli, 2004; Northcott, 2004). For instance, Cetorelli's (2004) and Ho's (2010) empirical results showed that the banking industry in the EU and Hong Kong have become more competitive following deregulation.

In this study, we analyze the quarterly panel data of 338 listed U.S. commercial banks over the period from the first quarter of 1990 to the fourth quarter of 2005. The total number of observations is 21,632. We apply the static Bresnahan–Lau model (SBLM) (Bresnahan, 1989) as well as the dynamic Bresnahan–Lau model with error corrections (DBLEC), derived from panel unit root tests and panel cointegration analyses, to examine the degree of short-run and long-run market power in the U.S. commercial banking industry.

The remainder of the paper is organized as follows. Section 2 reviews the literature on the SBLM and DBLEC in regard to market power. Section 3 constructs the empirical models and describes the data. Section 4 analyzes the empirical results, while the final section concludes the paper.

2. Literature review

In the past few decades, the SBLM (Bresnahan, 1982, 1989; Lau, 1982) has been used to examine market power in the banking,

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[1] Tel.: +886 2 29393091x81037.
[2] In the early 1990s, the banking industry in the European Union (EU) has also been deregulated to branch freely across borders within EU countries (Cetorelli, 2004).

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agriculture and fishery, and power industries around the world.[3] In the area of banking, Shaffer (1989, 1993) examined the U.S. and Canadian banking industries. The market power of both banking markets was found to be insignificant, while the loan, deposit (Bikker and Haaf, 2000) and consumer credit (Toolseman, 2002) markets in the Netherlands exhibited almost perfect competition after 1983. On the basis of Bikker's (2003) study, both the housing loan and corporate lending markets in nine members[4] of the EU were found to be highly competitive during the period 1976–1998. This result is consistent with Nathan and Neave's (1989) findings for the same industry.

However, Shaffer and Disalvo (1994) studied the banking duopoly that existed in Fulton County, Pennsylvania in the U.S. over the period from 1970 to 1986 and from 1976 to 1986, respectively. They concluded that during both periods, the situation in Fulton County was somewhere between that of a duopoly and a competitive market, whereas the degree of market power possessed by the two banks was relatively low. Zardkoohi and Fraser (1998) investigated the impact of regional deregulation on competition in the U.S banking industry during the period from 1964 to 1993, and found that the impact of deregulation on competition varied significantly from state to state. In some states, the abolition of controls actually led to an increase in the market power of banks operating in that state.

Neven and Róller's (1999) empirical results showed that the loans markets of seven European countries were characterized by collusive behavior between 1981 and 1990, but that the intensity of such behavior declined over time. Suominen (1994) revealed that monopoly power was statistically significant in the deposit and loan markets in Finland from 1986 to 1989. Angelini and Cetorelli (2000) and Canhoto (2004) respectively investigated Italian banking assets from 1983 to 1997 and Portuguese deposit markets in the early 1990s. They found that both markets enjoyed a certain level of market power. Bikker and Haaf (2000) also confirmed that the Portuguese deposit market was oligopolistic during the period from 1983 to 1998, while the loan and deposit markets were highly competitive in another eight European countries. Móré and Nagy (2004) found the Hungarian credit market to be much less competitive than those of the EU member nations during the period 1996–2003.

As for other industries, Hjalmarsson (2000) applied the DBLEC to an analysis of the electric power market in Norway and Sweden. In the agriculture and fishery industries, Buschena and Perloff (1991) showed that legal and institutional changes after 1973 significantly strengthened the market power of coconut oil refining and exporting firms in the Philippines. On the other hand, Deodhar and Sheldon (1997) found the world market for soymeal exports to be perfectly competitive, while Hatirli et al. (2003) faced the same consequence in regard to banana imports by Turkey. By employing the SBLM, Frank and Steen (2006) suggested that the Salmon Agreement in 1997 enhanced the Norwegian salmon's market power in the French market. By contrast, Steen and Salvanes (1999) found that the French salmon market was more or less fully competitive using the SBLM and DBLEC.

To date, relatively few studies have employed the SBLM and DBLEC to examine market power, particularly in the banking industry. The current study aims to fill this gap in the literature by exploring both short-run and long-run competition in the U.S. commercial banking industry over the period 1990–2005.

[3] Studies using the PR (Panzar and Rosse, 1987) model also focus on individual banks and their competitive strategies (Bikker and Haaf, 2002). However, the scaled variable acting as a dependent variable often leads to distorted parameter estimates. On the contrary, the BL (Bresnahan, 1982, 1989; Lau, 1982) model is much accurate and refined in the examination of market power (Bikker, 2003). Gunalp and Celik (2006) employed the PR model to assess the competitive environment and industrial characteristics of the Turkish banking industry especially during the financial liberalization era from 1980 to 2000.

[4] They are Belgium, the UK, France, the Netherlands, Germany, Italy, Portugal, Spain and Sweden.

3. Data and methodology

3.1. Definition of variables

Financial institutions produce a variety of outputs, whose characteristics differ significantly from those of manufacturing firms. The prices of a bank's inputs and outputs are difficult to measure. In addition, distinguishing between inputs and outputs is a challenging task.

Favero and Papi (1995) proposed five methods to measure bank inputs and outputs, based on the role played by a bank.[5] The intermediation approach has been extensively employed in the literature, due partially to the availability of financial data. This approach treats a bank as a financial intermediary that hires labor, capital and various funds to produce financial services, such as loans, investments and non-traditional activities. An ordinary U.S. commercial bank appears to operate in this manner. With this approach, the emphasis is on the process of transferring funds from companies and households that have surplus funds to those that are lacking in funds. We therefore adopt the intermediation approach to define banks' inputs and outputs.[6]

Following Shaffer (1989), Shaffer and Disalvo (1994), Zardkoohi and Fraser (1998) and Móré and Nagy (2004), a bank's balance sheet entry "total loans & leases, gross" is selected as the output quantity (Q). Its price (P) is calculated as the ratio of the interest and fee income from loans to total loans and leases. A higher price of the output is anticipated to be negatively associated with the quantity demanded for bank loans.

