ECONOMIC ANNALS, Volume LXVI, No. 228 / January – March 2021
UDC: 3.33 ISSN: 0013-3264

https://doi.org/10.2298/EKA2128101G

*Mohamed Sadok Gassouma**
*Kais Ben-Ahmed***

THE ROLE OF FOREIGN BANKS IN THE TRANSMISSION OF MONETARY POLICY: EMPIRICAL EVIDENCE FROM TUNISIA

ABSTRACT: *This paper presents an empirical analysis of the effect of monetary policy shocks on credit supply in Tunisia, using a vector autoregressive model and a nonlinear interactive model. The focus is on the magnitude of these shocks in the presence of foreign banks. The variables of interest are the concentration index of deposit banks, and monetary policy shocks based on the monthly data of 27 universal and business banks covering the period 1993 to 2016. The results support a positive and significant impact of concentration index on credit supply. However, monetary policy shocks appear to have no significant effect when the market is concentrated with the entry of foreign banks. The findings of this study also reveal that the entry of foreign banks neutralises monetary policy shock transmission in the credit supply, which may be offset by market discipline.*

KEY WORDS: *policy shocks, competitiveness, credit supply, banks, Tunisia.*

JEL CLASSIFICATION: E-E5-E52

* Higher Institute of Theology, Ez-zitouna University,Tunisia, ESCTRA Laboratory IHEC Carthage, Tunisia e-mail: gasadok@yahoo.fr

** University of Jeddah, College of Business, Department of Finance, Saudi Arabia
Higher Institute of Management, Sousse University, Sousse, Tunisia
e-mail: benahmedkaies@gmail.com

1. INTRODUCTION

Monetary policy shocks have gained the attention of policymakers and academics because they can be one of the primary causes of bank failure and trigger a financial crisis. However, there is still a lot of uncertainty around the effects of monetary policy shocks, despite twenty years of empirical research and many methodological advances. Studying the effects of monetary policy shocks is a difficult endeavour. Most researchers concentrate on examining the effects of monetary policy shocks on credit supply, while few works consider financial liberalisation as a relevant factor affecting the causal link between monetary policy shocks and credit supply. In light of this gap, several studies have examined the effects of financial liberalisation on credit risk. The current study bridges the existing gap in the literature by investigating the importance of liberalisation in the financial markets, measured by banking competitiveness.

This study refers to the theories of Bensaid and Palma (1995), Panzar and Rosse (1987), Gunji et al. (2009), and Gopalan and Rajan (2017). The objective is to demonstrate how a Nonlinear Interactive (NLI) model can connect these theories to evaluate the joint effect of policy monetary shocks and bank competitiveness on credit supply by regressing both monetary policy shocks and the combination of these shocks and the competitiveness index on the credit supply. The primary objective is to empirically investigate whether the concentration of foreign banks affects the causal relationship between the monetary shocks resulting from unexpected fluctuations in interest rates and credit supply in the Tunisian economy.

Regarding monetary policy shocks, our modelling is based on the works of Gunji et al. (2009), who use the residue of the interest rate equation in the VAR model. This model allows calculating the effect of monetary policy's exogenous shocks on economic factors. In other words, the authors consider the impulse response function, as the effect on the economic variables of exogenous shock changes in the interest rate, as a monetary policy.

The presence and design of foreign banks in the Tunisian market are formalised according to Panzar and Rosse's (1987) competitiveness index, showing the concentration of foreign bank ownership in the Tunisian interbank market. The underlying index measures the degree of competitiveness linked to the entry of

foreign banks. The underlying index differs from the classical Herfindahl-Hirschman Index (HHI) and the Concentration Ratios range (CRn) in two important respects. The latter indexes measure domestic banks' degree of concentration, whereas the Panzar-Rosse H-statistic defines a measure of competitive intensity that encompasses more than just domestic banks and includes foreign banks.

Primarily, there are two trends of opinion on this issue, with opposing approaches. The primary trend supports the adverse role of competitiveness in the connection between financial shocks and credit flexibility. Among other things, increased competitiveness and foreign concentration cause monetary shocks to increase the supply of credit and therefore attract more borrowers of lower quality, corrupting asset quality and increasing credit risk. Accordingly, financial deregulation together with the progression of the interest rate build the loan fee, resulting in an intermediation edge inconsistent with a market structure characterised by high competitiveness and an increased concentration of foreign banks. Therefore, this makes the decision criteria flow, increases the supply of credit, and degrades its quality (Grop and Vesela 2004; Bikker and Haaf 2002; Chan et al., 1986; Manove et al., 2001; Gehrig 1998; Marquez 2002; Bolt et Tieman, 2004; Reppulo 2004; Hellman et al., 2000; Ellizalde and Reppulo 2004).

On the other hand, in the presence of greater liberalisation following a rapid growth of credit supply, monetary policy shocks can negatively affect macroeconomic magnitudes. However, Ida et al. (2018) use a varying coefficient Bayesian panel VAR model, where the coefficients are allowed to vary as a function of the degree of financial, product, and labour market regulation, on data from 1976Q1–2006Q4 for 19 OECD countries. The object is to test whether the current account improves or deteriorates following a monetary policy expansion. Their empirical results support the theory. They therefore conclude that following a monetary policy expansion, the current account is more likely to go into deficit in countries with more liberalised financial markets. To this effect, Maudos and Fernandez (2004) argue that acquisitions can increase banking concentration but that they reduce the quality of loans. In addition, Alencar and Nakane (2004) observe that competitiveness makes the economy more sensitive to interest rates.

