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ORIGINAL RESEARCH

Growth, bank credit, and inflation in Mexico: evidence from an ARDL-bounds testing approach

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Abstract This paper explores the long-run effects of inflation on the dynamics of private sector bank credit and economic growth in Mexico over the period 1969–2011. With an ARDL-type model, the statistical results suggest that the availability of private sector bank credit in the economy exerts a positive impact on real GDP. In addition, inflation rates have contributed negatively to the increase in private credit, liquid liabilities, and financial development. A key outcome is that one percent increase in inflation is associated with a 0.07 % fall in long-run real rate of output through its effect on bank credit to the private sector. Another crucial finding is that policies of financial liberalization have helped stimulate economic growth. Reinforcing the literature on finance and growth, this study reaffirms that inflation rates are detrimental to long-run financial development and economic growth.

Keywords Inflation · Private sector bank credit · Financial development · Growth

JEL Classification E31 · G21 · E44 · O4 · C22

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Abbreviations

ADF	Augmented Dickey–Fuller test
AIC	Akaike information criterion
ARCH	Autoregressive conditional heteroskedasticity
ARDL	Autoregressive distributed lag
CUSUM	Cumulative sum
CUSUMQ	Cumulative sum squared
ECM	Error correction model
GDP	Gross domestic product
GFDD	Global financial development database
KMO	Kaiser–Meyer–Olkin coefficient
KPSS	Kwiatkowski–Phillips–Schmidt–Shin test
ISI	Import-substitution industrialization
OLS	Ordinary least squares
PCA	Principal component analysis
PP	Phillips–Perron
SBC	Schwarz Bayesian criterion

1 Introduction

In many developing countries policies were implemented in the 1980s and 1990s with the promise of enhancing financial systems, so that they could allocate resources efficiently to productive investments of innovative entrepreneurs and increase long-run growth (Fry 1997). Although it is well known that in developing countries banks dominate financial systems and at the same time are the main financing source, it is still a puzzle as to why they do not engage more aggressively in lending to small and medium-sized firms (Freeman and Click 2006).

Mexico is not the exception. Although the country began to liberalize the financial sector in the late 1980s, its intermediaries are still underdeveloped (De la Torre et al. 2012; Jiménez and Manuelito 2011). In the finance and growth literature, banking sector depth is measured mostly by bank credit to the private sector and deposit money bank assets as a percentage of gross domestic product (GDP) (King and Levine 1993a, b). Accordingly, in 1961 in Mexico the value of the first indicator was 20.49 % and of the second 21.21 %; whereas 50 years later bank credit to the private sector went down to 18.26 %, bank assets increased to 33.93 % (Global Financial Development Database 2013). Financial depth of Mexican banks is even worse than that of smaller economies such as Guatemala and Colombia. For instance, in 1961 their private sector bank credit as a percentage of GDP was 10.07 and 20.15, respectively, and in 2011 the values went up to 22.58 and 31.97. The same story repeats for bank assets: 10.20 and 20.50 % in 1961, and 35.15 and 39.78 % in 2011 (Global Financial Development Database 2013).

Some authors argue that in Mexico the lack of bank loans is associated to weak bankruptcy provisions (Hanson 2010; Bergoeing et al. 2002; Kehoe and Ruhl 2010);

overall macroeconomic conditions significantly affecting loan supply or demand, through lending capacity and interest rates, respectively (Barajas and Steiner 2002); a banking sector with a reduced number of institutions (Haber 2009); or falling demand for loans even if interest rates have declined steeply in the last decade (Ríos and Salazar 2012). However, none of these authors address the issue analyzed in the present paper, which is the effect of inflation in determining the persistent low rates of bank credit to the private sector in the Mexican economy over the period 1969–2011.

We assume that the dynamics of inflation rates have distorted bank managers' capacity to assess properly a firm's potentially profitable investment, thus reducing the allocation of resources to the real economy. It is important to investigate private sector bank credit, because it has been pointed out as a key factor in explaining the country's low investment rate and thereby economic growth (Moreno-Brid and Ross 2009)[1], as well as inflation since it has been demonstrated that in Mexico higher inflation rates lead to greater uncertainty and lower growth (Grier and Grier 2006).

To prove our hypothesis, the empirical approach used here is the Autoregressive Distributed Lag (ARDL) cointegration model. The estimated equations include four models to investigate any negative effects of inflation on growth through private sector loans and other financial variables such as bank assets, liquid liabilities, and an index of financial development that is built with the variance of the first three variables using the principal component analysis. The estimations also control for other macroeconomic variables, such as government consumption, gross fixed capital formation, exports of goods and services, and a financial liberalization dummy.

The main empirical findings can be summed up as follows: (a) inflation is directly detrimental to economic growth and indirectly through bank credit to the private credit; (b) real output is also lower if inflation slows down the process of financial development; and (c) financial liberalization or the elimination of financial repression is positively associated to growth. In practical terms to understand the economic size of the inflation coefficient, one percent increase in the inflation rate leads to 0.07 % fall in GDP through its adverse impact on bank loans to the private sector. Moreover, if the overall process of financial development is taken into account, inflation rates lower GDP growth rates by 0.11 %.

The paper is organized as follows: In Sect. 2 we briefly describe related literature. Section 3 presents our empirical strategy, the hypothesis testing process of cointegrating relationships and the description of data. Section 4 discusses the main econometric results and the robustness check. Conclusions from this research are summarized in Sect. 5.

2 Economic and financial background

One major concern of policymakers in developing countries is to control inflation rates, which are highly disruptive for all economic sectors. If the economic

[1] Bank credit is a major worry for Mexican authorities. On September 10, 2013 the Mexican Congress approved a new financial reform bill. One key objective of the bill is to promote and increase bank and non-bank credit at lower costs. For more details, see SHCP (2013).

environment is inflationary, firms and households find it difficult to make adequate investment or consumption decisions because inflation reduces planning periods and damages confidence. In the literature of inflation and growth, there are two main channels through which inflation reduces the real rate of investment: the level channel and the productivity channel (Temple 2000; Barro 1995; Fischer 1993).