We use the real GDP as an income level (Y) indicator, as did Shaffer (1989) and Móré and Nagy (2004). Any one of the interest rate on the U.S. three-month Treasury Bills, the interest rate on commercial paper, or the interest rate on informal loans is most often chosen as the exogenous variable affecting the demand for bank loans. Because of the unavailability of the rate on comprehensive commercial paper and the difficulty in acquiring informal loan interest rates during the sample period, we use the interest rate on three-month Treasury Bills to act as the exogenous variable (Z).[7] Shaffer (1993) adopted the same indicator to examine the impact of the indicator's fluctuations on lending in the Canadian banking industry.

Two bank-specific input prices are identified to explain the variations in output prices. One of them is the average wages (W^1) and the other is the interest rate on total deposits, calculated as the ratio of interest on deposits to total deposits (W^2). This definition is consistent with that in Zardkoohi and Fraser (1998). The higher are the input prices, the higher are the bank's output prices. The data required to compute the aforementioned variables are taken from the Bureau of Labor Statistics (BLS) and the Federal Reserve Banks of Chicago and St. Louis, respectively. They comprise quarterly data for 338 listed commercial banks in the U.S. stock market over the period 1990Q1–2005Q4. Table 1 describes the data sources and variable definitions.[8]

3.2. The Static Bresnahan–Lau Model (SBLM)

In the following, we begin by describing a simple algebraic demonstration and general theoretical structure of the Bresnahan–Lau Model.

[5] The five methods are: (1) the production approach; (2) the user cost approach; (3) the value-added approach; (4) the intermediation approach; (5) the asset approach.

[6] Shaffer (1989), Favero and Papi (1995) and Noulas (1997) all used the intermediation approach to measure bank inputs and outputs.

[7] Zardkoohi and Fraser (1998) adopted the interest rate on six-month commercial paper to investigate the U.S. banking industry, while the three-month data fit our quarterly time span.

[8] Appendix A shows that the pairwise collinearity hypotheses of the correlation matrix of the independent variables are rejected. The highest variance inflation factor (VIF) is that of the price (P), whose VIF equals 7.79. The second highest VIF is that of Z equaling 5.56, followed by income (Y) (1.52), the price of labor (W^1) (1.22), and the price of funds (W^2) (1.21), respectively. This implies that these variables are appropriate and can be employed in the regression analysis.

Table 1
The descriptions of data.

Variables	Definitions	Data sources
Quantity of output (Q)	Total loans & leases, gross	Federal Reserve Bank of Chicago
Price (P)	Interest and fee income from loans/total loans & leases, gross	Federal Reserve Bank of Chicago
Income (Y)	Real GDP	Federal Reserve Bank of St. Louis
Other price (Z)	Interest rate on government three-month Treasury Bills	Federal Reserve Bank of St. Louis
Price of labor (W^1)	Average wage	Bureau of Labor Statistics
Price of funds (W^2)	Interest on deposits/total deposits	Federal Reserve Bank of Chicago

Bresnahan (1982) and Lau (1982) subsequently showed that the market power can be identified by the rotation of the demand curve (Chintrakarn and Jindapon, 2012), while the equilibrium of marginal revenue and marginal cost are specified and estimated as a simultaneous system (Bask et al., 2011). Built upon the pioneering work by Bresnahan (1982, 1989) and Lau (1982), these recent studies analyze the extent of market power within a demand and supply framework.

Under the assumption of profit maximization, Bresnahan (1982, 1989) and Lau (1982) proposed the so-called SBLM, which argues that a firm's marginal costs should be equal to the realized marginal benefit when estimating market power for different types of market structures. The market demand equation may be formulated as $Q = D(P, X; \alpha) + \varepsilon$, where Q is the demand for some financial product, P is its price, X is a vector of exogenous variables, such as the interest rate on three-month Treasury Bills (Z) and income (Y), α consists of the corresponding parameters to be estimated, and ε denotes a random disturbance. For a supplier pursuing profit maximization, the following optimality condition must hold: $P = C'(Q, W^1, W^2; \beta) - \lambda \cdot h(\cdot) + \eta$, where $C'(\cdot) = \partial C(Q)/\partial Q$ is the marginal cost function, W^1 and W^2 represent the prices of labor and funds, respectively, β consists of the corresponding parameters, $h(\cdot) = Q\partial D^{-1}(Q)/\partial Q$, λ is the indicator of market power, and η denotes the supply shock. Note that here the term $P + \lambda \cdot h(\cdot)$ signifies the firm's marginal revenue. When $\lambda = 0$, the market is perfectly competitive. The output price is now equal to a price-taking firm's marginal cost. When $\lambda = 1$, the market is fully monopolistic and the sole firm has substantial market power. Finally, when $0 < \lambda < 1$, the market is imperfectly competitive and firms have some degree of market power. The demand and supply equations of the SBLM for the ith firm at time t are conventionally specified as:

$$Q_{it} = \alpha_0 + \alpha_P P_{it} + \alpha_Y Y_{it} + \alpha_Z Z_{it} + \alpha_{PZ} P_{it} Z_{it} + \varepsilon_{it}. \tag{1}$$

and

$$P_{it} = \beta_0 + \beta_Q Q_{it} + \beta_{W^1} W^1_{it} + \beta_{W^2} W^2_{it} - \lambda \tilde{Q}_{it} + \eta_{it}, \tag{2}$$

respectively, where $\tilde{Q}_{it} = Q_{it}/(\alpha_P + \alpha_{PZ} Z_{it})$. Gollop and Roberts (1979) and Appelbaum (1982) exploited firm-level data to conduct empirical studies of market power.

3.3. The dynamic Bresnahan–Lau model with error corrections (DBLEC)

The SBLM ignores the dynamic effects on the demand side and supply side since there are no lagged or lead variables contained. However, the dynamic formulation could be helpful in solving the short-run deviations from the market equilibrium and long-run dynamics in a market. In this section, we drive the Bresnahan–Lau model with error corrections (Steen and Salvanes, 1999) to capture the dynamics of the (non-)stationary panel data.