The second trend supports the proposition that competitiveness mitigates the impact of monetary policy shocks by either diminishing or maintaining the credit supply. This can improve loan quality and does not attract risky, bad borrowers. Within this paradigm, Chen and Haller (2003) and Dermigug-Kunt and Detragiache (2002) demonstrate that liberalisation better controls monetary policy shocks by further diversifying banking products. This makes up for intermediation losses with the profits from specialisation and accordingly makes it easier to control credit risk. In addition, Stiglitz and Greenwald (2003) use the mean-variance approach to analyse the banking system mechanism under the risk-averse constraint. They find that the increase in the interest rate applied to credits in a regulated market is particularly important in a competitive market and that this increase leads to lower deposits that can reduce credit supply.

On the other hand, competiveness can reduce the effect of monetary policy shocks on credit supply, as demonstrated by Adams and Amel (2005), who use U.S. data to test the impact of bank concentration on the transmission of monetary policy. They find that the effect of monetary policy on bank loans is weaker in concentrated banking markets. Their analysis differs somewhat from ours in some key areas. For example, they use the Herfindahl Index as a measure of bank concentration, while we use an index of the degree of bank competition.

VanHoose (1985) shows that competiveness cannot have any effect on the relationship between monetary shocks and credit supply by arguing that if the central bank uses the monetary market security rate as a policy instrument, changes in bank competition will have no impact on monetary control.

On the other hand, the positive effect of monetray policy shocks on credit supply in the presence of competiveness can be neutralised by substituting the interest rate with the cutting score as a thresholder credit decision. Chen (2005) studies the change in borrowers' behaviour when moving from a monopolistic market (in which a single bank operates) to a more competitive one (in which a foreign bank joins its domestic counterpart to compete). The passage between the two markets is governed by the arbitration between the interest rate and the score threshold offered by banks. The authors show that as part of a monopolistic market before the entry of the foreign bank, the interest rate is lower than the one that may be charged as part of a duopoly market. On the other hand, income

volatility is higher after financial liberalisation. The function of the revenue is concave, so that the revenue increases to an optimal threshold and then decreases.

The same author continues his study in 2007 to check the evolution of credit supply following such a gap in competiveness. He conducts a survey of banks in the European Union after globalisation. The research finds that the interest margin decreased while the index of competitiveness increased, and consequently the quality of loans improved. This leads to a decrease in credit risk. After the liberalisation of the banking system the lender's interest rate decreases even though the monetary market rate increases and the interest volatility falls. This result reflects the cautious behaviour of the bank because it is not based on the interest rate; rather it is based on information technology and the increase of the score threshold.

The current study is closely related to the above studies; however, when analysing bank supply and monetary policy shocks we found no empirical analysis that considers the financial liberalisation effect. Unlike previous studies that focus exclusively on examining the relationship between monetary policy shocks and credit supply, we show how introducing the competitiveness of foreign banks into the analysis can affect this relationship.

The rest of this paper is structured as follows. Section 2 presents the model specifications. The methodology is described in section 3. The statistical analysis, namely results and diagnostic checks, are presented in section 4. Section 5 provides a conclusion.

2. MODEL SPECIFICATIONS

The model specifications are derived by adopting a three-step procedure. First, we estimate a VAR model to generate impulse responses to monetary policy shocks. Second, we test if the liberalisation of the financial markets affects the impact of monetary policy shocks on credit supply. Third, we estimate our model by deriving both optimal loan supply and competitiveness.

2.1. The VAR models

In all the following, let MMR, CRS, LIQ, and GDP be the monetary market rate, credit supply, liquidity, and gross domestic product, respectively. Following the methodology of Gunji and Miura (2017), we employ a VAR model of the following form.

$$\begin{bmatrix} MMR_t \\ CRS_t \\ LIQ_t \\ GDP_t \end{bmatrix} = \begin{bmatrix} \sum_{i=1}^{2} \alpha_i MMR_{t-i} + \sum_{i=1}^{2} \beta_i CRS_{t-i} + \sum_{i=1}^{2} \theta_i LIQ_{t-i} + \sum_{i=1}^{2} \lambda_i GDP_{t-i} \\ \sum_{i=1}^{2} \alpha_i CRS_{t-i} + \sum_{i=1}^{2} \beta_i LIQ_{t-i} + \sum_{i=1}^{2} \theta_i GDP_{t-i} + \sum_{i=1}^{2} \lambda_i MMR_{t-i} \\ \sum_{i=1}^{2} \alpha_i LIQ_{t-i} + \sum_{i=1}^{2} \beta_i GDP_{t-i} + \sum_{i=1}^{2} \theta_i MMR_{t-i} + \sum_{i=1}^{2} \lambda_i CRS_{t-i} \\ \sum_{i=1}^{2} \alpha_i GDP_{t-i} + \sum_{i=1}^{2} \beta_i MMR_{t-i} + \sum_{i=1}^{2} \theta_i CRS_{t-i} + \sum_{i=1}^{2} \lambda_i LIQ_{t-i} \end{bmatrix}$$

Christiano et al. (2014) show that placing the impulse response function of the MMR in front of other economic variables does not depend on the order of the variables to be imputed in the VAR model. The monetary policy shocks (MPS) are the result of an unexpected change in the monetary market rate (MMR) because of a sharp credit supply (CRS) shock (cf. Equation 1 in the VAR model). These shocks consist of fluctuations in the MMR not considered by the Tunisian Central Bank (TCB), but which can cause an excess of credit risk in Tunisian banks. The details of the estimated error response function of the VAR model are given in the methodology.

