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This paper can be downloaded without charge from http://www.ecb. int or from the Social Science Research Network electronic library at: http://papers.ssrn.com/abstract_id=xxxxxx
2 *Directorate Monetary Policy, European Central Bank, Postfach 16 03 19, D-60666 Frankfurt am Main, Germany. E-mail: alessandro.calza@ecb.int, joao.sousa@ecb.int*
3 *Estudios Monetarios y Financieros, Servicio de Estudios, Banco de España, Calle Alcalá 50, 28014 Madrid, Spain. E-mail: mmanrique@bde.es*

Address	**Kaiserstrasse 29**
	D-60311 Frankfurt am Main
	Germany
Postal address	**Postfach 16 03 19**
	D-60066 Frankfurt am Main
	Germany
Telephone	**+49 69 1344 0**
Internet	**http://www.ecb.int**
Fax	**+49 69 1344 6000**
Telex	**411 144 ecb d**

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Contents

Abstract

This paper provides new evidence on the behaviour of euro area aggregate loans to the private sector. Using a sample covering the last twenty years, a cointegrating vector linking the real stock of loans to a small set of domestic macroeconomic variables is found. Besides real GDP and prices, this set includes a new measure of the cost of loans obtained as a weighted average of bank lending rates. The results are overall encouraging, though the recursive estimates of the long-run parameters suggest that in 2000 some disturbances, probably of a temporary nature, affected the system. The study then addresses the issue of the leading indicator properties of loans. It finds that the deviations of the real stock of loans from the equilibrium level implied by the model seem to contain information on future changes in inflation, though not on its level.

JEL classification: C32, C51.
Keywords: credit, loans, euro area, cointegration.

Non-Technical Summary

The analysis of monetary and credit aggregates plays a key role in the monetary policy framework of the European Central Bank (ECB). Loans to the private sector granted by Monetary Financial Institutions (MFI) are the largest counterpart to M3 and the most prominent among the credit aggregates monitored by the ECB. Recent studies have provided evidence that euro area loans to the private sector exhibit some desirable empirical properties. For instance, using cointegration techniques, Hofmann (2001) finds evidence of stable long-run relationships between the stock of loans to the private sector and domestic macroeconomic variables in several individual euro area countries, while Calza, Gartner and Sousa (2001) provide similar evidence for the area as a whole. In addition, based on out-of-sample forecasting techniques developed by Stock and Watson (1999), Nicoletti-Altimari (2001) shows that the developments in nominal loans contain substantial information about medium term prospects for area-wide inflation. These findings suggest that the monitoring of developments in loans to the private sector in a macroeconomic framework may yield useful information about the state of the economy and the emergence of risks for future price stability.

The purpose of this work is to empirically investigate the behaviour of euro area aggregate loans to the private sector and to analyse their leading indicator properties for inflation. Consistent with the approach followed in Calza, Gartner and Sousa (2001), the empirical analysis addresses the identification issue by means of modelling techniques conventionally used in money demand studies. A long-run relationship linking the stock of real loans to a small set of macroeconomic variables is estimated using multivariate cointegration techniques. Besides real GDP and prices, this set includes a new measure of the cost of loans, obtained as a weighted average of bank lending rates. The estimated vector error-correction system performs satisfactorily in terms of the statistical properties of its residuals. Nevertheless, the recursive estimates of the long-term parameters reveal that in 2000 some disturbances, probably of a temporary nature, affected the system. The main candidates to explain such disturbances are intense mergers and acquisitions activity (both domestically and abroad), rising real estate prices and increased need by telecom companies to finance the UMTS license auctions.

After estimating the loan error-correction term of the model, the study empirically investigates its information content on future inflation. In order to do so, the study applies a methodology commonly applied to the assessment of the usefulness of the yield spread to forecast future activity (see Hamilton and Kim (2002)). While the loan overhang/shortfall does not seem to have significant information content as regards the future level of inflation, our empirical results suggest that loan disequilibria should help to predict future changes in inflation.

1. Introduction

The analysis of monetary and credit aggregates plays a key role in the monetary policy framework of the European Central Bank (ECB).[1] Loans to the private sector granted by Monetary Financial Institutions (MFI) are the largest counterpart to M3 and the most prominent among the credit aggregates monitored by the ECB.[2] Recent studies have provided evidence that euro area loans to the private sector exhibit some desirable empirical properties. For instance, using cointegration techniques, Hofmann (2001) finds evidence of stable long-run relationships between the stock of loans to the private sector and domestic macroeconomic variables in several individual euro area countries, while Calza, Gartner and Sousa (2001) provide similar evidence for the area as a whole. In addition, based on out-of-sample forecasting techniques developed by Stock and Watson (1999), Nicoletti-Altimari (2001) shows that the developments in nominal loans contain substantial information about medium term prospects for area-wide inflation. These findings suggest that the monitoring of developments in loans to the private sector in a macroeconomic framework may yield useful information about the state of the economy and the emergence of risks for future price stability.