If one aims to examine dynamic competition in both the short run and the long run, then the use of a conventionally autoregressive distributed lag (ADL) model, such as

$$\begin{aligned} Y_t = {} & \alpha_0 + \beta_0 X_t + \beta_1 X_{t-1} + \beta_2 X_{t-2} + \ldots + \beta_p X_{t-p} + \gamma_1 Y_{t-1} \\ & + \gamma_2 Y_{t-2} + \ldots + \gamma_q Y_{t-q} + u_t, \end{aligned} \tag{3}$$

may create spurious relationships when time series data are non-stationary, while $|\hat{\gamma}_i| = 0, i = 1, 2, \ldots, q$ indicates the absence of a long-run equilibrium. Therefore, we have to transform these variables by taking the first differences. This process can easily be shown to convert non-stationary data into stationary data at the expense of removing information on the long-run equilibrium.[9] This problem can be resolved by relying on an error correction model (ECM) that dates back at least to Engle and Granger (1987). An ECM includes both short-run dynamics and a long-run equilibrium relationship with the latter being derived from a cointegration relationship. Both components of an ECM are stationary so that the difficulty faced by spurious regression is no longer present.

Numerous studies have applied the ECM to explain various economic hypotheses. Among them, Steen and Salvanes (1999) used an ECM to test for market power. Rewriting Eq. (3) as an ECM, which is referred to as the DBLEC, we obtain[10]:

$$\begin{aligned} \Delta Y_t = {} & \beta_0 \Delta X_t - \sum_{j=2}^{p} B_i \Delta X_{t-j+1} - \sum_{k=2}^{q} \Gamma_k \Delta Y_{t-k+1} - (1-\Gamma_1) \\ & \cdot [Y_{t-1} - \alpha_0/(1-\Gamma_1) - B_0 X_{t-1}/(1-\Gamma_1)] + u_t \\ \Delta Y_t = {} & \beta_0 \Delta X_t - \sum_{j=2}^{p} B_i \Delta X_{t-j+1} - \sum_{k=2}^{q} \Gamma_k \Delta Y_{t-k+1} - (1-\Gamma_1) \\ & \cdot [Y_{t-1} - \alpha_0/(1-\Gamma_1) - B_0 X_{t-1}/(1-\Gamma_1)] + u_t, \end{aligned} \tag{4}$$

where Δ denotes the first difference operator and $\xi_{t-1} = Y_{t-1} - \alpha_0/(1-\Gamma_1) - B_0 X_{t-1}/(1-\Gamma_1)$ is the error correction term obtained by using the cointegration parameter vector $(1 - \alpha_0/(1-\Gamma_1) - B_0/(1-\Gamma_1))$, where $0 \le \left|\sum_{i=1}^{q} \hat{\gamma}_i\right| = \Gamma_1 \le 1$. The coefficient of the error correction term $(1-\Gamma_1)$ represents the speed of adjustment towards the long-run equilibrium. It has to be nonnegative, implying that an excess demand, for example, for a product in the current period will be removed by $(1-\Gamma_1)$100% in the next period. The closer that the value of $(1-\Gamma_1)$ is to zero, the slower that the rate of adjustment speed will be. At the other extreme, the closer that the value is to 1, the faster that the rate of adjustment speed towards the long-run steady-state will be.

Hence, our DBLEC can be expressed as follows:

A. Demand function

$$\begin{aligned} \Delta Q_{it} = {} & \alpha_0 + \sum_{s=1}^{k_1-1} \alpha_{Q,s} \Delta Q_{it-s} + \sum_{s=0}^{k_2-1} \alpha_{P,s} \Delta P_{it-s} + \sum_{s=0}^{k_3-1} \alpha_{Y,s} \Delta Y_{it-s} \\ & + \sum_{s=0}^{k_4-1} \alpha_{Z,s} \Delta Z_{it-s} + \sum_{s=0}^{k_5-1} \alpha_{PZ,s} \Delta P_{it-s} Z_{it-s} \\ & + \hat{\gamma}(Q_{it-1} - \theta_P P_{it-1} - \theta_Y Y_{it-1} - \theta_Z Z_{it-1} - \theta_{PZ} P_{it-1} Z_{it-1}) + u_{it}. \end{aligned} \tag{5}$$

The lags for the independent and dependent variables in Eq. (5) are $k_j, j = 1, \ldots, 5$. The notations ΔQ_{it}, ΔP_{it}, ΔY_{it} and ΔZ_{it} represent the rate of change over time periods t and $t-1$ for each individual bank i. The $\alpha_{n,s}$, $n = Q, P, Y, Z, PZ$ are the coefficients of the differenced variables, representing the short-run marginal effects of these differenced variables on the dependent variable. The $\theta_n (= \alpha_n/\hat{\gamma})$ are the coefficients of the cointegration vector, reflecting the long-run equilibrium relationship between market demand and price. For example, θ_P shows the impact of P_{it-1} on Q_{it-1} in the long-run steady-state.

[9] This is also known as an over-differencing problem.

[10] The details of the derivation are described in Appendix B.

B. Supply function

$$\begin{aligned}\Delta P_{it} &= \beta_0 + \sum_{s=1}^{k_6-1}\beta_{P,s}\Delta P_{it-s} + \sum_{s=0}^{k_7-1}\beta_{Q,s}\Delta Q_{it-s} + \sum_{s=0}^{k_8-1}\beta_{W^1,s}\Delta W^1_{it-s} \\ &\quad + \sum_{s=0}^{k_9-1}\beta_{W^2,s}\Delta W^2_{it-s} + \sum_{s=0}^{k_{10}-1}\lambda_s\Delta\tilde{Q}_{it-s} \\ &\quad + \hat{\psi}\left(P_{it-1} - \xi_Q Q_{it-1} - \xi_{W^1}W^1_{it-1} - \xi_{W^2}W^2_{it-1} - \Lambda\tilde{Q}_{it-1}\right) + \eta_{it}.\end{aligned} \tag{6}$$

The lags for the independent and dependent variables in Eq. (6) are k_l, $l = 6, \ldots, 10$. The $\beta_{n,\,s}$, $n = P, Q, W^1, W^2$, are the coefficients of the differenced variables. In addition, the coefficients within the error correction term $\hat{\psi}$, $\xi_Q = \beta_Q/\hat{\psi}$, $\xi_{W^1} = \beta_{W^1}/\hat{\psi}, \xi_{W^2} = \beta_{W^2}/\hat{\psi}$ and $\Lambda = \lambda/\hat{\psi}$ are defined as the corresponding ones of Eq. (5), and can be explained in an analogous manner. Unlike the SBLM, the DBLEC distinguishes the impacts of short-run factors from the long-run equilibrium.

