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E. P. Ribeiro, K. Castor[1]

Inferring bank-to-bank competition from dynamic time series analysis of price correlations

Inferring bank to bank rivalry and competition generally requires the estimation of a full demand model, with high data requirements, unavailable to most researchers. We suggest dynamic time series analysis of price correlations to infer about bank to bank competition, taking into account the well-known criticisms to price correlations for delimiting relevant markets. The method is applied for credit markets in Brazil, where bank monthly loan interest rates time series are available. We conclude that there is little rivalry between large banks in most of the credit markets studied.

Keywords: bank competition; rivalry; mergers; Brazil.

JEL classification: D22; C20.

Introduction

Bank competition is an important issue in many jurisdictions, e.g. (OECD, 2010), with central place for the interplay between financial stability and competition. In the case of Brazil, where very high interest rates on loans coexist with marked concentration, banking competition is subject to a heated debate.

The literature on banking competition across countries often attempts to describe the degree of competition using the standard market structure classification (perfect competition, monopolistic competition, oligopoly, monopoly), applying well known economics methods, such as (Bresnahan, 1982; Lau, 1982; Panzar, Rosse, 1987) and more recent developments, as (Boone, 2008). An application for Russia is (Mamonov, 2015). For Brazil, the main conclusion of Barbosa et al. (2015), that use multiple methods, is that retail banking in the Brazil is neither in the extreme of perfect competition nor in the other extreme of tacit coordination structure.

While these papers try to summarize competition levels in a single metric, Competition Authorities face other challenges. The focus of their analysis is on changes in competition levels from a merger, given a market structure and characteristics and effects from the merger itself. In more detail, when evaluating a merger in highly concentrated markets, it is important to learn

[1] **Ribeiro, Eduardo Pontual** — Federal University of Rio de Janeiro, Economics Institute, Rio de Janeiro, Brazil; eribeiro@ie.ufrj.br.

Castor, Kamaiaji — Federal University of Rio de Janeiro, Economics Institute, Rio de Janeiro, Brazil; kamaiaji.castor@ufrj.br.

whether the merging parties still face strong competition from non-merging firms, or whether the merging parties are the closes competitors. This firm-to-firm competition analysis does not recommend the use of the above mentioned measures in merger analysis.

Often the information of such degree of rivalry is qualitative, such as the opinion of market players, the opinion on customer switching costs or indirect evidence such as idle capacity of the firms in the market. Alternative, the influential US merger guidelines (FTC, 2010) suggest the estimation of diversion ratios between firms, that require detailed consumer data or the estimation of full-fledged demand models, with high data demands. Most merger analysis outside Europe and the US do not employ such methods due to practical and time constraints.

This paper proposes the use of price correlations to provide useful information on firm-to-firm rivalry and whether merging parties would be close competitors or not. Price correlations have been extensively used in merger analysis for relevant market definition ((Donath, 2009) for the EU experience, and (Katsoulacos et al., 2014), for an example). We propose the use of dynamically complete price correlations to learn about competition between firms within a relevant market. Based on a reaction function specification in a differentiated product market, one can evaluate whether merging banks react to price (interest rate) changes at each other and from other banks.

We adapt the price correlation methodology to evaluate relevant markets for rivalry measurement. Although price correlation methods are quite popular their weaknesses are well known. Davis, Garcés (2009) show that given common cost and demand shocks, prices of different firms will co-move even if their products are not substitutes among themselves and the firms do not compete with each other. Hence the method may provide strong evidence on whether two firms *do not* compete (prices do not react to each other), while it provides weak evidence, in many cases whether they do compete.

A difficulty in estimating reaction functions includes working with simultaneous time series of many banks, leading to a very large number of coefficients in a standard VAR/VECM analysis for price correlations. We propose analyzing bank pairs to keep models manageable. While this introduces a secondary source of lack of identification of price reactions, due to chain of substitution effects[2], the null of no competition between firms is not influenced by omitted variables.

The use of price correlations to evaluate the price reactions requires much less information than more detailed indicators of firm competition such as diversion ratios and may be used as a simple filter for further competition study, given its correct size using a hypothesis test terminology (it correctly identifies no price reaction between two firms if this reaction does not exist).

