# THE DEMAND FOR BANK LOANS IN VENEZUELA: A Multivariate Cointegration Analysis

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**Abstract:** In order to disentangle the stagnant behavior of the credit market in Venezuela this study focuses on the evolution and specific role played by the demand for loans. Using monthly time series we estimate a credit demand function for the period 1986:1 to 2000:12. The estimation is based on a minimum theoretical specification which sees credit demand as driven by firm's financial decisions. Cointegration tests indicate that there is one stationary long-run relationship between the real stock of loans, an index of real sales, the interest rate on loans, the real exchange rate and the economy mark-up. The short-run dynamics of the demand for real loans is subsequently modeled by means of a Vector Error Correction Model. The general-to-specific methodology has served to restrict the error correction vector, thereby obtaining two final dynamic models to explain the short-term movements of real credit. The results are consistent with the view that the interest is exogenous and the quantity of bank borrowing is largely demand determined.

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# 1. Introduction

Macroeconomic literature has amassed a great deal of studies on money demand and its determinants, but interest in the role of the credit market and of credit demand has been relatively scarce. There are surprisingly few studies on the demand for bank credit and its determinants for industrialized countries, and interest in developing economies is quite recent.[1] As Fase (1995) has recently emphasized, such scarce attention to bank credit in comparison to the strong focus on the money is even more intriguing in view of the fact that the first econometric analyses of bank credit were published in the early thirties of the last century by Tinbergen (1934, 1937), much before the famous studies by Brown (1938, 1939) on the determinants of money demand.

However, in recent times the perception of the credit market's importance in understanding many macroeconomic and financial problems has been changing. A major reason for that is the acknowledgment of the credit market as a key factor in understanding the transmission channels and of monetary policy, as well as in understanding the mechanics of money creation. Actually, interest in the credit market is not new in economic analysis. For instance, for some years proponents of the endogenous nature of money have called attention to the role of banks in attempting to accommodate the variations in credit demand. From this viewpoint the Central Bank controls short-term (inter-bank) interest rates, and thereby the cost for banks of obtaining liquidity, and banks set loan interest rates as a mark-up of the official rate. Economic units borrow from banks, create deposits and bank money. Banks come to the wholesale market when they need liquidity. Interest rate changes on the wholesale liquidity market affect the rest of the rates and consequently, the willingness of

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[1] Empirical studies on the behavior and determinants of credit demand fall basically into four categories: First, those that estimate credit demand as a system. Indeed, the possibility of endogeneity coming from the interest rate frequently leads studies to estimate a system using two reduced equations (one for credit demand and another for the interest rate). This is very much the approach followed by Melitz and Pardue (1973), Heremas *et al.* (1976) Catao (1997) Friedman and Kuttner (1993) and Fase (1995). There are other studies that establish *a priori* that at any point in time the volume of loans corresponds to the level at which credit supply and demand are equal, and estimate the determinants of bank credit under equilibrium conditions, rather than a credit demand function. The simultaneous solution of credit supply and demand equations generates a single-equation model where both supply and demand factors appear as explanatory variables. Hendershott (1968), Hicks (1980) and Panagopoulos and Spiliotis (1998) follow this approach. A third approach (the disequilibium approach) uses a maximum likelihood method to estimate separate functions of credit supply and demand and compare their estimated values with the actual data in order to determine whether the market is supply or demand constrained. Following the pioneering studies by Laffont and Garcia (1977) and by Sealy (1979), during the last decade several studies such as Blundell-Wignal and Gizycki (1992), Pazarbasioglu (1997), Ghosh and Ghosh (1999), Literas and Legnini (2000), Barajas, López and Oliveros (2001) and Barajas and Steiner (2001) estimate loan supply and demand using the likelihood function derived by Maddala and Nelson (1974). Finally, some studies attempt to identify the demand curve by isolating those variables that affect one side of the market only (identification by parameters restrictions) and using as an auxiliary hypothesis the idea that banks operate in an environment of imperfect competition and that credit to corporations is determined by the demand side of the market at the interest rate set by the banks. Goldfeld (1969), Harris (1976), Moore and Theadgold (1985), Cuthbertson (1985), Arestis (1987), Arestis and Biefang-Frisancho (1995), Howells and Hussein (1992) and Calza, Gartner and Souza (2001) use this approach.

economic units to borrow from the financial system. But in the final analysis, credit demand (private decisions) is what determines the course of money supply.[2]

In the context of developing countries, the importance and predominance of bank loans for the process of accumulation have been in sharp contrast with their role in industrialized economies. In the early 1990s, for instance, the World Bank estimated that close to 55% of public and private investment in the developing world was financed by internal corporate resources. Such a broad self-financing design could be sustained in many economies due to the protection mechanisms that ensured high returns in real markets. Hence, in perspective, opening up the markets, as well as the pressures exerted by greater competition should have redirected the financing structure of the real sector towards bank loans, thereby encouraging further development of the credit market.

This lesser presence of domestic credit in the financing structure of private economic activity may be attributed to permanent shocks from the supply side of loans, and/or to disturbances from the demand side. Therefore, adequate control of supply changes with a minimum theoretical model that allows identifying the determinants of demand is required.  This study proposes to give adequate answers concerning the driving forces of the outstanding lending activity and to determine the specific role of credit demand in the evolution of this market. For that purpose this paper examines and specify a reliable credit demand function for Venezuela. The case of Venezuela is interesting in many ways since the credit market shows no signs of increasing strength in spite of having undergone opening and deregulation throughout the past few years.

From an empirical viewpoint, the paper shows that it is possible to specify a reliable credit demand function in Venezuela and to define the impact of its determinants, in the long term as well as in the process of dynamic adjustment towards long-term equilibrium.

The paper consists of seven sections. Following the introduction, section II presents some stylized facts about the behavior of the credit market in Venezuela. In section III we present a simple theoretical model, which derives from the financing constraints and corporate investment function of the firm. Section IV explains the statistical series to be used and provides a simple correlation analysis between the logarithm of credit in real terms and a set of twelve possible relevant independent variables. Section V tests the assumption that variables are non-stationary, and verifies the existence of unit roots in the series. The co-integration methodology is applied using the methods of both Engle and Granger, and Johansen, yielding a long-term equilibrium relationship from January 1986 to December 2000. The results lead to an interpretation that may be consistent with the specification of the theoretical model. Section VI attempts to obtain a dynamic representation of the credit demand function, using a vector error correction model (VECM) and applying the "the general-to-

[2] Modern contributions to the endogenous money theory and to the importance of credit demand in the creation of base money can be track back to Kaldor (1970), Davidson (1977), and Moore (1979), but in the last two decades the topic has witnessed increasing work, refinements and discussions among post-Keynesian scholars.

specific" methodology for the purpose of restricting and obtaining the best specification. Two error correction models with very subtle differences are obtained and subject to diagnostic and stability tests. The paper ends with some brief conclusions.

## 2. Some Stylized Facts about the Evolution of Credit in Venezuela

A preliminary exercise, for the purpose of characterizing the evolution of credit in the Venezuelan economy, is based on the availability of monthly and quarterly figures of the real stock of credit of commercial banks from 1986 to 2002. Figure Nº 1, for instance, shows that the quarterly evolution of outstanding commercial and universal bank loans to the private sector as a proportion of GDP expanded during the 1980s. This expansion ceased at the end of the decade, and a marked downward trend began that extended beyond the years of the 1994-1995 financial crisis. The year 1997 shows a slight recovery, followed by stagnation until the year 2000. Remarkably, the Credit-GDP ratio of the Venezuelan economy, that during the 1980s was even higher than that of its Latin American peers, became one of the lowest of the continent by the end of the 1990s.[13]

Figure 1
Stock of Credit as a Percentage of GDP
1980-2000

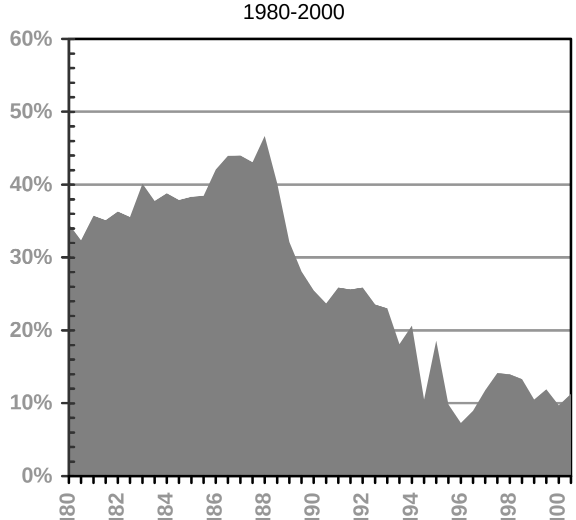
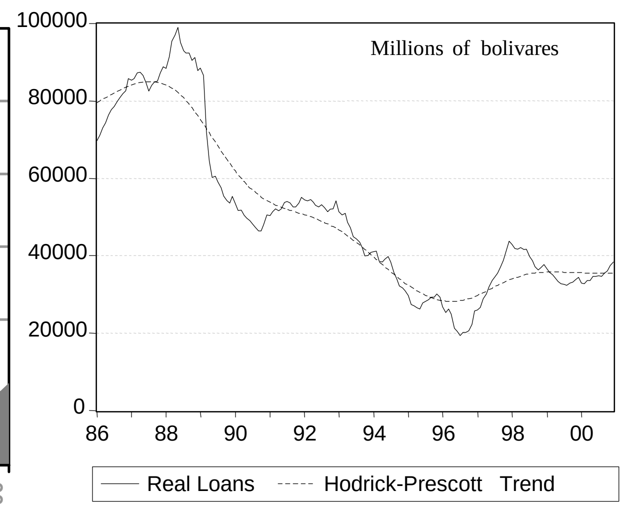



Figure 2
Evolution and Trend of the Real Stock of Credit
1986:1 - 2000:12




Calculations of the monthly credit stock in real terms and its trend value (using the Hodrick-Prescott filter) showed an almost uninterrupted drop since 1988 until mid-1996 that dragged real credit down to 20% of its peak value. Its temporary recovery in 1997 barely compensated this drop (see figure 2). The same pattern stands out when quarterly data are used in analyzing the evolution of per-capita real stock of credit (see figure 3). Indeed, by the end of 2000 real credit in per capita terms was barely 28% that of mid-1988.