At the outset, we have to pre-test whether individual variables are stationary. Several panel unit root testing procedures are thus employed. If all variables are found to be stationary, then the use of the conventional least squares method is legitimate. Otherwise, more attention must be paid. A test for the existence of panel cointegration among the variables of interest needs to be implemented.

Table 2 reveals that the variables output quantity (Q), income (Y), and the price of labor (W^1) are consistently found to be nonstationary by the six testing procedures, while the remaining variables of output price (P), other price (Z), the price of funds (W^2) and PZ have inconsistent outcomes among the six methods. For variables P and W^2, the three testing approaches that assume a common unit root confirm the presence of a panel unit root, whereas the other three approaches presume that individual unit roots confirm the absence of a panel unit root. Thus only two testing procedures, which assume that there is a common unit root, suggest that Z and PZ are nonstationary, while the other four procedures fail to do so. To sum up, we still regard all of the seven variables as having a panel unit root for the purpose of taking care of possible nonstationarity, on the one hand, and of avoiding the potential problem of over-differencing the model, on the other.

We utilize the SIC (Schwartz information criterion) to determine the optimal lag length for the vector autoregressive model. The results show that, for the variables ΔQ_{it}, ΔP_{it}, ΔY_{it}, ΔZ_{it} and $\Delta P_{it}Z_{it}$ of the demand function, the optimal lags (k_1,k_2,k_3,k_4,k_5) are found to be 2, 1, 1, 1 and 2, respectively. For the variables $\Delta P_{it},\Delta Q_{it},\Delta W^1_{it},\Delta W^2_{it}$, and $\Delta\tilde{Q}_{it}$ of the supply function, the optimal lags (k_6,k_7,k_8,k_9,k_{10}) are found to be 1, 2, 1, 8 and 3, respectively. We next apply the residual cointegration testing procedure, developed by Pedroni (2000, 2004) and Kao (1999), to investigate the existence of a long-run equilibrium for the demand and supply Eqs. of (5) and (6). As shown in Appendix C, the two test statistics reject the null hypothesis of no cointegration, indicating that there exist long-run equilibrium relationships among variables in the two equations. This finding justifies the superiority of DBLEC over the static model in the study of market power.

Eqs. (5) and (6) can now be reformulated as:

$$\begin{aligned}\Delta Q_{it} &= \alpha_0 + \alpha_{Q,1}\Delta Q_{it-1} + \alpha_{P,0}\Delta P_{it} + \alpha_{Y,0}\Delta Y_{it} + \alpha_{Z,0}\Delta Z_{it} \\ &\quad + \alpha_{PZ,0}\Delta P_{it}Z_{it} + \alpha_{PZ,1}\Delta P_{it-1}Z_{it-1} \\ &\quad + \hat{\gamma}(Q_{it-1} - \theta_P P_{it-1} - \theta_Y Y_{it-1} - \theta_Z Z_{it-1} - \theta_{PZ}P_{it-1}Z_{it-1}) + u_{it}.\end{aligned} \tag{7}$$

$$\begin{aligned}\Delta P_{it} &= \beta_0 + \beta_{Q,0}\Delta Q_{it} + \beta_{W^1,0}\Delta W^1_{it} + \sum_{s=0}^{7}\beta_{W^2,s}\Delta W^2_{it-s} + \sum_{s=0}^{2}\lambda_s\Delta\tilde{Q}_{it-s} \\ &\quad + \hat{\psi}\left(P_{it-1} - \xi_Q Q_{it-1} - \xi_{W^1}W^1_{it-1} - \xi_{W^2}W^2_{it-1} - \Lambda\tilde{Q}_{it-1}\right) + \eta_{it},\end{aligned} \tag{8}$$

where $\tilde{Q}_{it} = Q_{it}/(\theta_P + \theta_{PZ}Z_{it})$.

4. Empirical results

4.1. Empirical results of the SBLM

For the supply and demand specifications, Q and P are interrelated with each other. To identify the coefficients of the two variables, this paper uses the conventional two-stage seemingly unrelated regression (SUR) method, whose estimates can easily verify the consistency. The estimation results of the SBLM are shown in Table 3.

In the demand function, all of the coefficient estimates are significantly different from zero at even the 1% level of significance, whereas the value of the adjusted R^2 is not high. As expected, the coefficient of product price (P) "interest and fee income from loans/total loans & leases, gross," is negative, which is consistent with the law of demand. The price of the other product (Z), represented by the interest rate on three-month Treasury Bills, is found to have a negative association

Table 2
Panel unit root tests.

Method variable		Levin, Lin & Chu t-stat[a]	Breitung t-stat[a]	Hadri Z-stat[b]	Im, Pesaran & Shin W-stat[c]	ADF Fisher χ^{2c}	PP Fisher χ^{2c}
Q	Statistic	9.9184	3.2058	97.8381	30.0986	315.059	316.513
	P-value	1.0000	0.9993	0.0000	1.0000	1.0000	1.0000
	Observations	21,152	20,841	21,632	21,152	21,152	21,294
P	Statistic	41.1336	14.2983	33.3515	−8.6365	879.483	13900.3
	P-value	1.0000	1.0000	0.0000	0.0000	0.0000	0.0000
	Observations	19,835	19,479	21,624	19,835	19,835	21,280
Y	Statistic	28.7189	5.1166	104.525	54.2221	2.4088	7.5777
	P-value	1.0000	1.0000	0.0000	1.0000	1.0000	1.0000
	Observations	21,294	20,956	21,632	21,294	21,294	21,294
Z	Statistic	−19.8053	13.2254	39.7182	−23.4654	1619.35	1300.18
	P-value	0.0000	1.0000	0.0000	0.0000	0.0000	0.0000
	Observations	20,956	20,618	21,632	20,956	20,956	20,956
W^1	Statistic	5.2233	19.2432	104.232	31.7357	33.4229	22.7947
	P-value	1.0000	1.0000	0.0000	1.0000	1.0000	1.0000
	Observations	21,294	20,956	21,632	21,294	21,294	21,294
W^2	Statistic	6.9611	7.6521	78.2898	−20.7937	1452.74	8594.84
	P-value	1.0000	1.0000	0.0000	0.0000	0.0000	0.0000
	Observations	19,879	19,541	21,624	19,879	19,879	21,286
PZ	Statistic	−6.3943	11.2912	56.1820	−25.3147	1734.21	6548.88
	P-value	0.0000	1.0000	0.0000	0.0000	0.0000	0.0000
	Observations	19,913	19,575	21,624	19,913	19,913	21,280