2.2. A model for banking competitiveness

Contrary to the classic concentration indexes of Herfindahl–Hirschman (HHI), which are mostly used in the literature, the concentration index used in this paper aims to measure the concentration of foreign banks through a competitiveness index. This index provides a solid measure of the degree of competitiveness linked to the entry of foreign banks. The lower the index, the more the market is concentrated and uncompetitive, and the higher the index, the more the market is liberalised and competitive.

The competitiveness index model uses the work of Panzar and Rosse (1987) and Yeyati and Micco (1987; 2007). The model used to generate a concentration index for each year takes the form

$$\text{Log}(\text{ROA}_i) = \beta_{0i} + \beta_{1i}\text{CE} + \beta_{2i}\text{SE} + \beta_{3i}\text{AC} + \beta_{4i}\text{SD} \quad (1)$$

The parameters of the model represented by Equation (1) are $\beta_i = (\beta_{0i}, \beta_{1i}, \beta_{2i}, \beta_{3i}, \beta_{4i})$, where β_{0i} is the intercept, β_{1i} is the regression coefficient corresponding to the cost of equity (CE), β_{2i} measures the effect of the share of equity (SE), β_{3i} corresponds to the asset costs (AC), and β_{4i} measures the effect share of deposits (SD). The variable of interest is economic performance, the return on assets (ROA). More importantly, β_k, k = (1, 2, 3, 4) measures the elasticity between each of the variables CE, SE, AC, and SD on the one hand and ROA on the other.

A parameter estimation of the model in Equation 1 is performed using the biannual data of 27 Tunisian banks over the period 1993–2016. The panel data set for each year is composed of 54 (27×2) observations. Consequently, we get 24 parameters β_{kt} (k = (1, 2, 3, 4); t = 1993, …, 2016)) for each variable over the period 1993–2016. The competitiveness index, the crux of our analysis, denoted as H, is the sum of the elasticities β_k for each year. Thus, the H-index can be computed as

$$H_{1993 \le t \le 2016} = \sum_{k=1}^{4} \beta_{kt}$$

Estimated parameters of the model are presented in Table 4, while competitiveness index H is displayed in Figure 2.

2.3. Deriving optimal banking competitiveness

Considering bank competitiveness in this study is important when the object is to test the effect of political monetary shocks on credit supply. Furthermore, we want to see how the effect of political monetary shocks imposed on the credit supply by the monetary rate market differs between monopolistic and liberalised markets. We also want to demonstrate the effect of liberalisation on the

relationship between credit supply and monetary policy shocks in Tunisian banks. Therefore, we consider an interactive model taking the form:

$$\text{Log(CRS)} = \alpha_0 + \alpha_1 \text{MPS} + \alpha_2 \text{MPS} \times \text{CPS} \quad (2)$$

As in the case of the VAR model, the credit supply (CRS) is expressed in logarithm, and the CSP variable represents banking competitiveness. The logarithm is used for standardisation reasons: it does not affect the logic but approximates the other variables just for the credit supply value expressed in Tunisian dinars, and consequently facilitates the interpretation of the results. As has been noted before, monetary policy shocks (MPS) are the response of the monetary market rate (MMR) when a credit supply shock (CRS) occurs.

Banking competitiveness allows splitting the effect of shocks on the interest rate in two ways: if the competitiveness index is low, shocks decrease the supply of credit. However, if competitiveness is high, monetary policy shocks are no longer regulated by the supply of credit. This is shown through the following relationship:

$$\frac{\partial \text{Log(CRS)}}{\partial \text{MPS}} = \alpha_1 + \alpha_2 \text{CPS} \quad (3)$$

The range at which the causal relation between MPS and CRS changes is $\frac{\partial \text{Log(CRS)}}{\partial \text{MPS}} = 0$. Namely, for $\text{CSP} = -\frac{\alpha_1}{\alpha_2}$, MPS causes an increase or decrease in the CRS. It should be noted that the shock introduced in model 2 is expansionary (positive shock), reflecting an unexpected increase in MMR faced with a shock on the CRS.

3. MATERIALS AND METHODS

3.1. Data

This study uses the balance sheets of the 27 universal (25) and business (2) banks in Tunisia covering the period January 1993 to December 2016. The variables in the balance sheets relevant to this study are the CRS, MMR, LIQ, and GDP. Data for these variables were collected directly from the International Monetary Fund

(IMF), available on the Tunisian Central Bank (TCB) website (International Financial Statistics). With regard to the bank's competitiveness, this paper examines if MPS transmission, embodied by the interest rate, affects CRS.

Table 1: Summary statistics of the variables included in the VAR model

Variable	Mean	Std. dev	Min	Max	Skewness	Kurtosis	J-Bera	N°of obs.
CRS	5.881	1.743	3.160	10.810	1.169	1.169	3.629	288
LIQ	9.329	0.720	8.167	10.381	–0.720	–0.720	9.329	288
MMR	10.044	0.685	8.987	11.202	0.050	0.050	8.790	288
GDP	3.906	2.006	–1.900	7.150	–0.840	–0.840	4.339	288

Source: Authors' computation

It is important to indicate that credit supply (CRS) and liquidity (LIQ) are expressed in logarithm, while monetary market rate (MMR) and gross domestic product (GDP) are expressed in percentages. Descriptive statistics of these macroeconomic variables, the data for which are available monthly, are presented in Table 1. Consequently, we obtain 288 observations (24 x12), which cover the period 1993 to 2016.