With non-stationary time series, Pesaran (2007) suggests that the joint non-stationarity of a group of variables may be evaluated with the proportion of rejections of unit root tests found for each pair of series, instead of evaluating a model with all variables simultaneously. We follow this argument to evaluate price reactions and explore bank differences in error correction duration as in (Iregui, Otero, 2013). While Iregui and Otero use the log difference between prices of two firms (a log price ratio) as a possibly non-stationary variable, we recognize that a log price ratio is a cointegration (1; –1) vector and prefer to allow non-fixed cointegra-

[2] Chain of substitution effects are mentioned in the EU Relevant Market definition guideline (1997 Commission Notice on the definition of relevant market for the purposes of Community competition law, par. 57). For example, there is no price reaction from firm A to firm C, conditional of firm B price, and firm B price reacts to price A and price C movements.

tion vector for all price pairs. We then use the analysis for cointegration of general cointegration vectors. Note that the argument for price cointegration by Iregui, Otero (2013) is the Law of one price, we motivate the use of the price correlations under a reaction function for differentiated product competition.

Advancing the results we conclude that firm overdraft credit banks of different size compete less intensively and on individual overdraft credit interest rates for banks of differing size compete more intensely.

In summary, this paper provides new evidence on competition among banks in the banking sector in Brazil, filling a gap in the literature and in merger analysis carried by antitrust authorities. The method, based on price correlations, can provide information on the lack of close competition among banks in specific markets. Our central point is that price correlations, given their well-known weaknesses for revealing price competition from extensive false positives, can nevertheless be used to provide relevant information on the lack of competition among firms.

The paper is organized as follows: the next section provides an overview of the use of price correlations to competition and economic problems leading to the methodology used in this paper; the following section describes the data and presents the empirical results. The last section contains concluding comments.

1. Price correlations as price competition measure

There is an extensive literature on the use and misuse of price correlations for relevant market delineation (Haldrup, 2003; Werden, Froeb, 1993; Coe, Krause, 2008; Davis, Garcés, 2009). Davis and Garces show that price correlations may be justified under a reaction function model of competition (such as Bertrand). Yet, price correlations between markets of firms would be significant when common costs or demand shocks are available, even if markets or firms are not competitively related. This conclusion does not depend on whether prices are cointegrated or not, as cointegration is a time series property associated with effective correlation of nonstationary time series. The bias from identifying product competition from price correlation is one of omitted variable bias, namely, cost and demand shocks[3].

Coe, Krause (2008) on a more positive perspective suggest that, in large samples, and in the absence of common shocks, price correlations do reflect product substitution. But they are wary of time series methods to identify these effects. Nevertheless, price correlation tests continue to be used and are mentioned in the 2016 Horizontal Merger Guidelines by the Brazilian Competition Authority (CADE, 2016) as an empirical technique.

We argue that given the criticism, nevertheless price correlations provide effective information on whether products or firms are *not* substitutes or close competitors. In other words, if price correlation between products is not significant, one has constructive evidence that the products do not compete. On the other hand, significant correlation does not identify product substitution due to the potential omitted variable bias from common demand and cost shocks.

[3] Under a hypothetical monopolist framework for market delineation, a second, more important criticism of price correlations as a measure to delimit markets is based on the fact that while price correlations reflect possible product substitution, a Critical Loss Analysis requires substitution and margins interplay to indicate profitable price increases from the hypothetical monopolist. (Werden, Froeb, 1993).

While the price correlations are extensively discussed as a relevant market delineation tool, we take the reaction function framework and use it to provide evidence on firm competition. Reaction functions, under a given competition model, such as Bertrand, would require the estimation of a large system of equations (e.g., (Moita, Silva, 2014)). These systems in a time series settings (VAR/VECM models) may generate a very large number of parameters to estimate. We follow Iregui, Otero (2013) that argue that bivariate comparisons may consistently be used to evaluate non-correlations, even in a time series settings, based on Pesaran (2007) result that the non-stationarity of a panel of ratios may be evaluated using pairwise comparisons. The size of the test is based on the proportion of bivariate test rejections. As a ratio between two variables is a specific cointegration vector, we apply this to general cointegration tests, such as Johansen or an ECM test (see (Enders, 2014) for a description of these tests).