With respect to the level channel, in some exogenous growth models inflation may shrink the steady-state capital stock. In Tobin (1965) inflation could induce higher levels of capital stock, while Fischer (1979) uses the Tobin effect to argue that the anticipation of inflation produces changes in capital stock. Moreover, Stockman (1981) asserts that if inflation rates were higher, individuals would decrease their real money holdings because money is more expensive to accumulate. In such setting it is unwise to invest more, given that it would imply additional money holdings, higher costs, and therefore lower returns on investments.

In relation to the productivity channel, De Gregorio (1993) establishes that "[a] reduction in the rate of inflation by half is found to increase GDP per capita growth by 0.4 per year" in twelve Latin American countries during the 1950–1983 period. However, Jones and Manuelli (1995) found modest empirical support for changes in monetary policy (inflation) on long-run output through investment decisions[2]. Finally, Sirimaneetham and Temple (2000) demonstrate that over a period of 30 years an enhancement of one standard deviation in their index of macroeconomic stability could increase annual output by around 0.5–0.7 % points.

It is frequently contended that higher inflation pushes up costs of transactions and information. For instance, firms are unwilling to enter long-term contracts when they are uncertain if the evolution of future prices erodes incoming real cash flows. From the financial intermediaries perspective, it is reasonable to believe that bank managers will also be reluctant to grant loans to firms or enter longer-term contracts when they are unable to assess properly investment projects and risks. For example, higher inflation translates into higher unit production costs and, as a consequence, higher prices of products. Since consumers are less willing to buy more expensive products, there will be a fall in sales and the present value of cash flows.

Naturally, financial markets are one channel through which inflation might restrict investment productivity and hence real growth[3]. The financial liberalization theory sentences that inflationary finance retards the development of financial systems, thereby provoking lower levels of investment and reduced capital productivity, unless governments introduce measures to liberate financial markets and promote free-market competition (Shaw 1973; McKinnon 1973; Roubini and Sala-i-Martin 1992). In an inflationary environment, private agents usually avoid longer-term contracts because of possible cost raises related to interest rate

[2] Money may not be the only determinant of long-run inflation. Van der Ploeg and Alogoskoufis (1994) show in an endogenous growth model with overlapping generations, money in the utility function, and inelastic labor supply, that demand-side effects have real effects; for instance, both public debt and consumption hurt growth.

[3] King and Levine (1993a, b) were among the first researchers to show theoretically and empirically a causality running from finance to growth. Nonetheless, there is another strand in the literature that proves a bi-directional or reverse causality. For an extensive review on the subject, see Ang (2008a).

uncertainty. In addition, financial intermediaries would prefer to keep liquid portfolios to avert undue risks.

Under the endogenous growth framework, recent studies have reached the following consensus[4]:

1. Steady state inflation leads to diminished real activity (Huybens and Smith 1999), because inflation distorts the credit allocation process and deteriorates credit quality when the financial sector is unable to distinguish good borrowers from bad ones (De Gregorio and Sturzenegger 1994)[5]. Baum et al. (2006, 2009) prove that under higher and volatile inflation (macroeconomic uncertainty) banks and firms display homogenous behavior in the sense of reducing supply and demand for loans, respectively (see also Talavera et al. 2012).
2. There exists strong negative association among inflation, bank credit to the private sector, bank assets, and bank liabilities (Boyd et al. 2001; Rosseau and Wachtel 2002; Rosseau and Yilmazkuday 2009; Huang et al. 2010; Bittencourt 2011; Choi et al. 1996)[6]. Furthermore, it has been found a positive link between inflation and interest rates which in turn inhibits bank loan demand (Calza et al. 2006; Ibrahim and Shah 2012).
3. The relationship between inflation and finance is nonlinear (Lee and Wong 2005; Keho 2010), which means that at low inflation levels there are growth-enhancing effects arising from financial development—the effects might be reduced or nullified when inflation increases to more than 15 %.

In sum, the literature of finance and growth has firmly established theoretically and empirically a negative impact of inflation on growth through bank credit under either exogenous or endogenous growth models.

3 Strategy for cointegration and data issues

This section presents a cointegration method to demonstrate any long-term relationship among economic growth, private sector bank credit, and inflation in Mexico over the period 1969–2011. We apply the ARDL-bounds testing approach model developed by Pesaran and Shin (1996); Pesaran and Smith (1998) and Pesaran et al. (2001).

For our purposes, the ARDL approach to cointegration has three advantages with respect to the two most popular approaches, namely the Engle–Granger two-step method and Johansen's system-based reduced rank regression method. First, cointegration can be carried out even if variables are $I(0)$, $I(1)$, or mutually cointegrated (Pesaran and Shin 1996; Pesaran and Smith 1998). Thus the ARDL approach is suitable for econometric models that combine level and growth

[4] For an early survey of inflation and financial market performance, see Boyd and Champ (2003).

[5] In contrast, Hung (2003) develops and endogenous growth model of a three-period-lived overlapping generations and information imperfections within financial markets. In the model, adverse selection causes credit rationing, and financial development could increase inflation rates if in the initial status of equilibrium they were already high. In the end, economic growth will be reduced.

[6] See Andrés et al. (2004) for an application to a sample of developed countries.

variables (for instance, inflation with GDP, government consumption, exports, etc.). Second, cointegration is possible even when independent variables are endogenous. The method computes accurate long-run parameters and valid t-values; moreover, the endogeneity bias tends to be irrelevant and very small (Ang 2008b; Inder 1993). And third, in small sample sizes (more than 30 observations) the estimates of the short-run model are highly consistent with their respective long-run parameters, and therefore inferences are based on standard normal asymptotic theory (Ang 2008b).

There are two stages in the estimation of the ARDL model. The first stage consists in verifying the optimal number of lags for the first difference of variables with the Akaike Information Criterion (AIC) or the Schwarz Bayesian Criterion (SBC). The optimal and sufficient lag structure is a fundamental test in ARDL models to eliminate any endogeneity problems (Pesaran and Smith 1998). The second step refers to testing the existence of cointegration. First, the coefficient of the Error Correction Model (ECM) must be negative, which indicates that exogenous variables return to long-run equilibrium levels. And second, Pesaran et al. (2001) computed critical t-values (lower and upper bounds) to test the validity of cointegration, whereas Narayan (2005) calculated critical value bounds for the F statistic when the sample size is smaller than 80 observations, as it is done in this study. In both cases cointegration is accepted when the corresponding critical values lie above the upper bounds.