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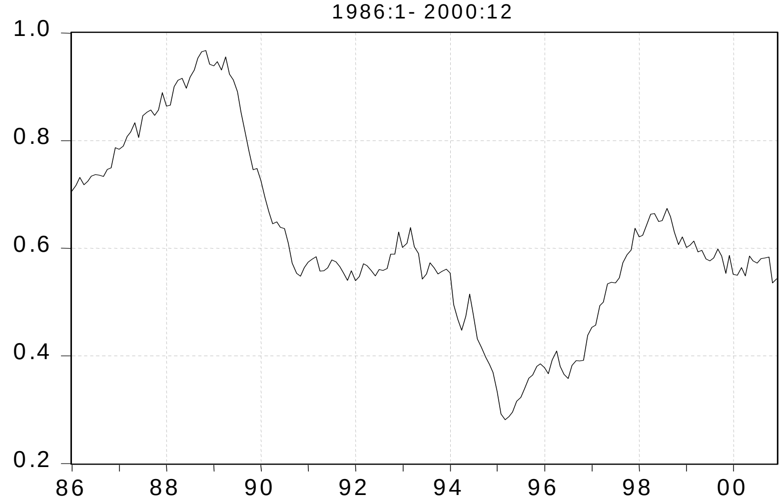
[3] This is clearly seen when the credit-GDP ratio calculated here is compared with that shown by Barajas and Steiner (2002) for several countries of the region.

The gradual contraction of credit to the private sector was accompanied by growing intermediation towards the public sector. As a glance at Figure 3 reveals, gradual changes in composition of bank assets occurred throughout the past decade. The growing share of investment in securities and domestic government bonds went from barely 7% at the end of 1988 to nearly 17% in November 1990, peaking at 38% of the nominal amount of bank assets in April 1995, when the economy was undergoing the effects of the banking crises.

Figure 3
Credit/Deposit Ratio
1986:1- 2000:12




Source: Banking system balance sheets

It is difficult to determine the extent to which this gradual substitution of credit to the private sector for credit to the public sector was due to portfolio decisions of the private agents (including the banks), which in last instance respond to factors associated with the risk, returns and comparative liquidity of the instruments. However, there is no doubt that the more active development of the market for bonds placed by the monetary authority and by the public sector throughout the 1990s provided more options to financial investors, at a time when the Venezuelan economy had been experiencing strong macroeconomic volatility and substantial imbalances.[4]

The question that this relatively simple facts rises is what are the factors driving for real credit in Venezuela? We will show that the demand side of the credit market plays a major role in this history confirming the post-Keynesian preoccupations and focus on firms as a source of credit and new money.

4 Radical changes began in late 1989 with the appearance of the well-received Zero Coupon Bonds. The stock of these bonds climbed to an equivalent of 7% of GDP in mid-1991 (García, Rodríguez and Salvato, 1997). Fiscal expansion, based on the temporary oil-derived resources resulting from the Persian Gulf conflict (that began in August 1990) forced the Central Bank to make much stronger use of its bonds to drain liquidity. Moreover, as a result of the dramatic drop of credit and higher liquidity risk that the banks had to face during the years of the financial crisis, the banking system turned *en masse* to Central Bank securities. Currently, the delicate fiscal situation of the Venezuelan economy, and the country's limited access to foreign markets have made domestic financing from the banking system the most immediate source of resources in recent years to cover the fiscal gap, as well as a liquid and relatively low-risk alternative for investors.

## 3. A Theoretical Specification of a Credit Demand Model for Venezuela

Independently from the empirical estimation method, few studies of the credit demand function formally justify the inclusion of certain explanatory variables. What predominates is the *ad-hoc* introduction of variables based on intuitive reasoning, with a remarkable diversity of explanatory variables. Hicks (1980) in this regard points out that even when the variable to be included is generally accepted, disagreements arise out of not knowing whether they should be included as levels or as first differences.[5] In order to reduce the risk of misspecification, we will require a minimum theoretical model for the estimation of a credit demand function.

Following the recent post-Keynesian tradition it will be assumed that banks operate under conditions of imperfect competition, and set the lending rate on the basis of the rate that provides access to the wholesale market of liquid resources plus a risk-related mark-up. Once the rate is set, banks attend to the demand and establish collateral, term and structure of payments contractually with the client. Demand thus determines the stock of credit at an exogenous interest rate, $r_L$. These assumptions are not only consistent with widespread modern banking practice of overdraft facilities and liability management, but also convenient to the objectives of this paper by making it easier the identification of a demand function.

It is further assumed that the corporate agent applying for credit acts in the framework of a financially open economy, and therefore explores and openly evaluates the possibility of converting their liquid resources into fixed investment or into external financial assets. This relatively ignored aspect of the credit demand problem, however, may be very important in open economies such as that of Venezuela, subject to highly mobile capital flows and ongoing capital flight throughout the period considered here.

We will assume that in order to make investment decisions, corporations face the following financing constraint,

$$I + \Delta B^{*} = CF + \Delta CR + \Delta P \qquad (1)$$

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5 There are very obvious disagreements, even using the more traditional determinants of credit demand. In the case of the economic activity variable, for instance, some studies in line with arguments by Kashyap, Stein and Wilcox (1983) prefer to emphasize the positive relation between economic activity and credit demand, on the basis that more economic activity has positive effects on income and expected profitability, hence in the financial situation of households and businesses, who would be willing to increase their debt in order to promote higher levels of consumption and investment. In contrast, other studies (following Benanke and Gertler, 1995) indicate that during expansionary phases, businesses may prefer to rely more on internal sources of finance and reduce the relative proportion of external financing. The relation between cost of financing and credit demand causes equally complex disagreements. First, there is relatively little accord as to the rate (or rates) that should be used. A second problem is to decide whether the rate should be taken in real or in nominal terms, with a separate term that would quantify the impact of the expected inflation rate (see the discussion in Cuthbertson, 1985). An additional problem is whether the rate should be taken in log form, in level form, in first differences, or as a deviation from another rate.

In equation (1), firms are equating total uses of funds for investment with total sources of funds. Thus, corporations must decide on the magnitude of their investment outlays in real assets, $I$, and on the magnitude of their investment in external financial assets, $\Delta B^*$. Firms finance these decisions from three sources of funds: cash flow, $CF$, variations in bank loans, $\Delta CR$, and/or variations in paper issuance, $\Delta P$. Eventually, it is also possible to finance the increase in the position of one asset by liquidating another.[6]

The limited deepness of private securities markets and the small importance of issuance of commercial papers or shares in Venezuela justify the low-risk assumption that $\Delta P = 0$.

Equation (1) is a simple accounting relation with no theoretical content and subject to several hypotheses regarding the behavior of firms. For instance, it could be assumed that increases in the real stock of credit, $\Delta CR$, may be accompanied by increases in $\Delta B^*$ without changing the restriction of equality, in which case corporations would be increasing their credit demand to finance the purchase of external financial assets. Under other conditions variations in the real position of external assets, $\Delta B^*$, could be related to increases in the position of fixed assets, in which case investment decisions are financed by liquidating external assets. Relative profitability or risk decisions could generate this portfolio change. Another possibility could be to finance investment with increases of the real credit demand. It is even possible to associate changes in the financing composition of the corporations, whereby an increase in the cash flow, $CF$, would be associated with a reduction of credit demand, $\Delta CR$. Then, it is clear that for the financing constraint of corporations to be operational under any theory, a model must be specified and constructed on the basis of some of the above behavioral assumptions.