3.2. Methodology

Based on the VAR model, we generate the impulse responses of the variables in MMR_t to monetary policy shocks, which are identified by imposing a triangular orthogonalization. However, first the stationarity[1] and the optimal lag order of the VAR model are checked. The issue of lag-length[2] selection is examined based on the Augmented Dickey-Fuller test and Akaike, Schwarz, and Hannan-Quinn Information Criterions.

[1] Stationarity model VAR is verified but does not appear in the paper.

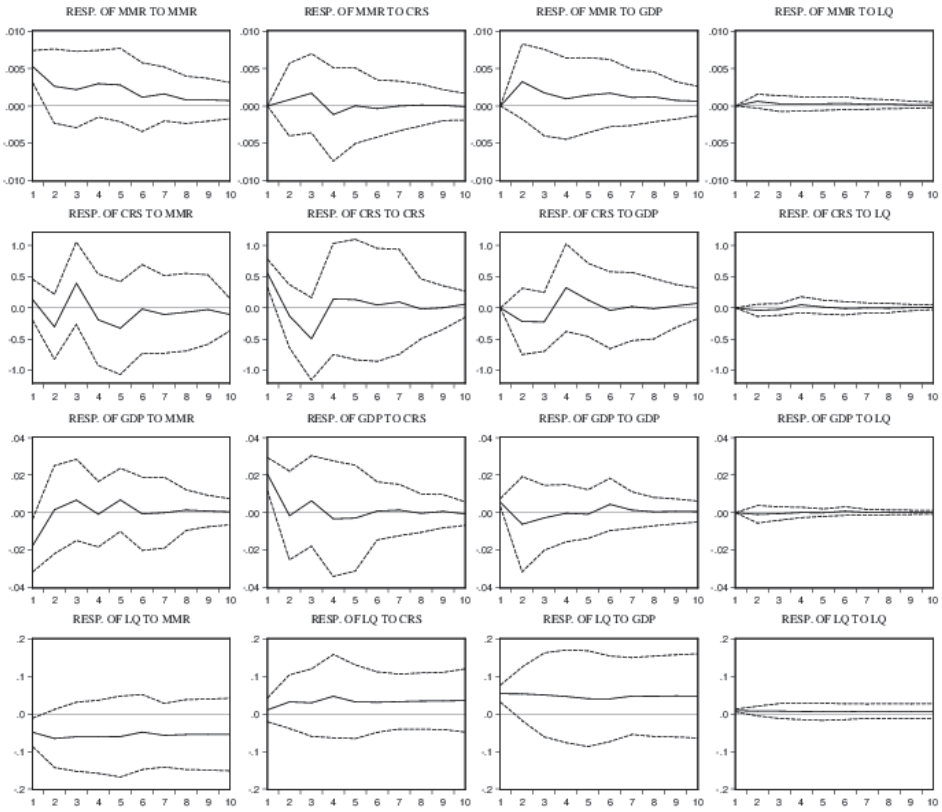
[2] The lag order selection criteria of the VAR model are verified but do not appear in the paper.

Table 2: Stationarity of VAR model

Variable	ADF statistics	Test critical values	Probability
MMR	–3.4271	–3.1449	0.0315
CRS	–4.3672	–3.1753	0.0078
LIQ	–3.4462	–3.2126	0.0353
GDP	–3.1478	–3.1199	0.0477

Table 2 shows that all used variables introduced in the VAR model are stationary at a 5% significance level (p-value < 0.05). Only the first lag is significant and turns out to dominate the others in size.

Figure 1: Empirical Impulse Responses



Note: Orthogonalized error responses to monetary policy shocks. The solid line displays empirical error responses. The dashed lines are 90% error bounds.

Figure 1 clearly shows that the impulse response of credit supply (decrease/increase) resulting from a monetary policy shock is significant. The solid line that displays error responses is well within the 90% confidence interval. Therefore, bank loans decline persistently for the first year and then increase gradually for about two years. The drop-in bank loans continue after around three years, reach a peak after four years, and subsequently return to the baseline. This result shows that the credit supply is highly sensitive to unexpected fluctuations in the MMR. Furthermore, a shock in the MMR results in an immediate and automatic adjustment of the credit supply. Moreover, the monetary policy shocks of the TCB are resolved by adjusting the credit supply to the economy. The effect of the monetary policy shocks reflects the sensitivity of certain economic agents – including borrowers – which are faced with systematic risk. Therefore, an unexpected change in interest rates affects the solvency and the behaviour of borrowers.

Table 3: Summary statistics

Variable	Mean	Std. dev.	Min	Max	Skewness	Kurtosis	J–Bera	N°of obs.
CA	0.0063*	0.0023	0.0038	0.0104	0.4776	1.9270	0.8598	1296
CE	0.3525*	0.0879	0.1537	0.4450	–1.3250	3.6949	3.1272	1296
SE	0.0139*	0.0024	0.0100	0.0176	0.0032	1.8563	0.5450	1296
SD	0.5393*	0.0911	0.3246	0.6493	–1.2912	4.1920	3.3706	1296
ROA	0.0081*	0.0032	0.0042	0.0150	0.8039	3.2219	1.0975	1296

*Mean is statistically not different from zero at the 5% significance level.