To investigate the main hypothesis posed in this article, the following model may be estimated:

$$\begin{aligned} \mathrm{GDP}_t = & \beta_0 + \beta_1 \mathrm{GOV}_t + \beta_2 \mathrm{XP}_t + \beta_3 \mathrm{INV}_t + \beta_4 \mathrm{INF}_t + \beta_5 \mathrm{FD}_t \\ & + \beta_6 \mathrm{INF} \times \mathrm{FD}_t + \mu_t, \end{aligned} \tag{1}$$

where GDP_t is gross domestic product; GOV_t is general government consumption; XP_t is exports of goods and services; INV_t is gross fixed capital formation; FD_t is a measure of financial development that assumes either of the following variables: private sector bank credit, bank assets, liquid liabilities, or an index of financial development; INF_t is a measure of inflation (average annual change in the consumer price index); and μ_t is the error term.

Based on the bounds-testing approach proposed by Pesaran and Smith (1998) and Pesaran et al. (2001), any long-run relationship may be given by the equation

$$\begin{aligned} \Delta \mathrm{GDP}_t = & \alpha_0 + \sum_{j=0}^{p} \beta_j \Delta \mathrm{GDP}_{t-j} + \sum_{j=0}^{p} \gamma_j \Delta \mathrm{GOV}_{t-j} + \sum_{j=0}^{p} \phi_j \Delta \mathrm{XP}_{t-j} \\ & + \sum_{j=0}^{p} \zeta_j \Delta \mathrm{INV}_{t-j} + \sum_{j=0}^{p} \eta_j \Delta \mathrm{INF}_{t-j} + \sum_{j=0}^{p} \varphi_j \Delta \mathrm{FD}_{t-j} \\ & + \sum_{j=0}^{p} \psi_j \Delta \mathrm{INF}_t \times \mathrm{FD}_{t-j} + \sigma_1 \mathrm{GDP}_{t-1} + \sigma_2 \mathrm{GOV}_{t-1} + \sigma_3 \mathrm{XP}_{t-1} \\ & + \sigma_4 \mathrm{INV}_{t-1} + \sigma_5 \mathrm{INF}_{t-1} + \sigma_6 \mathrm{FD}_{t-1} + \sigma_7 \mathrm{INF} \times \mathrm{FD}_{t-1} + \mu_t, \end{aligned} \tag{2}$$

where p is the optimal lag length and Δ refers to the first difference of variables.

The hypotheses for testing the existence of any long-run cointegration among the proposed variables in this paper are as follows:

$$\begin{aligned} H_0: &\quad \sigma_1 = \sigma_2 = \sigma_3 = \sigma_4 = \sigma_5 = \sigma_6 = \sigma_7 = 0 \\ H_1: &\quad \sigma_1 \neq 0, \sigma_2 \neq 0, \sigma_3 \neq 0, \sigma_4 \neq 0, \sigma_5 \neq 0, \sigma_6 \neq 0, \sigma_7 \neq 0 \end{aligned} \tag{3}$$

That is, the joint null hypothesis of no cointegration against the existence of a valid relationship between GDP_t and the set of explanatory regressors.

Lag orders were selected using AIC because results are usually better and more consistent than utilizing other information criteria (Lütkepohl 2006). Once we reject the null hypothesis of no cointegration, we can proceed to estimate the short-run model by approximating the ECM:

$$\begin{aligned} \Delta\mathrm{GDP}_t = \quad & \alpha_0 + \sum_{j=0}^{p} \beta_j \Delta\mathrm{GDP}_{t-j} + \sum_{j=0}^{p} \gamma_j \Delta\mathrm{GOV}_{t-j} + \sum_{j=0}^{p} \phi_j \Delta\mathrm{XP}_{t-j} \\ & + \sum_{j=0}^{p} \zeta_j \Delta\mathrm{INV}_{t-j} + \sum_{j=0}^{p} \eta_j \Delta\mathrm{INF}_{t-j} + \sum_{j=0}^{p} \varphi_j \Delta\mathrm{FD}_{t-j} \\ & + \sum_{j=0}^{p} \psi_j \Delta\mathrm{INF} \times \mathrm{FD}_{t-j} + \theta \mathrm{ECM}_{t-1} + \mu_t, \end{aligned} \tag{4}$$

where ECM_{t-1} is the error correction term that in turn is equal to:

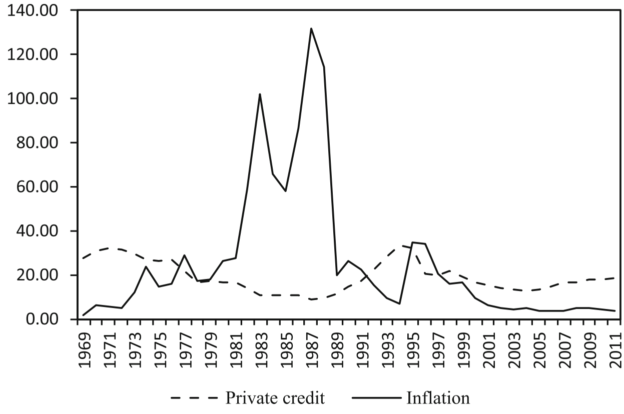
$$\begin{aligned} \mathrm{ECM}_{t-1} = \quad & \mathrm{GDP}_{t-1} - (\alpha_0 + \sigma_1 \mathrm{GOV}_{t-1} + \sigma_2 \mathrm{XP}_{t-1} + \sigma_3 \mathrm{INV}_{t-1} \\ & + \sigma_4 \mathrm{INF}_{t-1} + \sigma_5 \mathrm{FD}_{t-1} + \sigma_6 \mathrm{INF} \times \mathrm{FD}_{t-1}) \end{aligned} \tag{5}$$

The coefficients β_j, γ_j, ϕ_j, ζ_j, η_j, φ_j, and ψ_j represent the short-run dynamics of the variables, while the coefficients $\sigma_i (i = 1,\ 2,\ 3,\ 4,\ 5,\ 6)$ indicate the long-term dynamics. The term θ is the coefficient of correction in disequilibrium.