Different cointegration tests have different data generating model assumptions. We use two cointegration tests, Johansen and an ECM test. Johansen´s test does not require weak exogeneity of any series, but requires a correctly specified system of equations. The ECM test requires the specification of a single equation, but requires weak exogeneity of the explanatory variable for consistency. Their use is complementary for robustness. For a given bank $i=1,..,I$ with credit product price (ln (1+interest rate)) $p_{i,t}$ the error correction models are of the form

$$\Delta p_{i,t} = \theta_0^i + \alpha_i^j p_{i,t-1} + \theta_i^j p_{j,t-1} + \sum_{k=1}^{n} \gamma_k^i \Delta p_{i,t-k} + \sum_{k=1}^{m} \delta_k^i \Delta p_{j,t-k} + u_{i,t}, \qquad (1)$$

where $p_{j,t}$ is the price series of another product or firm. The model is estimated for every pair of banks $i \neq j$.We use the empirical version of the ECM suggested by Ericsson and MacKinnon (2002) for testing cointegration in an ECM model, namely, we do not restrict the cointegration vector and adjustment coefficient $\theta_0^i + \alpha_i^j p_{i,t-1} + \theta_i^j p_{j,t-1}$ as $\alpha_i^j (p_{i,t-1} - \beta_i^j p_{j,t-1} - \beta_{0i}^j)$.

We explore a measure of competition proximity between two banks, namely the speed of adjustment coefficient in an error correction model as in (Iregui, Otero, 2013), calculated in half life form $\ln(1/2)/\ln(1-a_i^j)$. Two banks would be considered close competitors if the adjustment coefficient for the pairwise error correction models suggests a short time to adjust to interest rate differences changes. Two banks will not be considered competitors if they do not cointegrate. This criterion implicitly assumes that that competition requires relative price levels to matter, not just price growth correlation. Note that effects may be asymmetric, namely, the speed of adjustment coefficient for bank *i* when it responds to changes in bank *j* interest rate changes is not imposed to be symmetric to a model where bank *j* interest rate is the dependent variable and bank *i* is the explanatory variable.

We calculate speed of adjustment coefficients for each pair of banks with long enough continuous time series information on credit product interest rates.

Once the speed of adjustment coefficients is calculated we explore patterns of response across bank types, running a regression of these coefficients to bank characteristics as in (Iregui, Otero, 2013). This allows us to provide general comments on the pattern of competition between banks. Namely, we regress the speed of adjustment measured as median lag on bank size differences, and whether the banks involved in the pairwise comparison are state owned and whether any is a foreign bank. In detail, we estimate

$$median\_lag_i^j = b_0 + b_1 \left(|\ln A_i - \ln A_j| \right) + b_2 State_{i,j} + b_3 Foreign_{i,j} + u_{ij}, \qquad (2)$$

where $median\_lag_i^j = \ln(1/2)/\ln(1+a_i^j)$ (see, e.g., (Harvey, 1990)), $|\ln A_i - \ln A_j|$ measures the bank size (total assets) differences between bank *i* and bank *j*, $State_{i,j}$ — a dummy that measures if either bank *i* or bank *j* is a state owned bank and $Foreign_{i,j}$ is one if either bank *i* or bank *j* is a multinational bank. Equation (2) results can provide evidence on whether banks of different size influence each other or not (or alternatively whether there is market segmentation across bank sizes). It can inform on the common knowledge that public banks may not compete directly with private banks, as their profit maximizing behavior may differ. Foreign banks, on the other hand may compete more intensively than Brazilian banks as they have a different institutional control setting. We use the median lag it provides a transformation of the negative speed of adjustment coefficient. The dependent variable becomes positive and informs about the closeness of competition. If the absolute value of the regression coefficient a_i^j, is small the speed of adjustment low, and the median lag variable will be high. Importantly, as the dependent variable is an estimated coefficient, each observation is weighted by the standard error of the speed adjustment coefficient in each regression.

2. Data description and empirical results

The Central Bank of Brazil publishes weekly average interest rates of new credit contracts in each bank for each product. We consider four types of credit lines, two for individuals and two for firms. The choice of credit lines is restricted by compatibility over time of the price series, as there was a methodological change in the price series in 2012. New credit lines were registered by the authority form 2012. For the selected credit lines, the measurement did not change from 2009. The series are, for firm credit, receivables credit and firm overdraft. For individuals, vehicle acquisition and personal credit overdraft. The number of banks in each product line with sufficient observations for estimation (5 years) range from 9 to 13, including the largest banks and some medium banks. The list of banks in each case will appear in the tables below. The last week of the month interest rates are used in the analysis.