Henceforth, corporative demand of external financial assets, $\Delta B^*$, will depend on two elements: (a) the expected return on those assets, $r^*$, and (b) the financing gap of the corporations, $I\text{-}CF$, which is the difference between the required resources for the investment, $I$, and the portion available from internal financing, $CF$.

$$\Delta B^* = \Delta B^*(r^*, I - CF) \qquad (2) \quad \text{where} \quad \frac{\partial \Delta B^*}{\partial r^*} > 0 \quad \text{and} \quad \frac{\partial \Delta B^*}{\partial (I - CF)} < 0$$

On the one hand, equation (2) indicates that an improvement in the expected return on foreign financial assets lead corporations to acquire financial claims against non-residents. On the other hand, to the extent that the investment/internal financing gap becomes wider, part of it could be financed by liquidating positions in foreign currency, which explains the inverse relation between $\Delta B^*$ and $(I - CF)$.

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[6] Corporations may be assumed as final applicants for government securities also, in which case investment decisions are broader. However, our institutional characterization of the Venezuelan market in this case assumes that the main applicants and holders of these government securities are the banks, and not private non-financial agents. We should note that by the end of 2000 the stock of investments in securities of the financial system in Venezuela represented nearly 60% of the government's registered public domestic debt stock.

We will assume that real credit demand depends directly on three factors: (a) the cost of credit, $r_L$ , (b) the financing gap, and (c) variations in the real position of external assets, $\Delta B^*$. In other words,

$$\Delta CR = \Delta CR(r_L, I - CF, \Delta B^*) \quad (3) \quad \text{with } \frac{\partial \Delta CR}{\partial r_L} < 0 \,, \; \frac{\partial \Delta CR}{\partial (I - CF)} > 0 \,, \text{ and } \; \frac{\partial \Delta CR}{\partial \Delta B^*} > 0$$

In equation (3) if the cost of the external debt for corporations increases at the margin, then the flow of real credit demand decreases. Again, if there is no change in the cost and relative risk of the different sources of financing, a wider financing gap ($I - CF$), requires more financing and credit demand. Likewise, credit demand may grow with increases in the position of corporate net external assets.

The real investment function is an endogenous variable in the model (which makes the financing gap likewise partially endogenous). Real investment is assumed as determined by: (a) the GDP level of the economy, $Q$; (b) the cost of capital (given by the real interest rate, $r_L$); (c) the corporate cash flow, $CF$; (d) the risk status of the economy, $\sigma$, and (e) the return on competing assets, in this case the foreign financial assets, $r^*$.

$$I = I(Q, r_L, CF, \sigma, r^*) \tag{4}$$

The incorporation of these variables in equation (4) requires a minimum justification. Investment responds positively to output or any index of economic activity following, for instance the Kalecki-Steindl approach or the Keynesian accelerator theory. In both cases and under certain conditions, the value of the desired stock of capital of a corporate firm is a positive function of output.

The negative relationship between investment and the interest rate is justified in the literature in various ways. One, for instance, assumes that firms "rank" various investment projects depending on their "internal rate of return" (or "marginal efficiency of investment") and thereafter, faced with a given rate of interest, chose those projects whose internal rate of return exceeded the rate of interest. Consequently, investment falls as the interest rate rises and rises when the interest rate falls.

The idea that investment responds positively to increases in corporate cash flows gains support in a considerable number of macroeconomic and financial studies. Though this idea can be track back to the writings of Keynes (1936) and Kalecki (1937), nowadays it has been based on the notion that imperfect information in credit markets. Imperfect information gives rise to divergences in the costs of internal financing and external financing; therefore, the consequent financing constraint make investment very sensitive to variations in the supply of internal funds. Corporate investment will be more sensitive to cash flow changes to the extent that the difference between these costs becomes larger (Fazzari, Hubbard and Petersen, 1988).

The presence of some measure of macroeconomic risk in equation (4) reflects the view that investment is an irreversible decision and that an increase of uncertainty may play an important role in investment decisions. In this case, an inverse relation should be expected between investment and any measure of macroeconomic risk.[7]

Finally, it is proposed that higher returns on financial assets in the rest of the world make the option of investing in fixed assets less attractive.

At a formal level, closing the model now simply requires making explicit the determinants of the return on external financial assets, $r^*$. Indeed, $r^*$ will be determined endogenously as the sum of the foreign currency return on external assets, $i^*$, and the expected exchange rate depreciation, $\hat{e}^e$. Formulating expectations around exchange rate variations is relevant in the case of Venezuela, at least since the early 1980s, when credibility in the system of fixed exchange rate fully disappeared. In this case, it is assumed that the expectation of variations in the nominal exchange rate arises as a result of the deviation between the long-run equilibrium real exchange rate, $\underline{q}$, and the effective real exchange rate, $q$. The resulting representations are:

$$r^* = i^* + \hat{e}^e \tag{5}$$

$$\hat{e}^e = \phi(\overline{q} - q) \tag{6}$$

If the interest rate on external financial assets, $i^*$ and the equilibrium real exchange rate are taken as given, then (5) and (6) imply that the return on external assets is a function of the behavior of the real exchange rate. A real exchange rate appreciation (relative to its long-term value), for instance, will generate expectations of depreciation of the nominal exchange rate and consequently, an increase in the return on the external financial assets will be expected as well.

Substituting (5) into (2) and (4), and then (2) and (4) into (3), will result in a reduced form credit demand function as follows:

$$\Delta CR = \Delta CR(r_L, Q, CF, q, \sigma) \tag{7}$$

where $\frac{\partial \Delta CR}{\partial r_L} < 0$, $\frac{\partial \Delta CR}{\partial Q} > 0$, $\frac{\partial \Delta CR}{\partial CF}$?, $\frac{\partial \Delta CR}{\partial q}$?, $\frac{\partial \Delta CR}{\partial \sigma} < 0$

In theory, equation (7) is a function of five exogenous variables: the interest rate on loans, the level of economic activity, the corporate cash flow, the real exchange rate, and the index of macroeconomic risk.

The signs of the partial derivatives for the interest rate on loans, the macroeconomic risk and the level of activity are evident; not, however, when evaluating the impacts of the cash flow variable and of the real exchange rate. An increase in corporate cash flow reduces the financing gap and consequently credit demand. However, it may also increase investment and the financing gap (increasing credit demand). The sign then will depend much on how sensitive is the investment function to changes in *CF.* In

[7] According to Caballero (1993) investment is irreversible in developing countries as a consequence of severe imperfections in the secondary market of capital assets and the several kinds of adjustment costs.

terms of the impact of the real exchange rate on credit demand, the investment channel (equation 4) must be differentiated from the channel through variations in the position of external financial assets (equation 2). On one hand, an appreciation of the real exchange rate (a fall of $q$) increases the expected return on external financial assets and discourages investment in real domestic assets, reducing the financing gap and credit demand. On the other hand, credit demand may increase in order to finance acquisition of the more attractive external financial assets.

## 4. Data and Correlation Analysis

Initially, equation (7) will be estimated with a set of monthly figures for the Venezuelan economy over a sample period that goes from January 1986 to December 2000. This means, in the best case, a time series analysis comprising a total of 180 observations. In order to preserve variability in the figures as much as possible, the series are not seasonally-adjusted. A logarithmic transformation of the data has been carried out prior for pretesting for correlation analysis, estimation in levels and the presence of unit roots.

Following most studies on credit demand, real credit demand is defined here as the real stock of credit that commercial and universal banks grant to the private sector (as shown on balance sheets for publication) divided by the monthly consumer price index (CPI) with 1984 as the base year.

The interest rate is the value reported by the Central Bank of Venezuela for the monthly average lending interest rate of commercial and universal banks. Its real value is the difference between this rate and the expected inflation rate. In the best case the latter corresponds to the cumulative current inflation rate measured by the CPI's monthly variation with 1984 as the base year. The annualized rate of the recorded monthly inflation has been added as an alternate measure of the expected inflation.

A rate of return on public securities, $r_b$, and the spread that develops between this and the average rate on loans, have been calculated in order not to disregard the some possible systematic relation between the market for government securities and credit demand. In view of the length of the period under study, the series consists of the returns on the most profitable paper on the market at the time, starting with Central Bank Certificates of Deposit from January 1986 to December 1989, following with yield of the Zero Coupon Bonds from January 1990 until December 1999 and ending with National Public Debt Bonds from July 1999 to December 2000.

Two proxies are used as an indicator of economic activity: the GMIEA (General Monthly Index of Economic Activity) published by the Central Bank, and the monthly sales index, also published by the Central Bank, and this latter adjusted for inflation on the basis of the 1984 CPI.

The effective real exchange rate is measured as the ratio between the nominal exchange rate and the consumer price index with 1984 as the base year. It is important to point out that the relevant variable in the theoretical model is the deviation between the long-run equilibrium real exchange rate and the current real exchange rate. But we can impose the strong assumption that the real equilibrium exchange rate is constant throughout the period. An additional limitation is that we have minimized the effect of the evolution of the price index of the rest of the world on the real exchange rate.

The macroeconomic risk affecting investment decisions has been determined to a frequently used conventional measurement, namely, the variance of the last twelve months moving average of the rate of inflation.[8] As an alternative we have also calculated and used the variance of the last twelve months moving average of the GMIEA. This considering that in defining macroeconomic risk, the variability of economic activity could be more important than that of inflation. In spite of the fact that Venezuelan bonds participate in the global market since the 1990s it is very unfortunate indeed that the length of the period under study prevents the use —as in other studies— of an emerging markets bond index (EMBI).