[a] Denotes Null hypothesis: unit root (assumes common unit root process).
[b] Denotes Null hypothesis: no unit root (assumes common unit root process) (Hadri, 2000).
[c] Denotes Null hypothesis: unit root (assumes individual unit root process).

Table 3
The estimation results of the SBLM using SUR.

Demand function		Supply function	
Coefficient	Estimate	Coefficient	Estimate
α_0	−9.0069* (0.3645)	β_0	−0.0553* (0.0055)
α_P	−6.0490* (0.3460)	β_Q	−0.0022* (0.0002)
α_Y	3.7684* (0.0898)	β_{W^1}	0.0286* (0.0018)
α_Z	−0.0388* (0.0037)	β_{W^2}	1.7857* (0.0086)
α_{PZ}	0.6689* (0.0407)	λ	0.0032* (0.0004)
R^2	0.1389	R^2	0.7041
$\bar{R}^2$	0.1388	$\bar{R}^2$	0.7040
D-W statistic	0.0653	D-W statistic	1.1412

The standard errors are included in parentheses.
* Indicates that the parameter estimate is significantly different from zero at the 1% level.

with the demand for loans. An increase in real GDP (Y) incurs an increase in the demand for loans. This is in conformity with the results obtained by Shaffer (1989) and Móré and Nagy (2004).

As far as the supply function is concerned, all of the coefficient estimates are significantly different from zero even at the 1% level and the value of the adjusted R^2 is quite high. According to Table 3, a negative output coefficient (β_Q) shows that an increase in lending supply will lead to a reduced interest rate on loans. A rise in wages or in the funding price raises the production costs, and consequently the output price has to be increased to at least cover the incremental costs. Finally, since the market power indicator λ is significant but almost zero, the market power of the sample banks is not well-pronounced. In other words, the U.S. commercial banking industry appears to be quite close to perfect competition.

4.2. Empirical results of the DBLEC

To validate the dynamic demand function, we first calculate several long-run elasticity measures. The long-run own-price elasticity is defined as $\varepsilon_P = \left[\theta_P + \theta_{PZ}\bar{Z}\right] \cdot \left[\bar{P}/\bar{Q}\right]$, the income elasticity as $\varepsilon_Y = \theta_Y\left[\bar{Y}/\bar{Q}\right]$, and the cross-price elasticity as $\varepsilon_Z = \left[\theta_Z + \theta_{PZ}\bar{P}\right] \cdot \left[\bar{Z}/\bar{Q}\right]$, where $\bar{Z}$, $\bar{P}$, $\bar{Q}$, and $\bar{Y}$ denote the sample means of the corresponding variables. Using the parameter estimates shown in Table 4 together with the data, we compute the various elasticity measures. The long-run own-price elasticity (ε_P) is equal to −1.1731, again implying a downward-sloping demand curve, which exceeds the same measure yielded by the SBLM in absolute value terms. The income elasticity (ε_Y) of 10.2156, similar to the result obtained using the SBLM, suggests that output loans is a luxury good. The long-run cross-price elasticity (ε_Z) equals 0.7047, meaning that the three-month Treasury Bills is a substitute for the loans.

Although the estimated error correction coefficient ($\hat{\gamma} = -0.0024$) is significant and has the expected sign, its absolute value is nearly zero. This implies that the adjustment speed of the demand for loans market towards its long-run steady-state is slow. Such a tardy adjustment speed may be attributed either to the fact that the deviation from the steady-state may not be far away, to the fact that the customers' habit is highly persistent, requiring a much longer time to be altered, or to the fact that, given the use of quarterly data, the market just does not have enough time to respond to the excessive demand.

The empirical results of the dynamic supply function reveal that the short-run market power indicators λ_s, $s = 0, 1, 2$, are statistically significant, while their absolute values are small. This implies that the U.S. commercial banking market is perfectly competitive in the short run, similar to the SBLM results. However, some commercial banks are able to exercise a certain degree of market power in the long run, because the value of the long-run market power indicator Λ is equal to 0.0931 and is significantly different from zero at the 1% level. Furthermore, as

Table 4
The estimation results of DBLEC with two-stage SUR.