We start the analysis evaluating the joint non-stationarity of the interest rate time series, using panel unit root tests. The Im–Pesaran–Shin test and the Levin–Lin–Chu test have different alternative hypotheses. The later test does not allow for heterogeneity in autoregressive coefficients across banks. The results in table 1 suggest that receivables and vehicle acquisition may be considered jointly stationary, depending on the test and number of lags used, while all tests and selected lags suggest that firm and individual overdraft credit lines interest rates are non-stationary.

Table 1. Panel unit root tests

	Im–Pesaran–Shin (1)	Im–Pesaran–Shin (3)	Levin–Lin–Chu (1)	Levin–Lin–Chu (3)
Receivables	0.001	0.047	0.015	0.168
Firm overdraft	0.142	0.579	0.160	0.659
Vehicles acquisition	0.001	0.001	0.003	0.042
Individual overdraft	0.999	0.999	0.999	0.999

Notes. Im–Pesaran–Shin test $\mathbf{H_0}$: all series has a unit root, $\mathbf{H_1}$: some series are stationary.
Levin–Lin–Chu test $\mathbf{H_0}$: all series has a unit root, $\mathbf{H_1}$: all series are stationary. $(\cdot)$ is the lag selection.

We ran Johansen cointegration tests for each pair of banks for each credit product, and report the proportion of cases where the tests suggest one cointegration vector. Lag selection was based on the SIC criteria (up to 3 lags). Table 2 shows that as a group the banks interest rates do not cointegrate. We also ran a robustness analysis controlling for the aggregate (government benchmark interest rate *Selic*) and the proportions change very little. Using Pesaran (2007) interpretation, the proportions at table 1 suggest that the null of no-cointegration would not be rejected in a panel setting, as the proportion of no-cointegrations are larger than 5 or 10%.

Table 2. Proportion of bank pairs that Johansen cointegration tests indicate at least one cointegration vector

	Proportion of cointegrations
Receivables	0.5217
Firm overdraft	0.4510
Vehicles acquisition	0.2449
Individual overdraft	0.3396

Notes. Authors' calculations are based on raw BCB interest rate data. Johansen cointegration tests using a 5% significance level and optimal lag length is based on minimum SIC criteria.

The above table is evidence that not all banks with the same product compete directly or indirectly, based on a panel cointegration. We thus proceed to the second set of models, namely pairwise ECM models. Note here that the regression models differ depending on which bank interest rate, from pair (*i, j*) is explained and which is the explanatory variable. Our tables below are such that the row bank is the explained variable and the column bank is the explanatory variable. Only coefficients significant at the 5% level are presented. The critical values for the significance test are from Ericsson and MacKinnon ECM cointegration test (Ericsson, MacKinnon, 2002; Enders, 2014), that is robust to non-stationarity of the bank pairs interest rate series. The banks are ordered by decreasing market share in the credit market, i.e. the largest banks are presented first and the smallest banks are presented to the right and down in the tables.

The matrix representation highlights patterns of competition. For receivables advances credit (table 3), reading down across columns, Caixa, HSBC and Bradesco are associated with interest rate changes that influence more banks. Looking across rows, Triangulo, Citibank and Fibra are the banks that respond to interest rate changes from more banks. Interestingly, these are medium to small banks. What may be important, no single bank interest rate change is associated with interest rate responses of three of the five largest banks (Banco do Brasil, Caixa and Itaú Unibanco). The other two largest banks react to only one (large) bank interest change. Regarding bank relative sizes we see that the upper right corner is empty of significant responses. This means that the largest banks (the upper rows) do not react to interest rate changes by smaller banks (rightmost banks). On the other hand the majority of smaller banks (the lower rows) tend to react to price changes from all bank sizes.

Coefficient wise, the results can be interpreted under a simple transformation, namely a half-life measurement. For example, mean length of response by Fibra bank to an interest rate change by Banco do Brasil is 0.65 $\left(=\ln(0.5)/\ln(0.345)\right)$ of a month. Mean time length of a response by Citibank to a change in interest rate by Banco do Brasil is 3.8 $\left(=\ln(0.52)/\ln(0.833)\right)$ months.