Measuring the cash flow of corporations is a serious drawback in Venezuela due to major limitations in the recollection and acquisition of industrial data. The fact that we use monthly data in this study raises an additional difficulty. However, with due care, the cash flow variable can be approximated calculating the ratio of the consumer price index, CPI, to the wholesale price index, WPI (measured on the same base year). Such an approximation is based on the normal association between the mark-up and the cash flow of corporate firms. If the behavior of the wholesale price index adequately reflects the behavior of the costs, then the CPI to WPI ratio should be a fairly good measure of the mark-up.[9]

The correlation vector between the logarithm of credit in real terms and a set of twelve possible relevant independent variables is shown in Table 1. Even though correlation analysis could not lend support any theory at all, it may well be desirable if the sign and size of the observed covariation provide clues for the selection of variables for some econometric estimation. Table 1, for instance, clearly shows that the correlation index with the real sales index (in logarithmic form) is far superior, to the correlation with the logarithm of the GMIEA. This would indicate that the real sales index would do better as a proxy level of economic activity than the GMIEA. Likewise, the nominal interest rate correlates much better with the real credit than the real interest rate. Indeed, many empirical studies attribute the fail to find any significant relation between the real interest rate and the credit to the difficulty of how to measure the expected inflation rate. It is interesting to note that the inflation rate shows no correlation with real credit (which does not help the notion of separating the nominal interest rate from the rate of inflation in an econometric estimation). Neither the interest rate on public securities, nor the U.S. inter-bank rate, nor the spread of rates between the return on securities and the lending rate correlate clearly with the logarithm of the real credit. The real exchange rate and the cash flow variable correlate well, being acceptable as factors accompanying an econometric estimation.

[8] It is important to point out that observed variability of any macroeconomic variable may not be a valid proxy for uncertainty (which is not directly observable) because it could well has been forecast by economic agents.

[9] Strictly speaking the mark-up would be the difference between the CPI and the WPI divided by the WPI; however, the possibility of obtaining negative values in some observations faces the possibility of using a specification in logarithms.

Macroeconomic risk measures, on the other hand, show practically no correlation with real credit.

Table 1
Correlation Vector between Log(*CR*) and the rest of the variables[1]
1986:1 to 2000:12

| | Ln(*Q*) | Ln(*GI*) | Ln(R) | $r_L$ | $(r_b - r_L)$ | Ln(i*) | Ln(*CF*) | Ln(q) | Ln($\sigma_1$) | Ln($\sigma_2$) | Ln($\pi$) | Ln($r_b$) |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Ln(*CR*) | 0.74 | -0.2 | -0.60 | -0.05 | 0.43 | 0.32 | -0.53 | 0.63 | 0.11 | -0.05 | -0.22 | -0.38 |

[1]/ *Q* = index of real sales, *IG* = GMIEA (General Monthly Index of Economic Activity, R = Average nominal lending rate, $r_L$ = Real lending rate, $(r_b - r_L)$ = spread between the return on government securities and the average rate on loans, i* = Federal Fund Rate (in the U.S.), *CF* = mark-up, *q* = effective real exchange rate, $\sigma_1$ = variability of the inflation rate, $\sigma_2$ = variability of the index of economic activity, $\pi$ = CPI inflation rate, $r_b$ = yields on government securities.

## 5. Unit Roots and Cointegration

Taking into account the theoretical specification of the model (presented above) as well as the degree of correlation between the real stock credit to the private sector and the rest of the variables, we have made a preliminary attempt to estimate the credit demand function in levels.[10] We found, however, that all estimations presented serious problems of serial correlation (with very low Durbin-Watson statistics). Though the inclusion of an AR(1) term in the estimations improved the $R^2$, the serial correlation problem still persisted. An interested aspect of these results is that the combination of a high $R^2$ with a low DW statistic may be an indication of spurious regression, increasing the suspicion that the series contain a stochastic trend, and thus the usual *t* statistics having nonstandard distributions may give seriously misleading inferences.[11]

Nevertheless, we do know that if the variables involved in the estimation have the same order of integration a linear combination of these non-stationary series may be stationary. Therefore, even individual non-stationary variables can form a cointegrating relationship. In order to detect the presence of nonstationarity in the series, we use widely popular unit root tests. In conducting the augmented Dickey-Fuller test (ADF) and the Phillips-Perron (PP) test we have assumed an intercept term in the regressions and the *t* statistics is referred to the MacKinnon critical values. Table 2 gives the results only for those variables that turned out to be nonstationary, that is, those that can be potentially cointegrated.[12]

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10 The underlying assumption here is that the series that we employ appear to behave as if the process that generates them does not contain a unit root. In this case the typical problems of spurious relationships are ignored, but one must guard against their presence.

11 In this case the usual *t* ratio does not possess a limiting distribution but diverges with increasing sample size, thus the variance of the asymptotic approximation to the unknown finite sample distribution can not be defined.

12 Proxys for macroeconomic risk, the GMIEA, and the real interest rate on loans turned out to be all I(0). The same occurs with the rate of inflation when the Phillips-Perron test is applied. This avoid the possibility of a cointegrating relationship between these variables and real loans.

Both unit roots tests the ADF and PP, using the MacKinnon critical values, cannot reject the null hypothesis of nonstationarity in the series for the log of real credit, *CR*, the log of the nominal interest rate on loans, *R*, the log of the real sales index, *Q*, the log of the real exchange rate, *q*, and the log of the cash flow variable, *CF*, at any of the reported significance levels. However, as we can see from Table 2, when variables are taken in first differences, both tests rejects the null in first difference, indicating that all the series contain one unit root and are of integrated order one I(1). As a consequence, the first necessary condition in order to find a cointegrating relationship among the variables is fulfilled.

Table 2
Unit Root Tests*

Null Hipothesis: $x_t$ is not stationary

| Variable | ADF Test | | | Phillips-Perron Test | | |
|---|---|---|---|---|---|---|
| **Ln (*CR*)** | -1.74 | 1% Critical Value** | -3.468 | -1.161 | 1% Critical Value** | -3.468 |
| **DLn (*CR*)** | -3.7 | 5% Critical Value** | -2.878 | -7.821 | 5% Critical Value** | -2.877 |
| | | 10% Critical Value** | -2.575 | | 10% Critical Value** | -2.575 |
| **Ln (*R*)** | -1.997 | 1% Critical Value** | -3.468 | -1.932 | 1% Critical Value** | -3.468 |
| **DLn (*R*)** | -4.754 | 5% Critical Value** | -2.878 | -10.147 | 5% Critical Value** | -2.877 |
| | | 10% Critical Value** | -2.575 | | 10% Critical Value** | -2.575 |
| **Ln (*Q*)** | -1.249 | 1% Critical Value** | -3.468 | -2.107 | 1% Critical Value** | -3.468 |
| **DLn (*Q*)** | -7.546 | 5% Critical Value** | -2.878 | -17.569 | 5% Critical Value** | -2.877 |
| | | 10% Critical Value** | -2.575 | | 10% Critical Value** | -2.575 |
| **Ln (*q*)** | -0.815 | 1% Critical Value** | -3.468 | -0.788 | 1% Critical Value** | -3.468 |
| **DLn (*q*)** | -6.315 | 5% Critical Value** | -2.878 | -15.403 | 5% Critical Value** | -2.877 |
| | | 10% Critical Value** | -2.575 | | 10% Critical Value** | -2.575 |
| **Ln (*CF*)** | -0.743 | 1% Critical Value** | -3.468 | -0.405 | 1% Critical Value** | -3.468 |
| **DLn (*CF*)** | -5.394 | 5% Critical Value** | -2.878 | -10.754 | 5% Critical Value** | -2.878 |
| | | 19% Critical Value** | -2.575 | | 19% Critical Value** | -2.575 |

* The regression includes an intercept term
** MacKinnon critical values for rejection of the hypothesis of a unit root

A number of methods for testing cointegration have been proposed in the literature. We consider the Engle and Granger (1987) method as well as the Johansen and Juselius (1990) method. We follow here the advice given by Gregory (1991) and Gonzalo and Lee (1995) who argue in favor of adopting and reporting several cointegration tests.

We have applied first the Engle-Granger method to the set of nonstationary variables, we have regressed the *Ln*(*CR*) on the set of independent variables (*R*, *Q*, *q* and *CF* in logarithmic form) over the period 1986-2000, and we have performed the unit root test on the residuals and obtained the results presented in Table 3. The statistic is 2.91, which does not fail to reject the null hypothesis (of nonstationarity) at about the 1% level. Hence, the results justifies at best that real loans, the nominal interest rate, the index of real sales, the real exchange rate variable, and the cash flow proxy form a cointegrating vector.