Source: Authors' computation

Now, we estimate Equation (1) for each year, to obtain an annual concentration index H. Consequently, we get 24 models for each year, and each model is for 27 banks and two semesters. We obtain a panel data set of 27 banks and 2 semesters over the 1993–2016 period. Descriptive statistics of the variables considered by the Equation (1) model are presented in Table 3.

Table 4 presents estimates of concentration indexes, represented by the coefficients in Equation (1). As defined by Panzar and Rosse (1987), if the concentration index is negative ($H \leq 0$) the market is monopolistic; if the concentration index is between 0 and 1 ($0 \leq H \leq 1$) the market is semi-

competitive; and if the concentration index is equal to unity (H = 1) the market is perfectly competitive.

Table 4: Estimation of the concentration index†

Year	β_0	β_1	β_2	β_3	β_4	R-squared	Fisher	DW
1993	–0.004	–0.003	–0.854	–0.233	–0.042	0.513	0.006	2.010
	(–2.084)	(–2.018)	(–2.131)	(–2.048)	(–2.290)			
1994	–0.001	–0.019	–0.175	0.019	0.024	0.481	0.004	2.610
	(–2.084)	(–2.018)	(–2.031)	(–2.278)	(–2.409)			
1995	–0.001	–0.087	–2.879	–0.491	–0.005	0.767	0.036	2.172
	(–2.040)	(–6.395)	(–3.613)	(–2.333)	(–2.167)			
1996	0.020	–0.078	–1.929	–0.049	0.027	0.981	0.046	2.835
	(–4.849)	(–8.364)	(–9.105)	(–17.963)	(–3.742)			
1997	0.005	0.033	–1.429	–0.113	–0.036	0.949	NA	NA
	(–2.200)	(–5.263)	(–4.771)	(–2.494)	(–5.262)			
1998	0.018	–0.021	–0.959	–1.156	0.008	0.711	0.003	2.040
	(–2.688)	(–2.528)	(–2.762)	(–2.715)	(–2.999)			
1999	0.032	–0.026	–1.141	0.274	0.007	0.708	0.013	2.314
	(–6.201)	(–2.712)	(–3.288)	(–2.880)	(–2.801)			
2000	0.031	–0.032	–1.210	0.826	0.007	0.798	0.001	2.705
	(–6.223)	(–4.992)	(–4.735)	(–3.501)	(–2.387)			
2001	0.0266	0.026	–1.114	0.947	0.007	0.981	0.007	2.250
	(–45.135)	(–23.122)	(–18.312)	(–5.393)	(–7.957)			
2002	0.026	–0.040	–1.329	0.473	0.019	0.916	0.007	2.153
	(–5.040)	(–5.998)	(–5.635)	(–2.267)	(–2.679)			
2003	0.025	–0.033	–1.567	0.512	0.020	0.931	0.019	2.367
	(–4.381)	(–6.342)	(–7.379)	(–2.426)	(–3.130)			
2004	0.032	–0.050	–1.597	0.923	0.014	0.881	9.234	2.469
	(–4.161)	(–4.557)	(–6.746)	(–2.395)	(–2.237)			
2005	0.042	0.049	–1.089	–0.080	–0.002	0.921	0.006	2.773
	(–4.480)	(–9.9132)	(–4.277)	(–2.832)	(–2.255)			
2006	0.030	–0.190	–0.058	–5.945	0.094	0.975	0.016	2.206
	(–2.770)	(–10.026)	(–2.350)	(–4.524)	(–2.501)			
2007	0.024	–0.030	–0.063	–0.127	0.006	0.915	0.001	2.246
	(–6.092)	(–5.483)	(–2.207)	(–2.019)	(–2.211)			
2008	0.012	–0.014	–0.285	0.328	0.007	0.532	0.001	2.050
	(–3.992)	(–4.062)	(–2.511)	(–2.356)	(–2.920)			
2009	0.028	–0.029	–0.944	0.325	0.005	0.553	0.000	2.749
	(–7.331)	(–5.533)	(–5.006)	(–2.211)	(–2.894)			
2010	0.0327	–0.057	–0.074	–0.227	–0.007	0.742	0.000	2.333
	(–6.015)	(–6.350)	(–2.313)	(–2.658)	(–2.937)			

2011	0.017	–0.003	–0.506	0.509	0.006	0.982	0.006	2.196
	(–4.121)	(–3.663)	(–2.367)	(–2.742)	(–2.664)			
2012	0.019	–0.021	–0.613	0.543	0.006	0.892	0.007	2.142
	(–4.618)	(–3.611)	(–2.631)	(–1.780)	(–2.545)			
2013	0.013	–0.016	–0.272	0.509	0.006	0.681	0.050	2.127
	(–2.925)	(–2.591)	(–2.075)	(–2.597)	(–2.466)			
2014	0.008	–0.015	0.079	0.734	0.001	0.872	0.055	2.056
	(–2.622)	(–2.054)	(–2.247)	(–2.935)	(–2.169)			
2015	0.011	–0.0$16	–0.170	0.627	0.005	0.612	0.054	2.043
	(–2.320)	(–2.363)	(2.618)	(–2.838)	(–2.179)			
2016	0.027	–0.047	–0.066	–0.1574	–0.004	0.963	0.000	2.407
	(–6.340)	(–6.193)	(–2.238)	(–2.231)	(–2.650)			

†The Hausman test accepts the fixed effect for all models.