Table 4 presents results for the other firm credit line, overdraft credit. Again, less than half of the possible price correlations pairs have significant adjustment coefficients, suggesting that competition across banks is not wide spread. A similar pattern from table 3 appears regarding bank relative sizes influence. The largest banks (the upper rows) do not react to interest rate changes by smaller banks (rightmost banks). On the other hand the majority of smaller banks (the lower rows) tend to react to price changes from all bank sizes. And largest banks (the upper rows) when they react to other banks are mostly from largest banks (left columns).

The following two tables deal with credit lines for individuals. Vehicle acquisition (table 5) has the largest set of banks involved. It is a market where niche banks can operate, using a network of first and second hand car dealers. There is a group of medium sized banks that are vehicle manufacturer banks, such as Yamaha, for motorcycles, or GMAC for GM vehicles. This market was studied by Moita, Silva (2014) using a VECM specification on a few selected banks under an ex-ante hierarchy (large banks, followers, niche banks). The main pattern of niche banks competing more against each other's and largest commercial banks being more isolated is also found here. Banks of middle size (middle rows) are influenced by other banks, while smaller, non-manufacturer banks are isolated competition-wise from other banks (lower rows).

The market with the least number of significant speed of adjustment coefficients is individual overdraft credit. Table 6 indicates that only the smallest, non-public banks react to other banks interest rate changes (Alfa, BIC). An immediate interpretation is that there is very little competition in this credit line. This is coherent with this credit line characteristics. It is extremely expensive, with rates varying from 200 to 500% a.a., compared to a basic interest rate of 12% a.a. in government bonds. Even with this extremely high rates its use is due more to of lack of control over ones' finance than a planned contracting. The same banks offer, in all checking account statements, regular individual loans with interest rates (*crédito pessoal*) a third or a fifth of the cost of overdraft. The 'impulse' use and the required link with a (fee based) checking account, imposing very high costs for customers to switch banks suggest little competition among banks, as consumers seem less able to switch between them. Vehicle acquisition, on the other hand, does not require one to have a checking account in the credit supplier and the banks compete at the dealer level to offer credit.

Comparing across tables 3–6, i.e., comparing across markets, some overall patterns can be seen. First, there are fewer significant coefficients from smaller banks to larger banks. Second, there is non-symmetry in responses, with most responses by smaller banks. Third, less than half of adjustment coefficients are significant, suggesting weak rivalry across banks.

To save space we do not present robustness analysis where the benchmark interest rate on government bonds (*Selic*) rate is used. The main patterns across matrices maintains, with most coefficients of adjustment significance confirmed. The results below appear robust to the inclusion of this common cost.

To close this section, we take the estimated speed of adjustment coefficients from tables 3–6 and use them as dependent variable in a model to associate their magnitude with bank pair characteristics, as in (Iregui, Ortega, 2013). Central is the hypothesis that banks of similar size compete more intensely. To ease interpretation, we transform the speed of adjustment coefficients into median lags. So the hypothesis would mean that banks of different size would have a positive coefficient (longer time to adjust, if any finite period). The regression is weighted by the coefficients standard errors, recognizing that the dependent variable is composed of estimates.

Table 3. Speed of adjustment coefficients matrix: Firm credit receivables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
(1) BANCO DO BRASIL*													
(2) BRADESCO					–0.550								
(3) ITAÚ UNIBANCO													
(4) SANTANDER		–0.253											
(5) CAIXA*													
(6) HSBC		–0.460			–0.529		–0.459						
(7) BANRISUL*						–0.276							
(8) CITIBANK	–0.167	–0.337		–0.315	–0.328	–0.507	–0.342				–0.353		
(9) SAFRA													
(10) BRB*		–0.354	–0.475	–0.420		–0.402	–0.476						
(11) MONEO		–0.226			–0.259								–0.194
(12) TRIÂNGULO	–0.478	–0.451	–0.616	–0.498	–0.510	–0.437	–0.481	–0.493	–0.612				–0.455
(13) FIBRA	–0.655	–0.362	–0.360	–0.350	–0.440	–0.385	–0.355	–0.371	–0.364	–0.349		–0.372	

Notes. Cell is the adjustment coefficient in ECM model (1) where the row variable is bank *i* price response to a change in the column bank *j* price. ECM model with one lag in difference variables to accommodate serial correlation. Only coefficients significant at a 5% significant test are presented, based on cointegration test by Ericsson, MacKinnon (2002). Banks ordered by their market shares in the credit market, with largest banks first. (*) indicates state banks.