Table 3
Unit Root Test on the Residuals

| Variable | ADF Test | | | Phillips-Perron Test | | |
|---|---|---|---|---|---|---|
| $\varepsilon_t$ | -2.919 | 1% Critical Value* | -2.577 | -2.724 | 1% Critical Value* | -2.577 |
| | | 5% Critical Value* | -1.941 | | 5% Critical Value* | -1.941 |
| | | 10% Critical Value* | -1.616 | | 10% Critical Value* | -1.616 |

* MacKinnon critical values for rejection of the hypothesis of a unit root

Whilst this Engle-Granger approach for testing cointegration is quite simple and straightforward, it does not work all that well in some circumstances. More specifically, the test appears to be biased in finite samples, in the sense that it can reject rather more frequently than the nominal 5% of the time. Moreover, as Banerjee *et al.* (1986) and Inder (1993) have pointed out, in small samples, some risk of bias exists in cointegrating vectors estimated by OLS. One addition limitation is that the Engle-Granger approach does not provide an estimation procedure and tests for all possible cointegrating vectors (in a more than two variable case). These deficiencies are taken into account by the approach introduced by Johansen. Indeed Johansen (1988, 1991), Stock y Watson (1988) and Johansen y Juselius (1990) have proposed a maximum likelihood approach that allows for the estimation of all cointegrating vector, as well as for tests of hypothesis on the cointegrating parameters.[13] In this multivariate approach cointegrating vectors are obtained from the reduced form of a system where all of the variables are assumed to be jointly endogenous. Though cointegrating vectors cannot be interpreted as representing structural equations, cointegration relationships can be due to constraints that an economic structure imposes on the long-run relationship among the jointly endogenous variables.

Johansen and Juselius (1990) specify two likelihood ratio (LR) test statistics to test for the number of cointegrating vectors. First, the likelihood ratio test statistic for the null hypothesis of at most $r$ cointegrating vectors against a general alternative. This test is also called the trace statistics and is given by

$$LR_{trace} = -T\sum_{i=r+1}^{p}\log((1-\hat{\lambda}_i) \qquad (8)$$

where $T$ represents the number of observations, and $\hat{\lambda}_r$ represents the *i-th* largest eigenvalue that is obtained from the determinant of an equation associated with the factorization of an estimated matrix of cointegrating vectors and the associated weighting matrix. The second likelihood ratio test for the null of exactly $r$

[13] Nevertheless, Gregory (1991) shows that the power of the maximum likelihood test is very sensitive to the lag order, which is a disadvantage against the ADF test suggested by Engle and Granger.

cointegrating vetors against the alternative of $r$ +1 is known as the maximum eigenvalue statistic

$$LR_{\max} = -T \log(1 - \hat{\lambda}_{r+1}) \tag{9}$$

Distribution of the test statistics is nonstandard, and approximate asymptotic critical values have to be obtained by simulation. Osterwald-Lenum (1992) gives the most comprehensive set of critical values for VARs with up to 11 variables.

We estimate a vector autoregression model using the same set of non-stationary variables and implement VAR-based cointegration tests using the methodology developed by Johansen.[14] It is important to point out here that the Johansen estimation is highly sensitive to the number of lags included in the vector autoregressive model underlying the data-generating process assumed in this method. Therefore, in order to avoid any *ad-hoc* procedure, we determine first the optimal lag structure according to both the Akaike information criterion (AIC) and the Schwarz criterion (SC). Table 4 shows in line with the AIC the value that minimizes the sum of the square residuals (-15.49). The criteria indicate an optimal lag length is 3. The SC suggests that the inclusion of 2 lags is appropriate. Since overparameterization involves a lower risk of misspecification we choose 3 as the optimal lag length. Additionally, since there seems to be a trend in all the nonstationary series, cointegration tests are conducted with the inclusion of a quadratic deterministic trend.

Table 4
Optimal Lag Structure in the Unrestricted VAR

| **Order of the VAR** | **Akaike Information Criteria** | **Schwarz Information Criteria** |
|---|---|---|
| 1 | -14.81 | -14.28 |
| 2 | -15.49 | -14.5 |
| 3 | -15.46 | -14.72 |
| 4 | -15.4 | -14.5 |
| 5 | -15.37 | -13.02 |
| 6 | -15.35 | -12.53 |

In the framework of the Johansen procedure the number of cointegrating vectors is determined sequentially. We start with the null hypothesis of no cointegration ($r$ = 0) and we continue only if this hypothesis is rejected. Table 5 shows the trace statistic for the cointegrating rank as well as the critical values at 1% and 5% significance level. The LR test rejects the hypothesis of no cointegration but not the hypothesis of at most one cointegrating relation. In other words, Johansen cointegration test results confirm the previous result indicating there is only one stationary long-run relationship between real loans, real sales, interest rate on loans, the real exchange rate and the

[14] The representation is a VAR of order $p$ and is given by $y_t = A_1 y_1 + ... + A_p y_{t-p} + Bx_t + e_t$, where $y_t$ = {ln($CR$), Ln($R$), Ln($Q$), Ln($CF$), Ln($q$)} is the vector of variables I(1), $x_t$ is the vector $d$ of deterministic variables, $A$ and $B$ are coefficient matrices, and $e_t$ is the vector of innovations.

economy mark-up. Formal exclusion tests show that none of the system's variables can be excluded from the cointegrating vector.

Table 5
Johansen Cointegration Test

Sample: 1986:01-2000:12
Series: *Ln* (*CR*), *Ln* (*R*), *Ln* (*Q*), *Ln* (*FC*), *Ln* (*q*)
Test Assumption: A Deterministic Trend in the Data
Lag Intervals: 1 a 3

$$LR_{trace} = -T\sum_{i=r+t}^{k}\log(1-\lambda_i)$$

| $H_0$ | $H_1$ | Eigenvalue | $LR_{\text{trace}}$ | 5% Critical Value | 1% Critical Value |
|---|---|---|---|---|---|
| R = 0 | R = 1 | 0.1545 | 79.0465 | 77.74 | 85.78 |
| R ≤ 1 | R = 2 | 0.1340 | 49.5080 | 54.64 | 61.24 |
| R ≤ 2 | R = 3 | 0.0788 | 24.1717 | 34.55 | 40.49 |
| R ≤ 3 | R = 4 | 0.0324 | 9.7249 | 18.17 | 23.46 |
| R ≤ 4 | R = 5 | 0.0220 | 3.9280 | 3.74 | 6.40 |

* The critical values for the cointegration test are taken from Osterwald-Lenun (1992)

In order to interpret the estimated cointegrating vector, it is a common practice to normalize it on one of the variables by setting its estimated coefficient equals to -1. Since one of our interests is to obtain long-run elasticities of the credit demand function we normalize the coefficient of the variable *Ln*(*CR*) and divided the whole cointegrating vector by –1. The results are reported in Table 6. The equation shows signs on the variables that are consistent with theory. The long-run real loans demand are positively related to the level of economic activity, the real exchange rate and the mark-up, and negatively related to interest rates on loans (with a long-run elasticity of –0.47).

Table 6
Cointegrating Vector: Ln(*CR*)
Normalized Cointegrating Coefficients

Sample: 1986:01-2000:12
Included observations: 176
Lag Intervals: 1 a 3

| Ln(*CR* ) | Ln(*Q*) | Ln(*R* ) | Ln(*CF*) | Ln(*q* ) | Trend | C |
|---|---|---|---|---|---|---|
| -1.000 | 0.5654 | - 0.4756 | 2.9140 | 0.0278 | - 0.0088 | 13.2472 |
| Log likelihood | 1442.75 | | | | | |

In the context of our theoretical specification a positive relation between real loans and the mark-up may indicate that the cash flow impact on private investment, which indirectly affects the financing gap, is greater that the direct impact on (*I - CF*). [15]

The positive relationship between the real exchange rate and real credit demand turns out to be a very significant element of analysis for an economy that has enthusiastically adopted several prolonged experiments, during recent years, in the use of the exchange rate as a nominal anchor. Again in terms of the transmission mechanisms of the theoretical model, the empirical results here would be indicating that a real exchange rate appreciation would discourage investment and credit demand (in favor of investment in external financial assets) to a greater extent than the financial leverage arising from the search for funds to be invested in the rest of the world. However, we need to point out that the elasticity value is quite low.

## 6. The Short-run Dynamics of Credit Demand

The Granger representation theorem states that if a cointegrating relationship exists between a set of I(1) variables, then a dynamic error correction representation of the data also exists. The estimation of an error correction model allows us to capture the short-run dynamics that characterized the process of adjustment towards equilibrium.

Our general error correction relationship will have the following form:

$$\begin{aligned}\Delta Ln(CR)_t = C + \sum_{i=0}^{n} \alpha_i \Delta Ln(Q)_{t-i} + \sum_{i=0}^{n} \beta_i \Delta Ln(R)_{t-i} + \\ \sum_{i=0}^{n} \delta_i \Delta Ln(CF)_{t-i} + \sum_{i=0}^{n} \varepsilon_i \Delta Ln(q)_{t-i} + \phi Z_{\tau-1} + e_t\end{aligned} \tag{10}$$

where $Z_{\tau\text{-}1}$ represents the error correction term, and, in our case, it is based on the cointegration regression using Johansen's approach.