3.1 Data issues

As mentioned above, the estimation period spans from 1969 to 2011. In this period the Mexican economy evolved from an import-substitution industrialization model (ISI) until 1982 and a market-oriented model since then. After dismantling ISI in the early 1980s when the economy collapsed because of an unfortunate combination of expansionary monetary and fiscal policies, sharp deterioration in international oil markets, strong capital outflows, and the peso devaluation (Lustig 2002), the government nationalized the banking system and gradually introduced tougher measures of financial repression such as selective credit policies and reserve requirements. By the end of the decade, the authorities undertook several economic and financial reforms to open up the economy to foreign competition in goods and capital markets, as well as to liberate domestic financial markets from government intervention. Furthermore, in the early 1990s the banking system was reprivatized and in 1997 it was sold to foreign investors in an effort to reinforce the system's efficiency and capacity to intermediate more funds in the economy.

Figures 1 and 2 show the performance of private sector bank credit, financial development and average inflation rates. First, it is observed that bank loans decreased sharply during the 1980s, increased a little more in the next decade and fell again in the 2000s. In the figure we similarly see a strong inverse relationship between inflation and bank credit from 1969 to 1997, and from then on a minor rise



Source: GFDD (2013).

Fig. 1 Evolution of private sector bank credit and average inflation rates, 1969–2011

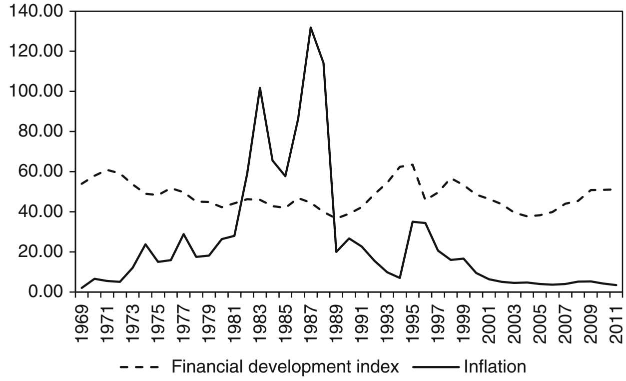


Fig. 2 Evolution of financial development and average inflation rates, 1969–2011

when inflation rates were stabilized under the new inflation-targeting regime imposed by the Central Bank. As for the overall financial development activity, we observe the same negative correlation throughout most of the period. In our econometric estimations we expect to prove statistically the negative link between inflation and bank credit.

Table 1 presents the definition of variables used in this study. Above we stated that the dependent variable is GDP_t, which is measured as the natural log of real gross domestic product. Moreover, we include a set of macroeconomic variables (the expected sign is in parentheses), namely government consumption (+), gross fixed capital formation (+), exports of goods and services (+), and a financial liberalization dummy, LIB_t (+). According to standard practice, the first three variables are also in natural logs. The dummy variable captures the overall liberalization process as follows: 1 if $t = 1969 - 1987$ (financial repression) and 0

Table 1 Definition of variables and data sources

GDP	Natural log of real domestic product. Base year is 2005. Source: WDI (2013)
GOV	Natural log of general government consumption expenditure in constant terms. Base year is 2005. Source: WDI (2013)
INV	Natural log of gross fixed capital formation in real terms. 2005 = 100. Source: WDI (2013)
XP	Natural log of exports of goods and services in real terms. Base year is 2005. Source: WDI (2013)
INF	Average annual increase in the consumer price index. 2005 = 100. Source: WDI (2013)
PC	Natural log of private sector bank credit in real terms, 2005 = 100. It is calculated with the following formula:
	{(0.5) × [PCt/P_et + PCt−1/P_et−1]}/[GDPt/P_at] where PC is credit to the private sector, P_e
	is end-of period CPI, and P_a is average annual CPI. Source: Global Financial Development Database (2013)
BA	Natural log of Deposit money bank assets in constant terms, 2005 = 100. It is computed using the following formula:
	{(0.5) × [BAt/P_et + BAt−1/P_et−1]}/[GDPt/P_at] where BA is deposit money bank claims, P_e
	is end-of period CPI, and P_a is average annual CPI. Source: GFDD (2013)
LL	Natural log of broad money or M3 in constant terms. It is calculated using the following formula:
	{(0.5) × [LLt/P_et + LLt−1/P_et−1]}/[GDPt/P_at] where LL is liquid liabilities, P_e
	is end-of period CPI, and P_a is average annual CPI. Source: Global Financial Development Database (2013)
FD	Financial development index that we calculated with the method of principal components.
	The index is a weighted average of three variables: private sector bank credit, deposit money bank
	assets and deposit money bank liabilities, all three extracted from Global Financial Development Database (2013)
INF × PC	Interaction variable of inflation and private sector bank credit
INF × BA	Interaction variable of inflation and bank assets
INF × LL	Interaction variable of inflation and liquid liabilities
INF × FD	Interaction variable of inflation and financial development index

otherwise (liberalization). The data were compiled from the *International Financial Statistics* of the International Monetary Fund, the *World Development Indicators* of the World Bank and the *Global Financial Development Database* of the World Bank as well.

The set of financial variables is defined as follows:

PC = natural log of bank credit to the private sector (+).

BA = natural log of total bank assets (+).

LL = natural log of bank liquid liabilities (+).

FD = financial development index (+).

In order to capture the effects of inflation on bank credit and real output, we created four interaction terms of inflation and finance that were included in the regressions accompanied by their respective source variable, for instance PC_t and

$\mathrm{INF}_t \times \mathrm{PC}_t$. As recommended by Ozer and Sørensen (2013), in a specification with interaction terms, we should always include the main variables since the interaction effect might be significant because of the ‘left-out variable bias’. Moreover, the regressions would not have collinearity problems between main variables and interaction variables if their nature were different, for example PC_t and $\mathrm{INF}_t \times \mathrm{PC}_t$ or LL_t and $\mathrm{INF}_t \times \mathrm{LL}_t$. Accordingly, out set of interaction indicators and their expected signs are $\mathrm{INF}_t \times \mathrm{PC}_t(-)$, $\mathrm{INF}_t \times \mathrm{BA}_t(-)$, $\mathrm{INF}_t \times \mathrm{LL}_t(-)$, and $\mathrm{INF}_t \times \mathrm{FD}_t(-)$.