Table 4. Speed of adjustment coefficients matrix: Firm overdraft

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
(1) BANCO DO BRASIL*			–0.233		–0.414	–0.515					
(2) BRADESCO											
(3) ITAÚ UNIBANCO	–0.234										
(4) SANTANDER											
(5) HSBC	–0.512					–0.683					
(6) BANRISUL*	–0.465				–0.490						
(7) CITIBANK	–0.608	–0.837	–0.640	–0.481	–0.626	–0.625					
(8) SAFRA											
(9) TRIÂNGULO											
(10) IND. DO BRASIL	–0.271	–0.270	–0.259	–0.269	–0.269	–0.273	–0.281	–0.274	–0.274		
(11) LUSO BRAS.	–0.848	–0.642	–0.737	–0.784	–0.777	–0.802	–0.655	–0.598	–0.609	–0.682	

Notes. See Notes to the table 3.

Table 5. Speed of Adjustment coefficients matrix: Vehicle acquisition

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
(1) BANCO DO BRASIL*				–0.374										
(2) BRADESCO			–0.446		–0.281			–0.216						
(3) ITAÚ UNIBANCO	–0.300	–0.620		–0.261	–0.259	–0.353								
(4) SANTANDER	–0.445													
(5) BANRISUL*	–0.297		–0.310	–0.286										
(6) J. SAFRA								–0.125						
(7) HONDA	–0.325	–0.330	–0.347	–0.317	–0.335	–0.316		–0.313	–0.322	–0.340	–0.293	–0.328	–0.344	–0.368
(8) YAMAHA		–0.472	–0.384		–0.393	–0.433			–0.406	–0.325		–0.339		–0.398
(9) BRADESCO FIN.			–0.346		–0.294			–0.277					–0.233	–0.379
(10) ITAUCARD												–0.218	–0.215	
(11) GMAC														
(12) RODOBENS								–0.227						
(13) A. J. RENNER										–0.328				
(14) PECUNIA			–0.306		–0.343			–0.285	–0.389	–0.345				

Notes. See Notes to the table 3.

Table 6. Speed of Adjustment coefficients matrix: Individual overdraft

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
(1) BANCO DO BRASIL*												
(2) BRADESCO												
(3) ITAÚ UNIBANCO												
(4) SANTANDER												
(5) CAIXA*	–0.314											
(6) HSBC				–0.461								
(7) BANRISUL*												
(8) CITIBANK				–0.271		–0.299						
(9) SAFRA												
(10) BRB*												
(11) ALFA	–0.621	–0.672	–0.643	–0.661	–0.661	–0.691	–0.620	–0.633	–0.631	–0.616		–0.613
(12) BIC	–0.772	–1.077	–1.141	–1.160	–0.734	–1.202	–0.732	–0.884	–0.894	–0.769		

Notes. See Notes to the table 3.

The explanatory variables are the following: the absolute size difference for each bank pair (measured as average deposits from 2009–2015); a dummy whether the bank is state owned and a dummy whether the bank is foreign. Only significant (at 5%) coefficients are used.

Table 7. Determinants of half-life of interest rate adjustment

	Receivables credit	Firm overdraft credit	Vehicle acquisition	Overdraft credit
		Coefficient		
Intercept	–0.372	–0.467	–0.343	–0.900
	(0.023)	(0.037)	(0.069)	(0.042)
$\ln A_i - \ln A_j$	–0.025	–0.056**	0.008	0.066***
	(0.016)	(0.024)	(0.038)	(0.021)
State Owned	–0.032	0.059	–0.007	0.102**
	(0.021)	(0.042)	(0.099)	(0.051)
Foreign	–0.025	–0.053	0.010	0.542***
	(0.068)	(0.092)	(0.160)	(0.068)
Observations	40	32	49	25
F-Statistics	2.42	2.59	0.99	21.32
R-Squared	0.17	0.22	0.06	0.75

Notes. Dependent variable is the half-life $\ln(1/2)/\ln(1+a_i^j)$ calculated from the speed of adjustment coefficient from tables 3–6, excluding non-significant ones.
Observations are weighted by the standard error of the estimated coefficients.
The variable $|\ln A_i - \ln A_j|$ measures the log size difference in 2013 average assets between banks *i* and *j*. State owned is 1 if any of the banks *i* or *j* is state owned.
Foreign equals 1 if any of the banks *i* or *j* is classified as foreign by the Central Bank of Brazil.
*, **, *** — significant at the 10%, 5% and 1% level respectively.