Since we are dealing here with the estimation of a demand function and given the features of the model adopted, it will be interesting to verify how sensitive is the specification to the problems of endogeneity. In particular, it is possible that the rate of interest on loans or the index of real economic activity can be deemed endogenous. It is in this sense that may be useful to find out the properties of a short-run estimation in the framework of a vector error correction model (VECM). Despite its general lack of structure, a VECM is intuitively attractive because it not only explicitly recognizes a wide range of short-run dynamics as well as the long-run equilibrium relationships among variables in the system, but also allows us to capture potential endogeneity of the determinants of credit demand. A VECM is a restricted VAR that has cointegration restrictions built into the specification, so that it is designed for use with nonstationary series that are known cointegrated. It is for that reason that

[15] A word of caution is needed here because we are not strictly estimating the determinants of a private investment function. In addition, the mark-up variable is still a rough approximation of the true cash flow variable.

cointegration tests are indeed useful as a previous step in the specification, diagnostic and analysis of a credit demand function.[16]

Table 7 shows the estimated VECM over a sample period that goes from may 1986 to December 2000 as well as some conventional diagnostic tests. The first column vector, which corresponds to the loans equation, shows that the coefficient of the error correction term is significantly different from zero. This would confirm the existence of a long-run relationship between real loans and the set of variables involved. The negative sign of the error correction coefficient is also important because it indicates that the real loans model, as it is presented and without being subject to a parsimonious process of selection, converge towards equilibrium. The magnitude of the coefficient, in terms of the monthly data set, is rather small suggesting that in case of deviation of the real stock of loans from their equilibrium level, this should be corrected only slowly.

Weak exogeneity tests are performed on the equations for index of economic activity and the interest rate, *Ln*(*R*) in order to determine whether, in the spirit of the general-to specific methodology, it would be legitimate to specify the demand for loans as a single-equation model instead of a system. The test is performed by assessing the statistical significance of the coefficient of the error correction term in each of the equations for *Ln*(*Q*) and *Ln*(*R*). Indeed, the respective *t* statistics 0.17 and -0.87 imply that there is no information loss from excluding those equations from the system. Therefore, both real sales and the interest rate on loans can be considered as weakly exogenous variables.[17]

A single-equation error correction model can be derived from a simple reparameterization of the column vector for *Dln*(*CR*). The general-to-specific reductions of the restricted and highly parameterized vector for *Dln*(*CR*) are designed to ensure that the parsimonious subset of the vector will convey all the information embodied in the original vector. In addition, this method will allows us to include, where necessary, contemporary values of the explanatory variables. This would not be possible in a pure VECM specification.

Several parsimonious procedures exist for model selection in a VECM. Krolzig (2000) suggests selection and diagnostic testing process: starting from the highly parameterized congruent general model, standard testing procedures are used to eliminate statistically-insignificant variables with diagnostic test checking the validity of the reductions and ensuring a congruent final selection.

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[16] In this case the representation is a VECM of order $p$ and it would be given by $Dy_t = A_1 Dy_1 + ... + A_p Dy_{t-p} + Bx_t + CZ_{t-1} + e_t$, where $y_t$ = {ln(*CR*), Ln(*R*), Ln(*Q*), Ln(*CF*), Ln(*q*)} es el is the vector of variables I(1), $x_t$ is the vector *d* of deterministic variables, $Z_{t-1}$ is the vector error correction, A, B and C are coefficient matrices, and $e_t$ is the vector of innovations.

[17] Granger causality test also confirm that the interest rate can be considered as an exogenous variable, which lends support to the idea sustained in the theoretical model that banks fix the interest rate in a context of imperfect competition.

Table 7

Estimation of the Vector Error Correction

Sample: 1986:05-2000:12

*t* statistics in parenthesis

| | D(Ln CR) | D(Ln Q) | D(Ln FC) | D(Ln R) | D(Ln q) |
|---|---|---|---|---|---|
| **Error Correction Term** | -0.034946 | 0.008497 | 0.024522 | -0.029011 | -0.015912 |
| | **(-3.02506)** | **(0.17092)** | **(-3.69936)** | **(-0.84853)** | **(-0.66610)** |
| **D(LnCR(-1))** | 0.324068 | -0.516937 | 0.061079 | 0.770507 | 0.126570 |
| | **(-3.55877)** | **(-1.31922)** | **(-1.16893)** | **(-2.85897)** | **(0.67216)** |
| **D(LnCR(-2))** | 0.045785 | -0.223796 | -0.003135 | -0.399708 | -0.251942 |
| | **(0.44987)** | **(-0.51101)** | **(-0.05369)** | **(-1.32701)** | **(-1.19714)** |
| **D(LnCR(-3))** | 0.040463 | 0.629384 | 0.053913 | -0.207464 | 0.171947 |
| | **(0.45429)** | **(-1.64213)** | **(-1.05488)** | **(-0.78703)** | **(0.93358)** |
| **D(LnQ(-1))** | -0.007529 | -0.229376 | 0.020288 | -0.162492 | -0.009317 |
| | **(-0.31515)** | **(-2.23111)** | **(-1.47985)** | **(-2.29804)** | **(-0.18858)** |
| **D(LnQ(-2))** | -0.044400 | -0.270248 | 0.016799 | 0.015735 | 0.006339 |
| | **(-1.90682)** | **(-2.69717)** | **(-1.25730)** | **(0.22833)** | **(0.13166)** |
| **D(LnQ(-3))** | 0.008071 | -0.030838 | 1.26E-05 | -0.025919 | -0.042514 |
| | **(0.40020)** | **(-0.35533)** | **(0.00109)** | **(-0.43424)** | **(-1.01940)** |
| **D(LnFC(-1))** | 0.446650 | 0.591758 | 0.290172 | -1.165 | 0.319578 |
| | **(-3.13595)** | **(0.96552)** | **(-3.55048)** | **(-2.76493)** | **(-1.08507)** |
| **D(LnFC(-2))** | 0.208239 | 0.136339 | -0.058745 | -0.863730 | 0.623434 |
| | **(-1.35055)** | **(0.20549)** | **(-0.66398)** | **(-1.89276)** | **(-1.95533)** |
| **D(LnFC(-3))** | -0.102229 | 0.105414 | 0.064807 | 0.651788 | -0.977203 |
| | **(-0.81269)** | **(0.19474)** | **(0.89785)** | **(-1.750)** | **(-3.75677)** |
| **D(LnR(-1))** | -0.055158 | -0.193210 | -0.016016 | 0.326941 | 0.134694 |
| | **(-2.02587)** | **(-1.64911)** | **(-1.02516)** | **(-4.057)** | **(-2.39240)** |
| **D(LnR(-2))** | 0.005993 | -0.103266 | -0.005506 | 0.031676 | -0.080381 |
| | **(0.20744)** | **(-0.83072)** | **(-0.33213)** | **(0.37050)** | **(-1.34560)** |
| **D(LnR(-3))** | 0.036766 | 0.109948 | -0.006642 | -0.072108 | 0.007020 |
| | **(-1.45486)** | **(-1.01106)** | **(-0.45805)** | **(-0.96410)** | **(0.13434)** |
| **D(Lnq(-1))** | -0.093097 | -0.189286 | -0.164803 | 0.214547 | -0.133180 |
| | **(-2.54522)** | **(-1.20260)** | **(-7.85208)** | **(-1.981)** | **(-1.76079)** |
| **D(Lnq(-2))** | 0.102636 | 0.017347 | 0.029227 | -0.422801 | -0.090234 |
| | **(-2.28729)** | **(0.08984)** | **(-1.13509)** | **(-3.18363)** | **(-0.97245)** |
| **D(Lnq(-3))** | 0.005761 | -0.054315 | -0.006181 | -0.311819 | 0.072424 |
| | **(0.12691)** | **(-0.27806)** | **(-0.23732)** | **(-2.32100)** | **(0.77155)** |
| **C** | -0.003420 | -0.006980 | 0.001606 | 0.003389 | -0.004853 |
| | **(-1.43693)** | **(-0.68150)** | **(-1.17592)** | **(0.48116)** | **(-0.98617)** |
| R-squared | 0.418280 | 0.197540 | 0.396925 | 0.282122 | 0.204019 |
| Adj. R-squared | 0.359742 | 0.116790 | 0.336239 | 0.209883 | 0.123920 |
| Sum sq. resids | 0.143860 | 2.663 | 0.047367 | 1.260 | 0.615147 |
| Standard Error | 0.030080 | 0.129 | 0.017260 | 0.089023 | 0.062200 |
| F statistics | 7.145 | 2.446 | 6.540 | 3.905 | 2.547 |
| Log likelihood | 375.8 | 119.049 | 473.654 | 184.925 | 248.02 |
| Akaike AIC | -4.078 | -1.159 | -5.189 | -1.908 | -2.625 |
| Schwarz SC | -3.772 | -0.853416 | -4.883 | -1.602 | -2.319 |
| Mean dependent | -0.003761 | -0.002623 | 0.002094 | 0.003479 | -0.004030 |
| S.D. dependent | 0.037592 | 0.137728 | 0.021185 | 0.100151 | 0.066454 |

| | |
|---|---|
| Determinant Residual Covariance | 5.29E-14 |
| Log Likelihood | 1441.58 |
| Akaike Information Criteria | -1.534.750 |
| Schwarz Criteria | -1.370.821 |

From the vector *Dln*(*CR*) we undertake pre-search simplification *t* test and exclude lag variables with reported insignificant values, after which the resulting single equation model is reformulated. Model A in Table 8 indicates that the first difference of the credit variable, the interest rate, the mark-up and the real exchange rate will enter into the model with one lag, the real sale index is retained with two lags and the real exchange rate variable with one and two lags respectively. The coefficient of the error correction term remains negative and is significantly different from zero. However, the sign of the real sales index is not consistent with the theory. Alternatively, we can either exclude the variable or use a different specification with one lag and/or the contemporary value. Model B, shows a version where the real sale index is excluded from the equation. The F test indicates that the joint significance of the variables improves (with respect to A), the coefficient of the error correction term keeps the right properties, but the information criteria do not improve. Version C, where the contemporary value of the real sales index is included, shows a positive sign coefficient for that variable, right properties for the coefficient of the error correction term, and better performance of the F test.[18] The information criteria of Akaike and Schwarz indicate that model C represents a better specification than A and B.