Demand function		Supply function	
Coefficient	Estimate	Coefficient	Estimate
α_0	−0.1107*(0.0215)	β_0	0.0942*(0.0052)
$\alpha_{Q,1}$	0.0550*(0.0069)	$\beta_{Q,0}$	−0.0395*(0.0018)
$\alpha_{P,0}$	−0.1316*(0.0288)	$\beta_{Q,1}$	−0.0205*(0.0018)
$\alpha_{Y,0}$	0.9239*(0.1380)	$\beta_{W_0^1}$	−0.1275*(0.0110)
$\alpha_{Z,0}$	0.0004 (0.0008)	$\beta_{W_0^2}$	0.9414*(0.0151)
$\alpha_{PZ,0}$	−0.0063 (0.0062)	$\beta_{W_1^2}$	0.0776*(0.0174)
$\alpha_{PZ,1}$	0.0148*(0.0025)	$\beta_{W_2^2}$	−0.1272*(0.0173)
$\hat{\gamma}$	−0.0024*(0.0004)	$\beta_{W_3^2}$	−0.2456*(0.0176)
Long-run parameters			
θ_P	−85.1968*(20.0591)	$\beta_{W_4^2}$	0.4559*(0.0179)
θ_Y	14.3340*(2.8020)	$\beta_{W_5^2}$	−0.0202(0.0167)
θ_Z	1.4790*(0.3078)	$\beta_{W_6^2}$	−0.1760*(0.0153)
θ_{PZ}	−9.9080*(4.0535)	$\beta_{W_7^2}$	−0.3877*(0.0136)
Long-run elasticity		λ_0	0.0592*(0.0081)
ε_P	−1.1731	λ_1	−0.0999*(0.0079)
ε_Y	10.2156	λ_2	−0.0949*(0.0075)
ε_Z	0.7047	$\hat{\psi}$	−0.6976*(0.0068)
		Long-run parameters	
R^2	0.0216	ξ_Q	0.0037*(0.0003)
$\bar{R}^2$	0.0210	ξ_{W^1}	−0.0226*(0.0023)
		ξ_{W^2}	0.3810*(0.0196)
		Λ	0.0931*(0.0039)
		R^2	0.9431
		$\bar{R}^2$	0.9430

The standard errors are included in parentheses.
* Indicates that the parameter estimate is significantly different from zero at the 1% level.

shown in Table 4, the significantly negative coefficients of lending supply to the interest rate on loans also confirm the conclusion above: some commercial banks have market power in the long run. It is noticeable that the supply-side speed of adjustment ($\left|\hat{\psi}\right| = 0.6976$) towards its long-run equilibrium through the error correction term is quite rapid. To be viable in such a fiercely competitive environment, each bank has to quickly respond to any market disequilibrium that is possibly caused by demand and supply shocks.

5. Concluding remarks

Prior to the 1990s, generally accepted restrictions on inter-state banking ensured that large banks could not become nationwide oligopolists. However, starting in the 1990s, the competitive impacts of the government's liberalization of the banking sector began to be more seriously questioned. To address this issue, in this study we have estimated models of the SBLM and DBLEC using data for listed commercial banks operating in the U.S.

Although the estimated SBLM market power indicator is statistically significant, its value is quite close to zero, indicating that the industry is nearly perfectly competitive. This is consistent with the results obtained by Shaffer (1989) who studied the same industry over the period 1941 to 1983. In addition, an increase in the interest rate on bank loans and in the interest rate on three-month Treasury Bills is found to reduce the demand for bank loans, while an increase in real GDP by contrast is found to raise the demand for bank loans.

The empirical results from the DBLEC demonstrate that the short-run market power indicator is significant but again close to zero, indicating that the banking industry is nearly perfectly competitive in the short run. However, as the long-run market power indicator is significant and deviates from zero, some degree of market power does exist in the industry in the long run.

This study mainly contributes to providing a more comprehensive analysis of market power in the U.S. commercial banking industry after 1990, while the SBLM and DBLEC approaches presented in this paper exhibit some limitations. Thus, we leave for future work the question of whether dynamic competition *among* individual banks with the Panzar-Rosse model could serve as a useful step in the right direction.

Appendix A. Pairwise colinearity of independent variables

	Output (Q)	Price (P)	Income (Y)	Other price (Z)	Price of labor (W^1)	Price of funds (W^2)
Output(Q)	1.0000					
Price(P)	−0.0580*(0.0000)	1.0000				
Income (Y)	0.1041* (0.0000)	0.2410*(0.0000)	1.0000			
Other price (Z)	−0.0620*(0.0000)	0.2175*(0.0000)	0.5667*(0.0000)	1 .0000		
Price of labor (W^1)	0.1047*(0.0000)	−0.2607*(0.0000)	0.9889*(0.0000)	0.6308*(0.0000)	1.0000	
Price of funds (W^2)	−0.0507*(0.0000)	0.8238*(0.0000)	−0.4103*(0.0000)	0.3952*(0.0000)	−0.4188*(0.0000)	1.0000

* Indicates that the null hypothesis is rejected at the 1% level.
The P-values are shown in parentheses.

Appendix B. The relationship between ADL model and ECM

For the following ADL(p,q) model, $Y_t = \alpha_0 + \beta_0 X_t + \beta_1 X_{t-1} + \beta_2 X_{t-2} + \ldots + \beta_p X_{t-p} + \gamma_1 Y_{t-1} + \gamma_2 Y_{t-2} + \ldots + \gamma_q Y_{t-q} + u_t$, we have that $Y_t = \alpha_0 + \sum_{i=0}^{p} \beta_i L^i X_t + \sum_{j=1}^{q} \gamma_j L^j Y_t + u_t$ (L indicates lag operator) and then

$$\left(1 - \sum_{j=1}^{q} \gamma_j L^j\right) Y_t = \alpha_0 + \sum_{i=0}^{p} \beta_i L^i X_t + u_t. \tag{A1}$$

Since the left-hand side of Eq. (A1) equals to $\left(1 - \sum_{j=1}^{q} \gamma_j\right) Y_t + \sum_{j=1}^{q} \gamma_j \Delta Y_t + \sum_{j=2}^{q} \gamma_j \Delta Y_{t-1} + \sum_{j=3}^{q} \gamma_j \Delta Y_{t-2} + \ldots + \gamma_q \Delta Y_{t-q+1}$; that is, $\left(1 - \sum_{j=1}^{q} \gamma_j L^j\right) Y_t = (1 - \Gamma_1) Y_t + \sum_{j=1}^{q} \Gamma_j \Delta Y_{t-j+1}$. Thus, we have that

$$\sum_{j=1}^{q} \gamma_j L^j Y_t = \Gamma_1 Y_{t-1} - \sum_{j=2}^{q} \Gamma_j \Delta Y_{t-j+1}. \tag{A2}$$

For X_t in the right-hand side of Eq. (A1), $\sum_{i=0}^{p} \beta_i L^i X_t = \sum_{i=0}^{p} \beta_i X_t - \sum_{i=1}^{p} \beta_i \Delta X_t - \sum_{i=2}^{p} \beta_i \Delta X_{t-1} - \sum_{i=3}^{p} \beta_i \Delta X_{t-2} - \ldots - \beta_p \Delta X_{t-p+1}$, where $\Gamma_1 \equiv \sum_{j=1}^{q} \gamma_j$; $\Gamma_2 \equiv \sum_{j=2}^{q} \gamma_j$; $\Gamma_q \equiv \gamma_q$ and $B_0 \equiv \sum_{i=0}^{p} \beta_i$; $B_1 \equiv \sum_{i=1}^{p} \beta_i$; …; $B_p \equiv \beta_p$,

$$\sum_{i=0}^{p} \beta_i L^i X_t = B_0 X_t - \sum_{i=1}^{p} B_i \Delta X_{t-i+1} \tag{A3}$$

exists.