Note: t-statistics are in parentheses.

Source: Authors' computation.

Figure 2 describes the evolution of the degree of liberalisation of Tunisian banks. Concentration indexes between 1997 and 2013 are less than zero, indicating that despite the entry of foreign banks, the interbank market remains monopolistic and Tunisian banks are rather concentrated. Therefore, most capital shares are held by the same people, which represent private institutions, industrial owners, and public institutions, with the state as controller.

Figure 2: Annual change in concentration indexes (H-Statistic) of Tunisian banks

The deconcentration shows that when the Tunisian banking sector is liberalised, the market changes from monopolistic to imperfect competition by differentiating banking products or substituting the increasing interest rate with the cutting score, as showed by Chen (2005). The concentration indexes were positive from 2013 to 2015, to decrease again in 2016, showing the outflow of foreign capital in the recent period.

4. STATISTICAL ANALYSIS

4.1. Results

Once the monetary policy shocks and concentration index have been estimated using the VAR model and the Equation (1) model respectively, they are included in the Equation (2) model to estimate their effect on credit supply. Descriptive statistics of the variables credit supply, monetary policy shocks, and competitiveness (H-index) retained in model (2) are presented in Table 5. Table 6 reports the estimates results using Least Square (LS) methodology for 288 time observations.

Table 5: Summary statistics of the effect of monetary policy shocks and competitiveness on credit supply

Variable	Mean	Std. dev.	Min	Max	Skewness	Kurtosis	J–Bera	N°of obs.
CRS	5.881	1.743	3.160	10.810	1.169	3.672	3.629	288
MPS	0.000	0.180	1.824	0.473	–5.103	44.910	2.354	288
H-index	–0.918	1.451	–6.192	0.797	–2.242	8.606	1.553	288

Source: Authors' computation

From Table 6 we can see that monetary policy shocks (MPS) have a significant negative effect on the credit supply. This result shows that in the event of exogenous monetary policy shocks by the TCB, the Tunisian banks reduce their credit supply to the economy to avoid the suspension of payments by taxpayers.

Table 6: Estimation of monetary policy shocks and bank's competitiveness effects on credit supply

Variable	α_0 – Intercept	α_1 – MPS	α_2 –MPS*CPS	R^2 (%)	F-statistics	DW
	16.607	–0.111	21.767	49.97	8.290	2.174
	(32.600)	(–2.572)	(2.420)		0.000†	

†The Hausman test accepts fixed effects for all models.

Note: *t*-statistics are in parentheses

Source: Authors' computation

The decrease in credit supply is the result of the increased refinancing cost of Tunisian banks. Thus, this drop-in credit supply ensures solvency between the TCB and banks on the one hand, and between banks and borrowers on the other hand, thus mitigating the the credit risk. Unlike what happens in a crisis, when facing a probable monetary policy shock, Tunisia might mitigate the credit risk by reducing the supply of credit to the economy. However, competitiveness allows splitting the effect of interest rate shocks in two ways. If the competitiveness index is low, shocks decrease the supply of credit. However, if competitiveness is high the monetary policy shocks are no longer regulated by the supply of credit. This is shown by the following equation:

$$\frac{\partial logCRS}{\partial MPS} = -0.111 + 21.767CPS = 0$$

As a result, if competitiveness is less than the value of 0.005 (α_1/α_2), clearly there is a significant negative relationship between credit supply and monetary policy shocks. However, this relationship changes from negative to positive if competitiveness exceeds the value of 0.005. Therefore, when the market is monopolistic ($H \leq 0$), interest rate shocks are regulated by reducing the supply of credit.

With the entry of foreign banks, the Tunisian bank market changed from being monopolistic to competitive. It therefore remains to be seen whether Tunisian banks operate within a framework of pure and perfect competitiveness. To do this, we follow Palma's approach, which shows that in an imperfect competition market the interest rate rises and does not fall.

Finally, we check the correlation between the concentration indexes and the interest rate for the 2013 to 2016 period, when the concentration index became positive. The correlation measured between the two variables is positive. Thus, when foreign banks enter the Tunisian banking market the interest rate rises and does not fall. This result is consistent with Gapalan and Rajan's (2017) findings, which report that with lower banking competition the interest rate declines, but with high competition the interest rate rises and enhances pass-through.

Accordingly, we conclude that when banking competitiveness is enhanced, the interest rate rises and systematically the credit supply rises also. In this sense, banking liberalisation leads to an increase in interest rate shocks and does not decrease the supplied credit (Gunji et al. 2009).

4.2. Diagnostic checks

All our results hinge on a string of three estimations: the VAR model, Equation (1), and Equation (2). Accordingly, a sequence of diagnostic checks must be initiated at each step and for each model. Our diagnostic checks are based on the Akaike Information Criterion (AIC), Schwartz Information Criterion (SIC), Log-likelihood (LV), and Hausman test. These tests help us to select which of the different models are significant. When test conditions are met, a null hypothesis can either be accepted, or rejected in favour of an alternative hypothesis.

The parameters αs, βs , θs, and λs in the VAR model are estimated using Ordinary Least Square (OLS) on each equation. However, first we need to select the optimal lag lengths *p,* since inference is dependent on the correctness of the selected lag order (Hacker and Hatemi, 2008). We get the following results.