Table 8

Dependent Variable: DLn(CR)
Sample: 1986:05 2000:12
Observations: 176
Method: OLS

| Regresor | Model A | *t* Value | Model B | *t* Value | Model C | *t* Value | Model D | *t* Value |
|---|---|---|---|---|---|---|---|---|
| DLn(CR)(-1) | 0,6575 | 6,23 | 0,59 | 5,74 | 0,6356 | 6,87 | 0,6735 | 8,19 |
| DLn(Q) | | | | | 0,1014 | 6,43 | 0,1004 | 7,18 |
| DLn(Q)(-2) | -0,0416 | -2,31 | | | | | | |
| DLn(CF)(-1) | 0,5016 | 3,89 | 0,4744 | 3,65 | 0,4814 | 4,12 | 0,2723 | 2,52 |
| DLn(R)(-1) | -0,0571 | -2,36 | -0,0523 | -2,14 | -0,0331 | -1,49 | -0,0114 | -0,57 |
| DLn(q)(-1) | -0,0817 | -2,33 | -0,0914 | -2,59 | -0,0631 | -1,97 | -0,0014 | -0,05 |
| DLn(q)(-2) | 0,1342 | 3,22 | 0,1184 | 2,84 | 0,1201 | 3,21 | 0,0813 | 2,42 |
| DUM | | | | | | | -0,1329 | -6,84 |
| Z(-1) | -0,3323 | -2,52 | -0,2622 | -2,02 | -0,2004 | -1,71 | -0,2909 | -2,78 |
| C | -0,002 | -0,88 | -0,0022 | -0,93 | -0,0017 | -0,81 | 0,0003 | 0,18 |
| R-squared | 0,3751 | | 0,3551 | | 0,4825 | | 0,5959 | |
| R-squared Adjusted | 0,349 | | 0,3322 | | 0,4609 | | 0,5766 | |
| D-W Statistic | 1,9467 | | 1,9864 | | 2,0649 | | 2,0407 | |
| Akaike Info Criteria | -4,109 | | -4,0889 | | -4,2976 | | -4,5338 | |
| Schwarz Criteria | -3,9649 | | -3,9628 | | -4,1535 | | -4,3716 | |
| F-Statistic | 14,4008 | | 15,51 | | 22,3802 | | 30,795 | |
| Prob (F-Statistic) | 0 | | 0 | | 0 | | 0 | |

Figure 4 shows the actual and fitted series, and the residuals of model C. The residuals look fairly satisfactory, but there are three strong deviations at 1989:3, 1996:5 and 1996:10. At least two of these spikes could be representing strong exchange rate adjustments occurred in 1989 and 1996 after severe balance of payments crisis. Inserting a dummy variable that takes on the value of 1 in 1989:3 and 1996:5 and 0

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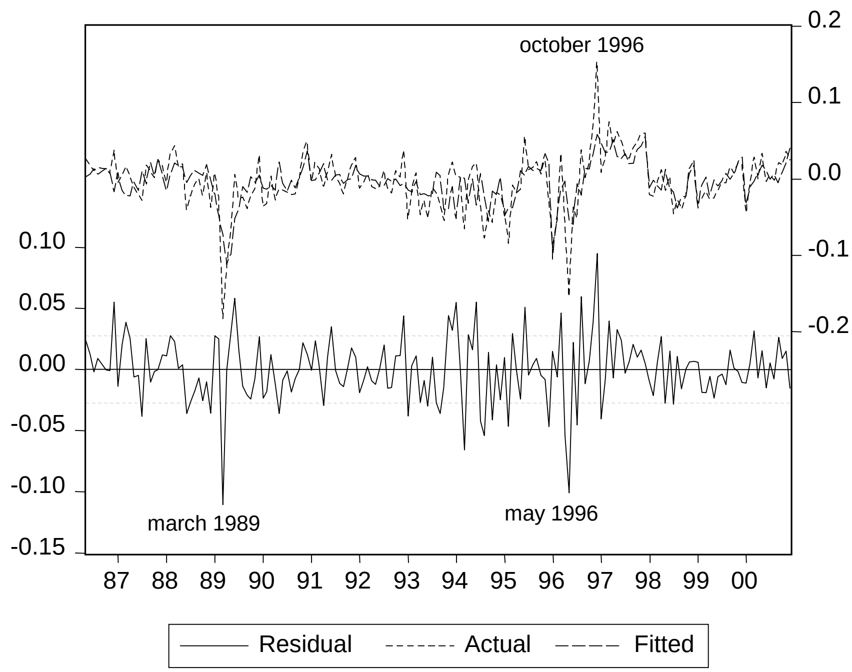
18 This is hardly surprising since we have already shown the very high contemporaneous correlation between L*n*(*CR*) and L*n*(*Q*).

elsewhere gives the regression output shown in Table 8 for model D.[19] The coefficient of the dummy variable is highly significant, the estimated coefficient of the error correction term keeps the expected sign and appears to be statistically significant, and the joint significance of the group (F-test) shows obviously an improvement on those from the previous single-equation models. However, lagged values of one month for the interest rate and the real exchange rate yield coefficients that are statistically insignificant. There is much to be gained by deleting these one-month lags from the equation since both the Akaike and Schwarz criterion move in the right direction and the F-test shows further improvement. Model E in Table 9 shows this reduction of the equation.

Figure 4
The actual and fitted value of DLn(CR) in Model C




It is interesting to note that in model E the impact of the variable interest rate has been absolutely removed (due to lack of statistically significance). However, one still may wonder whether the joint or individual inclusion of current values for the variables interest rate and the real exchange can be critical or not. We have chosen to recast in terms of the contemporary values of these variables and reestimate. The best specification turned out to be the one given by model F (in the same Table 9). The reader will note that the only difference between model E and F is the inclusion of the current value of the interest rate on loans. Evaluation of models E and F does not yield clear conclusions for final selection. In terms of AIC and SC the selection criteria is

[19] It is not clear what sort of event can capture the spike in 1996:10. Moreover, the deviation goes in opposite way to the spikes presented in 1989:3 and 1996:5. A separate dummy variable for this event did not yield a significant value for the respective coefficient.

ambiguous and even though the F-test clearly supports the selection of E, the results indicate that there is little to be gained by excluding *Ln*(R).

Table 9
Diagnostic Testing in the Error Correction Models E and F

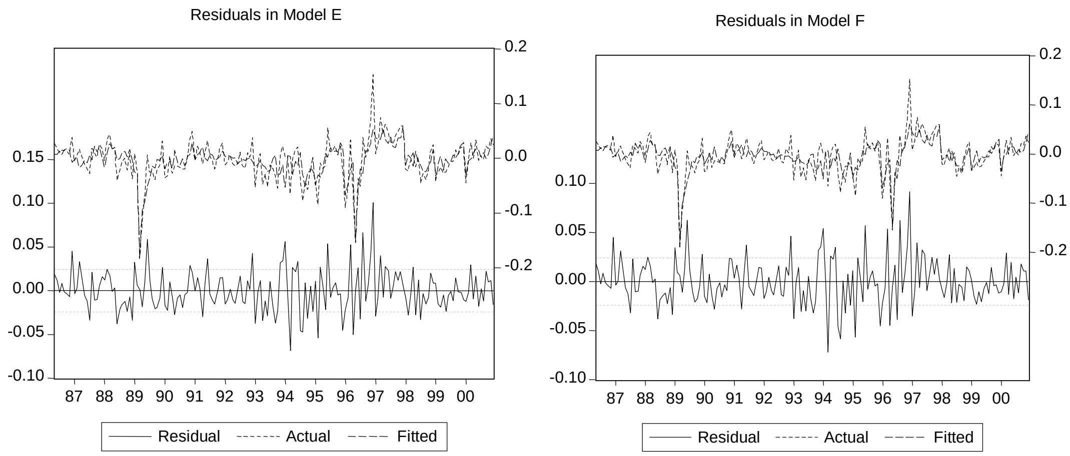
Dependent Varaible: DLn(*CR)*
Sample: 1986:05 2000:12
OLS and 176 Observations