For Eq. (A3), we further have that

$$\begin{aligned} \sum_{i=0}^{p} \beta_i L^i X_t &= B_0 X_t - B_1 \Delta X_t - B_2 \Delta X_{t-1} - B_3 \Delta X_{t-2} - \ldots - B_p \Delta X_{t-p+1} \\ &= (B_0 - B_1) \Delta X_t + (B_0 - B_1) X_{t-1} + B_1 X_{t-1} - B_2 \Delta X_{t-1} - B_3 \Delta X_{t-2} - \ldots - B_p \Delta X_{t-p+1} \\ &= \left(\sum_{i=0}^{p} \beta_i - \sum_{i=0}^{p} \beta_i\right) \Delta X_t + B_0 X_{t-1} - \sum_{i=2}^{p} B_i \Delta X_{t-i+1} = B_0 X_{t-1} + \beta_0 \Delta X_t - \sum_{i=2}^{p} B_i \Delta X_{t-i+1}. \end{aligned} \tag{A4}$$

Substituting Eqs. (A2) and (A4) into Eq. (A1) gives that:

$$Y_t = \alpha_0 + B_0 X_{t-1} + \beta_0 \Delta X_t - \sum_{i=2}^{p} B_i \Delta X_{t-i+1} + \Gamma_1 Y_{t-1} - \sum_{j=2}^{q} \Gamma_j \Delta Y_{t-j+1} + u_t. \tag{A5}$$

That is,

$$\begin{aligned} \Delta Y_t &= \alpha_0 + B_0 X_{t-1} + \beta_0 \Delta X_t - \sum_{i=2}^{p} B_i \Delta X_{t-i+1} - (1 - \Gamma_1) Y_{t-1} - \sum_{j=2}^{q} \Gamma_j \Delta Y_{t-j+1} + u_t \\ &= (1 - \Gamma_1) \cdot \frac{\alpha_0}{1 - \Gamma_1} + (1 - \Gamma_1) \frac{B_0 X_{t-1}}{1 - \Gamma_1} + \beta_0 \Delta X_t - \sum_{i=2}^{p} B_i \Delta X_{t-i+1} - (1 - \Gamma_1) Y_{t-1} - \sum_{j=2}^{q} \Gamma_j \Delta Y_{t-j+1} + u_t \\ &= \beta_0 \Delta X_t - \sum_{i=2}^{p} B_i \Delta X_{t-i+1} - \sum_{j=2}^{q} \Gamma_j \Delta Y_{t-j+1} - (1 - \Gamma_1) \cdot \left[Y_{t-1} - \frac{\alpha_0}{1 - \Gamma_1} - \frac{B_0}{1 - \Gamma_1} \cdot X_{t-1}\right] + u_t = \beta_o \Delta X_t - \sum_{i=2}^{p} B_i \Delta X_{t-i+1} - \sum_{j=2}^{q} \Gamma_j \Delta Y_{t-j+1} - (1 - \Gamma_1) \xi_{t-1} + u_t. \end{aligned} \tag{A6}$$

The ADL model of Eq. (A1) can be represented by ECM in (A6).

Appendix C. Panel cointegration test

Series: Q, P, Y, Z, PZ				
1. Pedroni residual cointegration test: null hypothesis: no cointegration				
a. Alternative hypothesis: common AR coefs. (within-dimension)				
			Weighted	
	Statistic	*P*-value	Statistic	*P*-value
Panel v-statistic	−7.2129	0.0000	−6.9265	0.0000
Panel rho-statistic	17.5621	0.0000	17.4070	0.0000
Panel PP-statistic	12.9288	0.0000	12.9860	0.0000
Panel ADF-statistic	17.4318	0.0000	17.1725	0.0000
b. Alternative hypothesis: individual AR coefs. (between-dimension)				
	Statistic	*P*-value		
Group rho-statistic	24.2084	0.0000		
Group PP-statistic	18.4396	0.0000		
Group ADF-statistic	32.4184	0.0000		
2. Kao residual cointegration test null hypothesis: no cointegration				
ADF	t-statistic	−3.2615	*P*-value	0.0006
Residual variance			0.0014	
HAC variance			0.0019	
Series: P, Q, W1, W2, QSTRD				
1. Pedroni residual cointegration test: null hypothesis: no cointegration				
a. Alternative hypothesis: common AR coefs. (within-dimension)				
			Weighted	
	Statistic	*P*-value	Statistic	*P*-value
Panel v-statistic	12.1758	0.0000	1.7249	0.0901
Panel rho-statistic	−33.4694	0.0000	−31.939	0.0000
Panel PP-statistic	−51.4629	0.0000	−48.3767	0.0000
Panel ADF-statistic	−27.7744	0.0000	−25.0997	0.0000
b. Alternative hypothesis: individual AR coefs. (between-dimension)				
	Statistic	*P*-value		
Group rho-statistic	−29.6278	0.0000		
Group PP-statistic	−56.2915	0.0000		
Group ADF-statistic	−26.3542	0.0000		
2. Kao residual cointegration test null hypothesis: no cointegration				
ADF	*t*-statistic	−19.9759	*P*-value	0.0000
Residual variance			0.0002	
HAC variance			4.19E−05	

References

Angelini, P., Cetorelli, N., 2000. Bank competition and regulatory reform: the case of the Italian banking industry. Federal Reserve Bank of Chicago. WPS No. 99–32.