Table 7: Model selection criterion of the VAR model

Order	LV	AIC	BIC
P = 1	1043.54	−11.56*	−11.44
P = 2	1045.48*	−11.54	−11.46*
P = 3	1040.12	−11.53	−11.33
P = 4	1036.52	−11.42	−11.25
P = 5	1033.69	−11.36	−11.15
P = 6	1030.54	−11.32	−11.10
P = 7	1028.24	−11.25	−11.05
P = 8	1021.27	−11.12	−11.02

* indicates the lag order selected by the criterion

Table 7 presents the LV, AIC, and BIC values for VARs with one to eight lags. These should be interpreted as fit statistics that describe the improvement in the log-likelihood, penalised for the additional lags. Smaller values of AIC and BIC fit statistics are better because they are based on the negative of the log-likelihood. However, higher values of LV fit statistics are retained. Referring to Table 7, the lowest AIC and BIC criterion is given for the lags of order 1 and 2 respectively. However, the highest LV criterion is given for the lag of order 2. A case could be made for 2 lags in view of the time coincidence between BIC and LV criteria, versus 1 lag for only the AIC criterion. Consequently, we will choose 2 lags for the lagged variables.

Table 8: VAR Model Estimation

	MMR	CRS	LIQ	GDP
MMR (−1)	0.777635 (0.53553) [2.45210]	−84.46201 (58.1543) [−2.45238]	0.865632 (7.60355) [2.11385]	−0.964185 (2.80596) [−5.34362]
MMR (−2)	0.188366 (0.52718) [3.35731]	63.38030 (57.2475) [3.10713]	−1.483116 (7.48499) [−2.19815]	1.609473 (2.76221) [2.58268]
CRS (−1)	0.003289 (0.00311) [2.05665]	−0.231271 (0.33803) [−3.68417]	−0.014252 (0.04420) [−2.32248]	0.001417 (0.01631) [3.08686]

CRS (–2)	0.001739 (0.00281) [2.61925]	–0.166084 (0.30490) [–1.54471]	0.003579 (0.03987) [3.08977]	0.011968 (0.01471) [1.81348]
LIQ (–1)	0.068463 (0.04938) [2.38644]	–4.204576 (5.36240) [–1.78408]	0.838212 (0.70112) [2.19553]	–0.108774 (0.25874) [–2.42041]
LIQ (–2)	–0.073409 (0.05045) [–3.45510]	5.795224 (5.47843) [2.05783]	0.175530 (0.71629) [2.24505]	0.097699 (0.26434) [3.36960]
GDP (–1)	–0.084729 (0.14313) [–2.59199]	2.023397 (15.5424) [2.13019]	1.536162 (2.03214) [1.75593]	–0.059935 (0.74993) [–2.07992]
GDP (–2)	–0.051611 (0.10021) [–2.51504]	–13.90551 (10.8818) [–2.27787]	–0.059891 (1.42277) [–2.04209]	0.146714 (0.52505) [3.27943]
R-Squared	0.927029	0.841547	0.987464	0.722001
Adj. R-squared	0.927029	0.841547	0.987464	0.722001
F-Statistic	0.005246	0.569643	0.074480	0.027485
Log-likelihood	7.259471	3.034873	45.01251	0.271388
Akaike IC	52.56864	3.682636	20.73118	32.69362
Schwarz IC	–7.428107	–1.947106	–2.121863	–4.115604
Log-likelihood		1047.5907		
Akaike Information Criterion		–19.26512		
Schwarz Information Criterion		–17.97204		
Number of coefficients		32		

Standard errors in ( ) & *t*-statistics in [ ]

Recall that our VAR specification has four ($k = 4$) endogenous variables, MMR, CRS, LIQ, and GDP, and includes lags 1 to 2 ($p = 2$). Thus, there are ($kp = 8$) regressors in each of the four equations in the VAR. The coefficient results are displayed in Table 8. Each column in the table corresponds to an equation in the VAR, and each row corresponds to a regressor in the equation. Note that the regressors are grouped by variable, so that all the lags for the first variable, here

MMR, are followed by all the lags for the second variable, CRS, and so on. The exogenous variables appear last.

For each right-hand-side variable, Table 8 reports the estimated coefficient, its standard error, and the t-statistic. For example, the coefficient for MMR (–1) in the GDP equation is –0.964185, the standard error is 2.80596, and the corresponding t-statistic is –5.34362. The table also displays additional information below the coefficient results. This information concerns summary statistics for the VAR system. These statistics include the determinant of the residual covariance, log-likelihood and associated information criteria, and the number of coefficients.

Looking a little more closely, we note that all the estimated coefficients in the VAR model are correctly signed and statistically significant, with an error probability of 5%. We accept the hypothesis that the defining variable is significant, since the result of the Student test is greater than +/– 1.96, (cf. t-statistics in [] in Table 8). Furthermore, the adjustment quality of the VAR model as measured by the determination coefficient (R^2 adjusted) is quite high, standing at 92%, 84%, 98%, and 72% for the MMR, CRS, LIQ, and GDP equations respectively.

5. CONCLUSION

In this article we consider a sample of Tunisian banks during the 1993 to 2016 period. The variables of interest consist of the concentration index for deposit banks, and the monetary policy shocks in the monthly data of 27 universal and business banks covering the same period. The credit supply was selected from the general balance sheet of all banks each month during the 24 years from 1993 to 2016. The concentration index and the monetary policy shocks were measured. The concentration index was calculated for each year for a panel-data set of 27 banks and 2 semesters over the 1993–2016 period. The monetary policy shocks are associated with the monetary market rate residual function issued by the VAR model.