The results in table 7 show that bank size influences the speed of adjustment, our strength of competition proxy, only on overdraft accounts (firm and individual). For firm overdraft when bank pairs are of similar size, the median time to adjustment increases. This suggests that this market is segmented as the more different sized the banks are the longer it takes for a bank to react to another bank price or loan interest rate increases. For individual overdraft credit, the result is the opposite: Bank pairs of similar size have quicker (smaller) mean adjustment time. For individual overdraft credit, when at least one of the banks is state owned, adjustment is longer (would be more isolated from competition). The same effect is measured when at least one of the banks is a foreign bank (such as Santander, or HSBC or Citibank).

In general, the bank characteristics are not related to the speed of adjustment coefficients. Apart from individual overdraft R-squares are small. In the Colombian market, Iregui, Otero (2013) found faster adjustment when banks were of different sizes for deposit rates and the opposite for lending rates (we use lending rates only).

Given the caveats in interpreting significant interest rate bank pairs coefficients, the results suggest that there is few direct (or indirect) competition on interest rate between banks in these markets in Brazil. Particularly for overdraft credit, large banks seem insulated, exploring situational monopolies from their checking account clients.

Concluding comments

While there are well known econometric methods to evaluate the overall competition level in the banking market such as the Boone index, there are few methods to infer about bank-to-bank competition, or closeness of competition. Estimating a full demand model of bank credit with volume and interest rate data, one could calculate diversion ratios used to infer closeness of competition across units. The very high data requirements of the demand estimation suggest the attractiveness of less demanding data methods. Such methods explore only prices in the markets and their correlations.

While price correlations are widely used to delimit relevant markets and have been explored by Iregui, Otero (2013) to evaluate the law-of-one-price argument of competition, we use dynamic price correlations, in a cointegration autoregressive distributed lag model (Ericsson, MacKinnon, 2002) to infer about non-price competition of different bank pairs.

Price correlations for relevant market definition have been subject to a number of studies that highlight its limitations. The very high likelihood of false significant correlation between prices of unrelated goods due to the omission of common cost or demand shocks are well documented. Nevertheless the low information requirements make price correlations attractive, as long as interpreted with appropriate care. We recognize the caveats and argue that price correlations of firms within a market may provide evidence on which firms *do not* compete with each other.

Applying the price correlations for four different credit products using monthly interest rates collected by the Central Bank of Brazil, we estimate cointegration models. The ECM models are interesting as they provide a metric that is robust to series non-stationarity and remain valid under cointegration (with an adjustment of the coefficient hypothesis test distribution).

Our results show that, even under a clear bias to falsely significant price correlations, many banks do not react to other bank interest rate changes. The least competitive market is the individual overdraft credit (compared to receivables, firm overdraft and vehicle acquisition), where only smaller banks, if any, may react to interest rate changes from other banks. This is expected as the overdraft loans are used mostly out of convenience and impulse as they are always available and very expensive, compared to other loan types such as consumer credit or vehicle purchase credit.

The estimated speed of adjustment coefficients from the pairwise ECMs were used in a regression model to differentiate coefficients by bank characteristics, namely size differences, state owned and foreign. If omitted variables, such as common cost or demand shocks or other prices, are influencing the results, we expect no explanatory power from our models. Our results show that the selected variables cannot explain the speed of adjustment patterns, apart from individual overdraft credit and, to a lesser extent, firm overdraft credit. In the latter case, price reactions are faster the more similarly sized banks are and the former case price reactions are faster the more different sized bank pairs are. The individual overdraft case suggests a more segmented market.

The relatively low competition between banks, as measured by large number of nonsignificant price correlations, call into attention the need to «grease competition wheels» in the banking sector in Brazil. A reduction in customer switching costs should be high in the regulator's agenda.

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