| **Regresor** | **Model E** | | **Model F** | |
|---|---|---|---|---|
| | Coefficient | *t* Value | Coefficient | *t* Value |
| DLn(CR)(-1) | 0,6814 | 8,45 | 0,6855 | 8,55 |
| DLn(Q) | 0,1015 | 7,42 | 0,1017 | 7,49 |
| DLn(CF)(-1) | 0,2686 | 2,52 | 0,2324 | 2,15 |
| DLn(R) | | | -0,0389 | -1,83 |
| DLn(q)(-2) | 0,0784 | 2,37 | 0,0673 | 2,02 |
| DUM | -0,1349 | -7,42 | -0,1192 | -5,96 |
| Z(-1) | -0,291 | -2,8 | -0,2745 | -2,64 |
| C | 0,0003 | 0,19 | 0,0003 | 0,19 |
| R-squared | 0,5951 | | 0,603 | |
| R-squared Adjusted | 0,5808 | | 0,5865 | |
| D-W Statistic | 2,0484 | | 2,0519 | |
| Akaike info criteria | -4,5545 | | -4,5629 | |
| Schwarz criteria | -4,4284 | | -4,4187 | |
| F-Statistic | 41,4115 | | 36,4675 | |
| Prob (F-Statistic) | 0 | | 0 | |
| | F-Test | Probability | F-Test | Probability |
| Breusch-Godfrey LM Test (order 1)[1] | 0,3067 | 0,5804 | 0,3611 | 0,5486 |
| Breusch-Godfrey LM Test (order 2)[1] | 0,4897 | 0,6136 | 0,4761 | 0,6219 |
| Breusch-Godfrey LM Test (order 3)[1] | 0,3245 | 0,8076 | 0,3265 | 0,8061 |
| ARCH(1) Test[2] | 1,6508 | 0,2 | 1,6508 | 0,2 |
| ARCH(2) Test[2] | 0,7696 | 0,4647 | 0,8135 | 0,445 |
| ARCH(3) Test[2] | 1,2834 | 0,2816 | 1,1294 | 0,3387 |
| Ramsey Reset Test (second power)[3] | 3,2026 | 0,0753 | 3,3771 | 0,0678 |
| Ramsey Reset Test (2nd and 3rd power)[3] | 3,298 | 0,0393 | 3,0804 | 0,0485 |

1/ Ho: No serial correlation in the residulas
2/ Ho: No hetedocedasticity of orden q
3/ Ho: No epecification error

Both short-run models have survived the same battery of tests. Visual examination of the residuals (in figure 5) shows that the 1989:3 and 1996:5 spikes has been drastically reduced and it is particularly remarkable the ability of both models to track very accurately the cyclical movements of the D*Ln*(CR) series. Additional diagnostic testing is presented in Table 10. The Breusch-Godfrey asymptotic test for serial correlation up to the third order gives a *P* value of 0.80, so the hypothesis of autocorrelation is rejected. Test of ARCH residuals with one up to three lags give *P* values between 0.2 and 0.46, so the assumption of homocedastic residuals is not rejected in favor of ARCH residuals. The Ramsey RESET test for specification error with one fitted term to include in the regression cannot reject the null of no significant evidence of misspecification (at 5% significance level).

Figure 5




Based on the behavior of D*Ln*(*CR*) and the evolution of the events over the period of study, we conduct (for each model E and F) a Chow forecast test that examine whether the parameters of the model are stable across various subsamples of the data.[20] We presume that a structural break might have taken place in 1996:5 or alternatively in 1996:10. Neither the F-statistic nor the log likelihood ratio reject the null of no structural change in the short-run credit demand function before and after 1996:5, and before and after 1996:10.

Table 10
Chow's Forecast Test

| | **Model E** | **Model F** |
|---|---|---|
| 1996:05 a 2000:12[1] | | |
| F-Statistic | 1,0384 (0,42) | 0,976 (0,53) |
| Log LR | 73,067 (0,06) | 69,95 (0,09) |
| 1996:10 a 2000:12[1] | | |
| F-Statistic | 0,8992 (0,65) | 0,8297 (0,77) |
| Log LR | 57,789 (0,23) | 54,336 (0,34) |

[1]/ Probability in parenthesis

Inspection of the short-run dynamics in the final specifications E and F indicates that the lagged value of the dependent variable D*Ln*(*CR*), appears strongly significant with the highest coefficient.[21] This reflects what in literature is known as the "loan-customer relationship" (LCR). The LCR factor operates in two ways: One the one hand,

[20] Once the data is split, the Chow forecast test estimate both models (E and F) for a subsample comprised of the first set of observations. The estimated model is then used to predict the values of the dependent variable in the remaining data points.
[21] See Wood 1974, Hicks 1980, Catao 1997, and Panagopoulus y Spilotis 1998 for similar findings.

clients keep a continues relationship with their banks, taking loans even above the short-run optimal requirements, as a precautionary measure against cyclical an unpredictable interruptions in the credit market; on the other hand, banks are interested in keeping such a long-run relationship with their clients and extends loans beyond the amount consistent with their short-run profit maximization program as a way to avoid customer mobility towards other competitors. For banks this ongoing relationship is important because it allow them to develop informational advantages and minimize elements of moral hazard.

The coefficient on $DLn(Q)_{t-1}$ (lagged real sales index) also features a significant asymptotic *t*-ratio of the theoretically anticipated sign (positive). Nevertheless, in both models the elasticity is lower than the one reported in the long-run relationship.

As for the contemporary value of interest rate variable, it plays no role in the E model and it takes the correct sign but it is still not significant even at the 5% level in the F model. The coefficient is rather small (it is only -0,03) and it seems reasonable to acknowledge that it is little use in explaining short-run dynamics.

The coefficient on $DLn(CF)_{t-1}$ is positive, highly significant, and the value of elasticity reported in both models is quite high. This may suggest that higher mark-ups and greater cash flow increases investment and in turn credit demand.

For a financially open economy as small as Venezuela, it is very interesting to find (even in the short-run), a mild impact of the real exchange rate on credit demand. With two lags the real exchange rate coefficient yields a positive sign and it is significant at the 5% level. Therefore, there arises the possibility that corporate firms substitute investment in fixed assets for foreign financial assets as soon as the real exchange rate appreciates.

The coefficient of the error correction term has a negative sign and is significantly different from zero. It ranges between -0.27 and -0.29, which indicates that less than one third of the disequilibrium at a given period in the credit market is corrected in one month.

## Conclusions

This study examines the recent marked slowdown in bank credit to the private sector in Venezuela and explores the determinants of the demand for loans and their possible role as the dominant side of the market. The estimation presented here has followed a minimum theoretical specification in an attempt to make up for the lack of theoretical rigor in much of the empirical work in this field. The theoretical model has post-Keynesian roots and assumes that banks operate in a context of credit market imperfections, setting the price of credit and responding to demand. Demand is driven by firms financial decisions which in turn are based on changes in balance sheets.

Cointegration tests using Engle and Granger and Johansen procedures coincide in proving the existence of a cointegration vector between the real stock of credit, the real sales index, the nominal interest rate, the mark-up, and the real exchange rate (all expressed in logarithms). The signs of long-term elasticities coincide with the theoretical assumptions. Real credit responds positively to increases in the rhythm of economic activity, to changes in the mark-up and to variations in the real exchange rate. The coefficient of the lending rate showed a negative sign.

The estimation of the vector error correction model has corroborated the weak exogeneity of both the interest rate and the real sales index. These results support the particular manner in which the theoretical model serves to identify credit demand: banks set interest rates and demand determines the volume of credit. We have restricted the error correction vector, thereby obtaining two final dynamic models to explain the short-term movements of real credit. The difference between the two models is the inclusion, or not, of the contemporary value of the interest rate, which in the best case has a significance only at 10 percent. Both error correction models confirm the importance, in short-term dynamics, of the lagged value of real credit. This may be interpreted as reflecting the need of clients to mitigate cyclical changes in credit, and the need of banks to maintain ongoing relations with their clients. Just as in the long-term equilibrium relationship, the contemporary value of the level of economic activity and the mark-up positively affect real credit, although their short-term elasticity are rather less than that reported for the long-run. Moreover, the effect of the interest rate on loans is virtually nil in the short term (elasticity value: -0.03). Both error correction models likewise corroborate the impact of the real exchange rate on the restructuring of corporate assets. The recurrent and prolonged periods of exchange appreciation in Venezuela seem to discourage credit demand even in the short term, in a process that may well be operating by means of fewer requirements for investment, in favor of increased positions in external financial assets, with expected returns that rise with exchange rate appreciation.

The two specifications were subject to intensive diagnostic testing with satisfactory results. The models are also stable according to Chow's forecast test, but the Ramsey's RESET test is inconclusive, which could indicate that some variables may have been omitted. Such results should not be surprising considering the simplifications of the theoretical specification and the empirical difficulties of capturing the effect of risk on credit demand in a single variable.

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