The most interesting feature of this paper's analysis of the effect of monetary policy shocks on credit supply is its consideration of foreign banks. When foreign banks are present, the relationship between monetary policy shocks and credit

supply changes in form and significance. The change in the form of the relationship shows that two situations have to be selected. The first is when banks' competitiveness is less than the threshold value (0.005) that is obtained by taking the partial derivative of the credit supply function with respect to competitiveness: a monetary policy shock is significantly and negatively related to credit supply. However, this relationship changes from negative to positive if competitiveness exceeds the threshold of 0.005. This suggests that there are strong threshold effects, in that foreign bank entry tends to enhance interest rate pass-through. The paper also concludes that when foreign bank entry leads to greater banking concentration, the extent of interest rate transmission is significantly lowered.

REFERENCES

Adams, R.M. & Amel, D.F. (2005). The effects of local banking market structure on the banking-lending channel of monetary policy. Board of Governors of the Federal Reserve System, *Finance and Economics Discussion Series*: 2005–16.

Alencar, L.S. & Nakane, M.I. (2004). Bank competition, agency costs and the performance of the monetary policy. *Working Paper Series*, 81, 1–46.

Bolt, W. & Tieman, A.F. (2004). Banking competition, risk and regulation. *The Scandinavian Journal of Economics* ,106, 783–804.

Bensaid, B., & Palma, A. De. (1995). Politique bancaire et concurrence bancaire. *Annales d'Economie et de Statistique*, 40, 93–123.

Bikker, J.A., & Haaf, K. (2001). Competition, concentration and their relationship: An empirical analysis of the banking industry. *Journal of Banking and Finance*, 26, 2191–2214.

Chen, X. (2005). Bank loan financial liberalization, competition, and quality. *Journal of Economic Integration*, 20, 109–122.

Chan, Y., Greenbaum, S.I. & Thakor, A.V. (1986). Information reusability, competition and bank asset quality. *Journal of Banking and Finance* 10, 243–253.

Chen, X. (2007). Banking deregulation and credit risk: Evidence from the EU. *Journal of Financial Stability*, 2, 356–390.

Chen, X. & Haller, H. (2003). Banking competition and screening of loan applicants. *Working paper, Virginia Tech, Blacksburg, VA*.

Christiano, L.J., Motto, R. & Rostagno, M. (2014). Risk shocks. *American Economic Review*, 104, 27–65.

Demirgug-Kunt, A., & Detragiache, E. (2002). Does deposit insurance increase banking system stability? An empirical investigation. *Journal of Monetary Economics*, 49, 1373-1406.

Ellizalde, A. & Repullo, R. (2004). Economic and regulatory capital for financial conglomerates. *Working paper, research series supervision*, 45.

Gapalan, S. & Rajan, R.S. (2017). Does foreign bank presence affect interest rate pass-through in emerging and developing economies? *Journal of Macroeconomics*, 54, 373–392.

Gropp, R., & Vesala, J. (2004). Deposit insurance and moral hazard. *Review of Finance*, 8, 571–602.

Gunji, H., Miura, K., & Yuan, Y. (2009). Bank competition and monetary policy. *Japan and the World Economy*, 21, 105–115.

Gehrig, T. (1998). Competing markets. *European Economic Review*, 42 (2), 277–310.

Gunji, H., Miura, K. & Yuan, Y. (2017). Competition among regional banks in Japan: Evidence from the Boone indicator. *Journal of International Economic Studies*, 31, 65–85.

Hellmann, T., Murdock, K. & Stieglitz, J. (2000). Liberalization, moral hazard in banking and prudential regulation: Are capital requirements enough? *American Economic Review*, 90, 147–165.

Hacker, R. S. & Hatemi-J, A. (2008). Optimal lag-length choice in stable and unstable VAR models under situations of homoscedasticity and ARCH. *Journal of Applied Statistics* 35 (6): 601–615.

Honohan, P. (2000). Banking system failures in developing and transition countries: Diagnosis and prediction. *Review of Banking , Finance And Monetary Economics*, 29, 83–109.

Ida, H., Martin, W. & Tomasz, W. (2018). How does financial liberalization affect the influence of monetary policy on the current account? *Journal of International Money and Finance*, 85, 93–123.

Marquez, R. (2002). Competition, adverse selection, and information dispersion in the banking industry. *Revue of Financial Studies*, 15, 901–926.

Maudos, J. & Fernandez, J. G. (2004). Factors explaining the interest margin in the banking sectors of the European union. *Journal of Banking and Finance*, 28, 2259–2281.

Manove, M., Padilla, A.J. & Pagano, M. (2001). Collateral versus project screening: a model of lazy banks. *Journal of Economics*, 32, 726-744.

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

Stiglitz, J. & Greenwald, B. (2003). Towards A New Paradigm In Monetary Economics. *Cambridge University Press*.

VanHoose, D. (1985). Bank market structure and monetary control. *Journal of Money, Credit, and Banking,* 17, 298–311.

Yeyati, E. L. & Micco, A. (2007). Concentration and foreign penetration in Latin American banking sectors: Impact on competition and risk. *Journal of Banking and Finance*, 31, 1633–1647.

Received: December 18, 2019
Accepted: December 14, 2020