Our measure of financial development was built combining the variables PC_t, BA_t, and LL_t. We applied the method of principal component analysis (PCA) to build FD_t. In particular, the method computes the principal components that aim at capturing the variance from source series. Usually, the first principal component contains the highest variance from the original series. Before proceeding with the PCA, we have to check the factorability of variables with the Bartlett’s test for sphericity and the Kaiser–Meyer–Olkin (KMO) coefficient. The Bartlett’s test converts the calculated determinant of the matrix to a χ^2 statistic, which is then tested for significance. In the test, the null hypothesis is that variables are non-collinear. On the other hand, the KMO indicators compare the size of variables’ correlation coefficients to the size of the partial correlation coefficients. In the latter test, a minimum value of 0.60 is necessary for an acceptable PCA. Results from both the specification tests and the PCA appear in Table 2, from which we can infer that the three variables may be assembled into another set of factors using the PCA. Hence, the values of the first PCA are used to calculate the weights for the financial development index (see also Fig. 2).

To verify the applicability of the ARDL bounds method, we employed three tests to evaluate the order of integration of variables (Table 3). Both the Augmented Dickey–Fuller and Phillips–Perron tests are employed to check for the null hypothesis of the existence of a unit root and the Kwiatkowiski–Phillips–Schmidt–Shin test verifies the null hypothesis of stationarity in variables. In Table 3 we see that all variables are $I(0)$, $I(1)$, or a combination of both, and that none is integrated of higher order. Therefore, based on these results we are allowed to apply the ARDL technique.

3.2 Testing for the hypothesis $\sigma_1 = \sigma_2 = \ldots = \sigma_7 = 0$

In this subsection we estimate whether Eq. (2) shows any degree of long-run cointegration among the proposed variables. We proceed in two steps to perform a Wald test for the joint null hypothesis using the F statistic, i.e. the coefficients of the level variables are zero as in $\sigma_1 = \sigma_2 = \ldots = \sigma_7 = 0$. In the first step we specify a restricted form of Eq. (2) by replacing the level variables with their first difference. For instance, we estimate an ordinary least squares (OLS) regression in first difference of the dependent variable GDP_t against the independent regressors GOV_t, INV_t, XP_t, INF_t, PC_t, $\mathrm{INF}_t \times \mathrm{PC}_t$, and LIB_t. In the second step we add the lagged terms of the variables to the previous result. The new OLS output gives the value of the F statistic that is represented by $F\ (\mathrm{GDP}_t|\mathrm{GOV}_t,\ \mathrm{INV}_t,\ \mathrm{XP}_t,\ \mathrm{INF}_t,\ \mathrm{PC}_t,\ \mathrm{INF}_t \times \mathrm{PC}_t, \mathrm{LIB}_t)$.

Table 2 Construction of financial development index with principal components

(a) Tests for factorability	
Determinant of the matrix of correlation	0.64
Bartlett's test for sphericity	17.896 (0.000)*
Kaiser–Meyer–Olkin measure	0.628

Number	Value	Difference	Proportion	Cumulative value	Cumulative proportion
(b) Principal component analysis					
Eigenvalues: (sum = 3, average = 1)					
1	1.779808	1.058529	0.5933	1.779808	0.5933
2	0.721279	0.222366	0.2404	2.501087	0.8337
3	0.498913	–	0.1663	3.000000	1.0000

Variable	PCA 1	PCA 2	PCA 3
Eigenvectors			
Private credit, PC	0.592926	−0.477843	−0.779316
Bank assets, BA	0.517591	0.842761	0.255323
Liquid liabilities, LL	0.616878	−0.582972	0.572257

Source: authors' calculation

* Statistically significant at 1 % level

P value is in parentheses

To accept or reject H_0, we compare the calculated F statistic with the critical value bounds obtained by Narayan (2005). The preceding process is repeated for each of the remaining three models.

Table 4 displays the results from the testing procedure. We used one and two lags, AIC, SBC and intercept. In all four models, we observe that the calculated F statistics are above the critical value bounds in first lag only, which provides enough evidence to reject the joint null hypothesis of no long-run cointegration. For example, in Model A the calculated F statistic is 24.912, a value that is well above the 1 % upper-bound critical value of 5.464 with $k = 7$ ($k = 8$ is not available on the tables). In the rest of the models, the null hypothesis is strongly rejected as well. However, in Model B the calculated F statistic is significant at the 10 % level for two lags, implying that in this specification the explicative variables may be viewed as 'long-run forcing variables' because the null hypothesis is weakly rejected. Therefore, from the F-test results we conclude that there exists a long-cointegration between GDP_t and the corresponding exogenous variables.

3.3 Testing for the hypothesis $\sigma_t = 0$

In the second part of the hypothesis testing process, we examine the value of the calculated t statistics against the upper bound critical values estimated by Pesaran et al. (2001). The calculated t statistic is obtained from the OLS regression with variable additions that was estimated in the first part of the testing process. In the

Table 3 Unit root test

Variables	ADF		PP		KPSS		Decision
	Level	1st diff.	Level	1st diff.	Level	1st diff.	
GDP	−2.706***	−4.794*	−2.552	−4.794*	0.807*	0.377***	I(0)/I(1)
GOV	−7.417*	−3.268**	−7.417*	−3.131**	0.751*	0.621**	I(0)/I(1)
INV	−1.834	−6.393*	−1.976	−6.393*	0.433***	0.310	I(1)
XP	−1.104	−4.398*	−0.995	−4.370*	0.811*	0.131	I(1)
INF	−1.624	−6.654*	−2.317	−5.841*	0.210	0.393***	I(1)
PC	−1.513	−3.218**	−1.339	−3.566**	0.495**	0.070	I(1)
BA	−3.847*	−6.330*	−2.728***	−6.552*	0.070	0.190	I(0)/I(1)
LL	−2.668***	−4.088*	−1.962	−3.993*	0.111	0.089	I(1)
FD	−2.867***	−4.972*	−1.917	−4.732*	0.139	0.084	I(1)
INF × PC	−5.023*	−5.411*	−5.035*	−28.574*	0.169	0.346	I(0)
INF × BA	−1.692	−6.929*	−2.371	−7.925*	0.224	0.500**	I(1)
INF × LL	−1.486	−7.014*	−2.185	−7.147*	0.235	0.247	I(1)
INF × FD	−1.615	−6.909*	−2.407	−8.553*	0.259	0.500**	I(1)

AIC was used for ADF to select the lag length; the maximum number of lags was set to five. Barlett–Kernel was used for PP and KPSS, as the spectral estimation method

*, ** and *** are statistically significant at 1, 5 and 10 %, respectively

example of the regression given above, we would look at the t statistic of the lagged dependent variable $\mathrm{GDP}_t(-1)$.