Appelbaum, E., 1982. The estimation of the degree of oligopoly power. Journal of Econometrics 19 (23), 287–299.

Bask, M., Lundgren, J., Rudholm, N., 2011. Market power in the expanding Nordic power market. Applied Economics 43 (9), 1035–1043.

Berger, A.N., Demirgüç-Kunt, A., Levine, R., Haubrich, J.G., 2004. Bank concentration and competition: an evolution in the making. Journal of Money, Credit and Banking 36 (3), 433–451.

Bikker, J.A., 2003. Testing for imperfect competition on EU deposit and loan markets with Bresnahan's market power model. De Nederlandsche Bank, Research Series Supervision No. 52.

Bikker, J.A., Haaf, K., 2000. Competition, concentration and their relationship: an empirical analysis of the banking industry. De Nederlandsche Bank, Research Series Supervision No. 30.

Bikker, J.A., Haaf, K., 2002. Competition, concentration and their relationship: an empirical analysis of the banking industry. Journal of Banking and Finance 26 (11), 2191–2214.

Bresnahan, T.F., 1982. The oligopoly solution concept is identified. Economics Letters 10 (1–2), 87–92.

Bresnahan, T.F., 1989. Empirical studies in industries with market power. In: Schmalensee, R., Willig, R. (Eds.), Handbook of Industry Organization 2. North Holland, Amsterdam, pp. 1011–1057.

Buschena, D.E., Perloff, J.M., 1991. The creation of dominant firm market power in the coconut oil export market. American Journal of Agricultural Economics 73 (4), 1000–1008.

Canhoto, A., 2004. Portuguese banking: a structural model of competition in the deposits market. Review of Financial Economics 13, 41–63.

Cetorelli, N., 2004. Real effects of bank competition. Journal of Money, Credit and Banking 36 (3), 543–558.

Chintrakarn, P., Jindapon, P., 2012. A dynamic analysis of market power in the U.S. mortgage lending market. European Journal of Science Research 69 (2), 218–223.

Demirgüç-Kunt, A., Laeven, L., Levine, R., 2004. Regulations, market structure, institution, and the cost of financial intermediation. Journal of Money, Credit and Banking 36 (3), 593–622.

Deodhar, S.Y., Sheldon, I.M., 1997. Market power in the world market for soymeal exports. Journal of Agricultural and Resource Economics 22 (1), 78–86.

Engle, R.F., Granger, C.W.J., 1987. Cointegration and error correction: representation, estimation and testing. Econometrica 55 (2), 251–276.

Favero, C.A., Papi, L., 1995. Technical efficiency and scale efficiency in the Italian banking sector: a non-parametric approach. Applied Economics 27 (4), 385–395.

Frank, A., Steen, F., 2006. When anti-dumping measures lead to increased market power: a case study of the European salmon market. Center for Economic Policy Research, DP No. 5781.

Gollop, F.M., Roberts, M.J., 1979. Firm interdependence in oligopolistic markets. Journal of Econometrics 10 (3), 313–331.

Gunalp, B., Celik, T., 2006. Competition in Turkish banking industry. Applied Economics 38 (11), 1335–1342.

Hadri, K., 2000. Testing for stationarity in heterogeneous panel data. Econometrics Journal 3 (2), 148–161.

Hatirli, S.A., Jones, E., Aktas, A.R., 2003. Measuring the market power of banana import market in Turkey. Journal of Agriculture and Forestry 27 (3), 367–373.

Hjalmarsson, E., 2000. Nord pool: a power market without market power. Department of Economics, Göteborg University. WPS No. 28.

Ho, C.Y., 2010. Deregulation, competition and consumer welfare in a banking market: evidence from Hong Kong. Journal of Regulatory Economics 37, 70–97.

Kao, C., 1999. Spurious regression and residual-based tests for cointegration in panel data. Journal of Econometrics 90 (1), 1–44.

Lau, L.J., 1982. On identifying the degree of competitiveness from industry price and output data. Economics Letters 10 (1–2), 93–99.

Móré, C., Nagy, M., 2004. Competition in the Hungarian banking market. Magyar Nemzeti Bank, Hungary. WP No. 2004/9.

Nathan, E.H., Neave, A., 1989. Competition and contestability in Canada's financial system: empirical results. Canadian Journal of Economics 22 (3), 576–594.

Neven, D., Róller, L.H., 1999. An aggregate structural model of competition in the European banking industry. International Journal of Industrial Organization 17, 1059–1074.

Northcott, C.A., 2004. Competition in banking: a review of the literature. Bank of Canada, Ottawa. WP No. 04–24.

Noulas, A.G., 1997. Productivity growth in the Hellenic banking industry: state versus private banks. Applied Financial Economics 7 (3), 223–228.

Panzar, J.C., Rosse, J.N., 1987. Testing for 'monopoly' equilibrium. Journal of Industrial Economics 35 (4), 443–456.

Pedroni, P., 2000. Full modified OLS for heterogeneous cointegration panels. Advances in Econometrics 15, 93–130.

Pedroni, P., 2004. Panel cointegration: asymptotic and finite sample properties of pooled time series tests with an application of the PPP hypothesis. Econometric Theory 20 (3), 597–625.

Shaffer, S., 1989. Competition in the U.S. banking industry. Economics Letters 29 (4), 321–323.

Shaffer, S., 1993. A test of competition in Canadian banking. Journal of Money, Credit, and Banking 25 (1), 49–61.

Shaffer, S., Disalvo, J., 1994. Conduct in banking duopoly. Journal of Banking and Finance 18 (6), 1063–1082.

Steen, F., Salvanes, K.G., 1999. Testing for market power using a dynamic oligopoly model. International Journal of Industrial Organization 17 (2), 147–177.

Suominen, M., 1994. Measuring competition in banking: a two-product model. Scandinavian Journal of Economics 96 (1), 95–110.

Toolseman, L.A., 2002. Competition in the Dutch consumer credit market. Journal of Banking and Finance 26 (11), 151–165.

Zardkoohi, A., Fraser, D.R., 1998. Geographical deregulation and competition in U.S. banking markets. Financial Review 33 (2), 85–98.