Calculated t statistics for the four models appear in Table 4 where it is also seen that all t statistics are above the upper-bound critical values for one lag. For instance, in Model A the calculated t statistic is −5.108 and the corresponding 1 % upper value critical bound with $k = 8$ is −5.07; therefore, we reject the null hypothesis of no long-run cointegration among the variables. Moreover, for Model B we also found weak evidence of long-run cointegration with two lags that is consistent with the Wald test carried out in the first part of the testing process. In any case, the hypothesis testing process confirms that the long-run determinants of inflation, bank credit, and economic growth may be estimated with at least one lag in the exogenous variables.

4 Empirical results

Our preliminary exploration of data indicate a negative association between inflation and bank loans, as well as the other two financial variables that are part of the financial development index. First, to examine in more detail the empirical correlation, we estimate four models of economic growth, bank credit and inflation with the interaction terms $\mathrm{INF}_t \times \mathrm{PC}_t$, $\mathrm{INF}_t \times \mathrm{BA}_t$, $\mathrm{INF}_t \times \mathrm{LL}_t$, and $\mathrm{INF}_t \times \mathrm{FD}_t$. Second, for the purpose of robustness checking we build a parsimonious error-correction model for the main model A to assess the statistical significance of the

Table 4 Testing for long run cointegration: F statistic and t statistic (Dependent variable: GDP)

Lag	Model A		Model B	
	(GOV, INV, XP, INF, PC, INF × PC, LIB)		(GOV, INV, XP, INF, FD, INF × FD, LIB)	
	$P = 1$	$P = 2$	$P = 1$	$P = 2$
F statistic	24.912*	2.888	34.294*	2.9833
t statistic	−5.108*	−2.576	−9.423*	−2.660
Lag length selection criteria				
AIC	−5.442	−5.826	−5.013	−5.472
SBC	−4.724	−5.066	−5.472	−5.504
Lag	Model C		Model D	
	(GOV, INV, XP, INF, BA, INF × BA, LIB)		(GOV, INV, XP, INF, LL, INF × LL, LIB)	
	$P = 1$	$P = 2$	$P = 1$	$P = 2$
F statistic	25.147*	2.925	24.2905*	2.379
t statistic	−8.304*	−2.900	−6.690*	−0.644
Lag length selection criteria				
AIC	−5.523	−4.954	−5.323	−5.446
SBC	−4.805	−5.440	−4.606	−4.686

Critical value bounds of the F statistic with $k = 7$ with constant ($k = 8$ is not available): (3.644, 5.464), (2.676, 4.130), and (2.260, 3.534) at the 1, 5, and 10 % level of significance, respectively

Critical value bounds of the t statistic with $k = 7$ with constant: (−2.58, −5.07), (−1.95, −4.43), and (−1.62, −4.09), respectively

*, **, and *** statistically significant at 1, 5, and 10 %, respectively

P is the lag length

ECM_{t-1} term and the long-run stability of regression coefficients with the cumulative sum (CUSUM) and the cumulative sum squared (CUSUMSQ) tests.

4.1 Long-run coefficients

Table 5 reports our main empirical findings of the estimated long-run coefficients for the four different versions of the economic growth equation. All versions include the set of control variables of macroeconomic determinants (GOV_t, INV_t, XP_t, INF_t and LIB_t) and for each model a different finance variable and its respective interaction term are added (PC_t, BA_t, LL_t, and FD_t). In all equations the liberalization binary variable is positive and highly significant at the 1 % level. It is important to note that all equations as well pass the Breusch–Godfrey test of serial correlation, functional form test, and the autoregressive conditional heteroskedasticity (ARCH) test. All the long-run coefficients are statistically significant with the exception of the coefficient of INV_t (it is insignificant in all specifications and does not have the expected sign), the coefficient of liquid liabilities, and the inflation rate in models B and D. In regard to the investment's coefficient, its insignificance may be related to the inclusion of financial variables, because they may be capturing

Table 5 Long-run determinants of growth, finance and inflation in Mexico

Regressor	Model A ARDL (1, 0, 0, 1, 0, 0, 1, 1)	Model B ARDL (1, 0, 0, 1, 1, 1, 1, 0)	Model C ARDL (1, 1, 0, 1, 1, 1, 1, 0)	Model D ARDL (1, 1, 0, 1, 0, 0, 0, 0)
Speed of adjustment (ECM_{t-1})	−0.724* (−9.119)	−0.371* (−4.463)	−0.532* (−6.351)	−0.400* (−4.646)
Intercept	7.756* (8.833)	11.047* (5.806)	12.08* (10.559)	7.04* (3.863)
Government consumption, GOV_t	0.352* (8.662)	0.526* (6.442)	0.488* (8.693)	0.621* (7.326)
Gross investment, INV_t	−0.009 (−1.213)	−0.010 (−0.590)	−0.033* (−2.678)	0.007 (0.376)
Exports, XP_t	0.203* (17.796)	0.165* (5.948)	0.195* (10.278)	0.19* (6.567)
Inflation, INF_t	−0.099* (−3.308)	0.320 (1.024)	−1.078* (−2.322)	0.248 (1.608)
Private credit, PC_t	0.260* (6.231)	–	–	–
Financial development, FD_t	–	−0.025*** (−1.792)	–	–
Bank assets, BA_t	–	–	−0.142* (−4.934)	–
Liquid liabilities, LL_t	–	–	–	−0.019 (−0.585)
Inflation × private credit, $INF \times PC_t$	−0.070* (−7.540)	–	–	–
Inflation × financial development index, $INF \times FD_t$	–	−0.106** (−2.020)	–	–
Inflation × bank assets, $INF \times BA_t$	–	–	0.298* (2.049)	–
Inflation × liquid liabilities, $INF \times LL_t$				−0.267* (−2.967)
Liberalization dummy, LIB_t	0.178* (9.891)	0.119* (3.312)	0.751** (2.894)	0.220* (4.335)
Adjusted R-squared	0.791	0.810	0.811	0.685
Diagnostic tests (LM version)				
Breusch–Godfrey	0.131 (0.718)	2.423 (0.120)	0.087 (0.769)	0.036 (0.849)
Functional form	0.889 (0.346)	0.273 (0.601)	0.948 (0.330)	0.634 (0.426)

Table 5 continued

Regressor	Model A ARDL (1, 0, 0, 1, 0, 0, 1, 1)	Model B ARDL (1, 0, 0, 1, 1, 1, 1, 0)	Model C ARDL (1, 1, 0, 1, 1, 1, 1, 0)	Model D ARDL (1, 1, 0, 1, 0, 0, 0, 0)
ARCH	0.001 (0.991)	1.966 (0.161)	0.038 (0.845)	0.023 (0.878)

Based on the Akaike Information Criterion and intercept only. The period of analysis is 1969–2011. Maximum lag = 1

t statistics are in parentheses. $n = 43$

p values of diagnostic tests are in brackets

*, **, and *** are statistically significant at 1, 5, and 10 %, respectively

some positive effects on economic growth.[7] Lastly, overall the coefficients of the four models have the expected sign and the error-correction term is negative and highly significant at the 1 % level. When the inflation rate is statistically significant, it has the expected negative sign. The value of the speed of adjustment term ranges from about −0.371 to −0.724.

Results from Model A confirm our hypothesis about the negative effects of inflation on private credit and economic growth as well as the hypothesis stated by the finance and growth literature in the sense that financial depth, as proxied by private sector bank credit, encourages both the level and rate of long-run real output. According to the estimation results and the insertion of interactions, an increase of 1 % in private credit generates a rise of 0.26 % in the long-run rate of economic growth. In this model, the inflation rate has a negative impact on output of around 0.099 % for every percentage increase in prices. In addition, inflation rates in Mexico have diminished the positive effects of bank loans to the private sector. An increase in long-run inflation rates of 1 % decreases real economic growth by 0.07 % through their effects on private credit. In addition, as underlined by the literature, financial liberalization has had positive effects on the Mexican economic growth.[8]

Our results differ from some studies. Employing econometric methodologies of cointegration analysis and causality tests—although different time periods—, Rodríguez and López (2009) and Arestis and Demetriades (1999) suggest that in Mexico there is bi-directional causality between financial development and growth. Contrarily, Bandiera et al. (2000) argue that in Mexico financial repression encouraged financial development and helped increase private savings. However, our results are consistent with Venegas-Martínez et al. (2009) in relation to private sector bank loans only, because they demonstrate a positive influence—albeit small—of financial development on growth and an inverse (positive) association between financial repression (liberalization and real output).

On the other hand, in Model B we modified the previous specification to include the financial development index and its interaction effect with inflation rates. The coefficient of financial development is around −0.025 (weakly significant at the 10 % level) and its interaction with inflation has a coefficient's value of −0.106. In comparison with private sector bank loans, the overall process of financial development has had negative effects on the economy's growth. This result deserves some further explanation. Possibly the index is capturing higher variance arising from the inclusion of bank assets and the broad money measure of M3. Its negative sign also points out to an inadequate bank structure and insufficient deposits that have reduced the efficiency of the banking system in intermediating and allocating resources to the economy. In fact, Ahmed et al. (2008) examined the finance-growth in Mexico with panel data (fixed and random effects methods) from 1971 to 2000

[7] Although not reported here, we run some alternative regressions with different transformations of the investment variable and we were unable to find acceptable long-run cointegration.

[8] We also run some regressions with the inverse of the financial liberalization dummy that represents financial repression, but we do not report results here. Unfortunately, such regressions caused invalid cointegration relationship because, according to the literature as well, financial repression is strongly associated to inflation.

and a Cobb–Douglas-type production function. The authors imply that it is probable that domestic credit could have been used to finance unproductive or speculative investments.

Finally, in Models C and D we report the results from the two other components of the financial development index, specifically bank assets and liquid liabilities. In both cases their coefficients are negative, but LL_t is statistically insignificant. In this sense, bank assets have been insufficient to encourage long-run economic growth, because an increase of 1 % in bank assets diminishes real rates of output by 0.142 %; in fact, an increase of 1 % in $INF_t \times BA_t$ leads to a 0.298 % in economic growth, which is inconsistent with BA_t's negative coefficient. The last result agrees with Boyd et al. (2001)'s findings who stress that the negative link between inflation and bank assets disappears when the inflation rate exceeds 15 % (during 1969–2011 the average inflation in Mexico was 25.7 %). Finally, in the two models the financial liberalization dummy preserves its positive and statistically significant sign.

4.2 Parsimonious error-correction model

According to Bahmani-Oskooee (2001), we can estimate a parsimonious dynamic relationship based on the ARDL model and AIC. The goal is to verify the stability of long-run coefficients in the growth equations along with the short-run dynamics represented by the coefficients of the error-coefficient model. In this section we estimate parsimonious error-correction model for our preferred specification (Model A).

First, we generate a series for the error-correction term using the long-run coefficients. Second, we estimate an OLS regression in first difference of the dependent variable ΔGDP_t against the regressors, including lagged variables of the dependent variable and the error term. And third, we employ CUSUM and CUSUMSQ tests to check for the stability of the long-run coefficients. The coefficients are said to be stable if both statistics remain within 5 % significance level or mainly in between the two outer lines drawn in the plot.

The two error-correction terms for Models A and B are obtained with the equations:

$$\begin{aligned} ECM_t \quad &= GDP_t - 0.352 \times GOV_t + 0.009 \times INV_t - 0.203 \times XP_t + 0.099 \times INF_t \\ &-0.26 \times PC_t + 0.07 \times (INF_t \times PC_t) - 0.0178 \times LIB_t - 7.756 \end{aligned} \tag{6}$$

Results from the estimations are shown in Table 6 and in Fig. 3. Most coefficients are significant at the 1 % level and they pass the Breusch–Godfrey, functional form and heteroskedasticity tests as well as the Wald tests with the exception of GDP_t and INV_t. Also, the ECM_{t-1} is significant at the 1 % level. Finally, according to the CUSUM and CUSUMSQ plots, the long-run coefficients are stable.

5 Concluding remarks

Recent studies have not been able to explain why in Mexico bank credit to the private sector have remained stuck in the last decades. In addition, some authors

Table 6 A parsimonious ECM of growth, bank credit and inflation in Mexico (Dependent variable: ΔGDP)

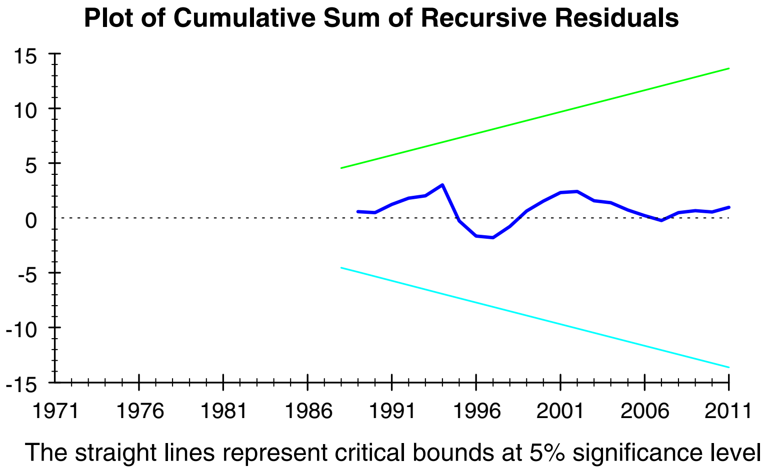
Regressor	Model A
Intercept	−0.020* (−3.810)
Gross domestic product, ΔGDP_{t-1}	0.009 (0.118)
Government consumption, ΔGOV_t	0.278* (4.066)
Gross investment, ΔINV_t	−0.007 (−0.990)
Exports, ΔXP_t	0.212* (6.256)
Inflation, ΔINF_t	−0.056* (−3.078)
Private credit, ΔPC_t	0.211* (5.798)
Inflation × private credit, $\Delta INF \times PC_t$	−0.040* (−5.968)
Liberalization dummy, ΔLIB_t	0.024* (1.515)
ECM_{t-1}	−0.677* (−7.298)
Adjusted R-squared	0.848
Durbin's h-statistic	−0.405
Standard error of regression	0.014
Breusch–Godfrey, $\chi^2(1)$	0.319 (0.572)
Functional form	0.461 (0.497)
Test for heteroskedasticity, $\chi^2(1)$	2.922 (0.087)
Wald tests	
H_0 = coefficient on $\Delta GDP_{t-1} = 0$	0.014 (0.906)
H_0 = coefficient on $\Delta GOV_t = 0$	16.536 (0.000)
H_0 = coefficient on $\Delta INV_t = 0$	0.98 (0.322)
H_0 = coefficient on $\Delta XP_t = 0$	39.14 (0.000)
H_0 = coefficient on $\Delta INF_t = 0$	9.48 (0.002)
H_0 = coefficient on $\Delta PC_t = 0$	33.62 (0.000)
H_0 = coefficient on $\Delta INF \times PC_t = 0$	35.61 (0.000)
H_0 = coefficient on $\Delta LIB_t = 0$	2.29 (0.130)
H_0 = coefficient on $ECM_{t-1} = 0$	53.26 (0.000)

* and ** are significant at 1 and 5 %, respectively

p values are in parentheses

have underlined that the lack of bank loans have constrained the level and efficiency of investment in the country. In this paper we address this question by proving statistically an explanation of how under macroeconomic instability, in particular through inflation rates, bank managers are unwilling to grant more loans to firms.

Our main contribution is that we assess empirically the effects of inflation on private sector bank credit and economic growth in Mexico over the period 1969–2011. Using the ARDL-bounds testing approach, we estimated the long-run effects of inflation rates through bank loans on real output in the long run. Moreover, we built a parsimonious error-correction model to test the stability of long-run coefficient and the statistical significance of the short-run error term. Among the principal results, we determined that bank credit exerts a positive and



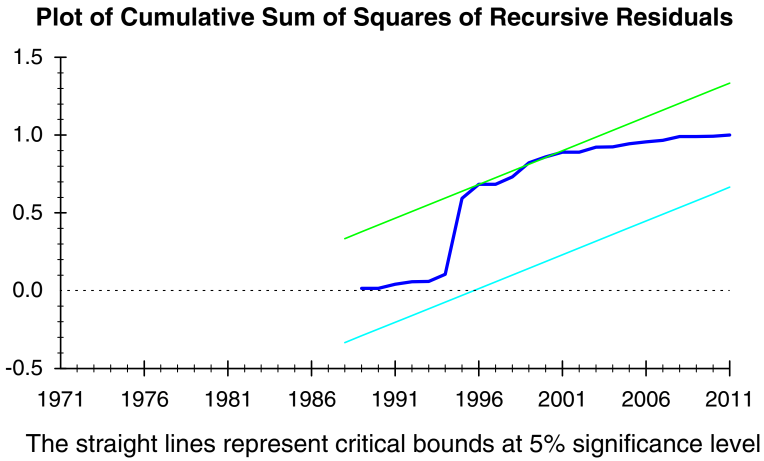


Fig. 3 CUSUM and CUSUMSQ tests of growth, private credit and inflation

significant influence on growth; inflation rates have diminished growth through their adverse effects on private credit, as indicated by related literature; financial liberalization has promoted the development of the Mexican financial sector; and not all aspects of financial development foment long-run real output. Also, most of the variables used in this study have the expected sign, except investment, which turned out to be insignificant and had a negative sign, and bank assets that had a positive sign.

Our results have several policy implications and suggestions for future research. First, policy makers should still promote intensively the deepening of the financial system, in particular the banking sector, as in the recently approved financial reform bill in Mexico. Second, the Central Bank should continue enforcing sound policies to control inflation rates like in the past decade where they were kept at around 4–5 % annually. As implied by the literature, a stable macroeconomic environment is a fundamental prerequisite for healthy financial sector development. Third, in the past decade there was a trivial increase in bank loans to the private sector under a stable macroeconomic environment, a phenomenon that is not yet well explained in the literature. It remains as a future line of research to understand why Mexican banks have been unable to grant more loans when inflation rates are low and stable. Finally, we are aware that even if we obtained econometrically-consistent estimates, there are modern techniques such as the Bayesian approach that could serve as an additional robustness check or to respond the questions mentioned in the research agenda.

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