Despite the relatively large dependence of the euro area on its banking sector, only few studies have investigated the determinants of aggregate loans to the private sector in euro area countries.[3] Recent studies include De Bandt and Jacquinot (1992) and Odonnat, Grunspan and Verdelhan (1997) on credit demand from corporations in France, Focarelli and Rossi (1998) on flow demand for credit to firms in Italy (also analysing behavioural differences across geographical regions), Panagopoulos and Spiliotis (1998) on loans to firms in Greece, Vega (1989) on aggregated credit demand in Spain and Manrique and Sáez (1998) on the demand for credit to households and firms in the same country. As mentioned above, Hofmann (2001) estimates long-run relationships linking real loans to the private sector, real GDP, short-term market interest rates and property prices in a number of industrialised countries, including Belgium, Germany, Spain, France, Ireland, Italy, Netherlands and Finland in the euro area. Calza, Gartner and Sousa (2001) models demand for euro area real loans to the private sector using real GDP and real market interest rates as explanatory variables. As part of systems also including supply equations, Kakes (2000) and Hülsewig et al. (2001) model the behaviour of bank lending demand in the Netherlands and Germany, respectively.

[1] The terms "credit" and "loans to the private sector" are not used interchangeably in this paper since, according to the ECB definition, the former is a broader category including both loans and the MFI holdings of shares and other securities. In October 2001 loans accounted for 89% of the outstanding amounts of MFI credit.

[2] See Masuch et al. (2000) for a description of the role of the counterparts of M3 (notably, loans to the private sector) in the ECB's regular assessment of risks for future price stability.

[3] The behaviour of bank lending is the subject of a fast growing strand of empirical literature inspired by the new "credit view", which emphasises the importance of imperfections in the financial markets. Empirical contributions to this literature are typically focused on testing for the existence of a credit channel of monetary policy transmission. Due to identification problems, these studies tend, though, to be based on micro-level - rather than aggregate - data.

Equations on loan demand in individual euro area countries can also be found in the monetary and financial blocks of some national and multi-country macro-econometric models. For instance, Jeanfils (2000) reports equations for mortgage loans and consumer credit in the Banque National de Belgique's quarterly model of the domestic economy. De Bandt et al. (2002) model separately loans to households and non-financial corporations within the Banque de France quarterly model of the French economy. Banca d'Italia's (1986, p. 168) model of the Italian economy includes one equation for total credit demanded by firms and households. The Dutch central bank's model of the domestic economy comprises individual equations for both short-term and long-term credit demand from households and firms (see Fase et al. (1992) and the updated versions by Van Els and Vlaar (1996)). Similarly, as part of the monetary block of its Euromon model, De Nederlandsche Bank (2000, pp. 83-85) provides error-correction models for the demand for bank loans to the private sector in Belgium, Germany, Spain, France, Italy, the Netherlands, Austria and Finland.

The relatively limited availability of studies on aggregate loans can be explained by both the traditional predominance of the "money view" (with its focus on the exclusive role of money in macroeconomic analysis) and the importance of identification problems. The identification issue regards the general difficulty of distinguishing between demand and supply factors when estimating behavioural relationships for monetary and credit aggregates. Authors adhering to the "credit view" argue that, due to informational asymmetries between lenders and borrowers, supply effects should play a particularly significant role in the market for loans via the so-called "balance sheet" transmission channel and/or the "bank lending" channel (see for instance Bernanke (1993)).

In general terms, the magnitude of loan supply effects should depend to a large extent on the role of banks in the financial structure of the economy and the way they react to changes in monetary policy stance or economic conditions (as well as the overall credit and liquidity conditions in the economy). Ehrmann et al. (2001) provide evidence on the relevance of loan supply effects in the euro area. Using micro bank data, the authors find that, following rises in policy rates, commercial banks generally reduce loan supply, with the size of the reduction inversely depending on their degree of liquidity. However, other bank-specific characteristics which have been found to amplify loan supply effects in the USA (e.g. size and capitalisation) do not seem to play a major role in the euro area. Angeloni et al. (2002) assess the existing evidence and conclude that institutional factors such as government guarantees, banking networks and the prevalence of relationship lending probably reduce the strength of the bank lending channel in the euro area.

The purpose of this work is to empirically investigate the behaviour of euro area aggregate loans to the private sector. Consistent with the approach followed in Calza, Gartner and Sousa (2001), the empirical analysis addresses the identification issue by means of modelling techniques conventionally used in money demand studies. A long-run relationship linking the stock of real loans to a small set of macroeconomic variables is estimated using multivariate cointegration techniques. A major novelty of the analysis consists of the use of a newly constructed measured of the cost of loans obtained as a weighted average of bank lending rates. In addition, the study addresses the issue of the leading indicator properties of loans for inflation. In particular, it investigates the information content on future inflation of the loan error-correction term using a methodology commonly applied to the study of the usefulness of the yield spread to forecast future activity (see Hamilton and Kim (2002)).

As a caveat to the analysis, it should be noted that the modelling techniques used are based on the assumption of limited supply effects.[4] Therefore, given the uncertainty regarding the importance of supply factors related to the balance sheet channel, it cannot be ruled out that the estimated relationships may to some extent capture also supply factors. Another important caveat regards the relatively high degree of data aggregation of the analysis. Due to the lack of sufficiently long harmonised historical data, the present work focuses on total loans to the private sector, rather than their sectoral breakdown, which may potentially generate "aggregation bias" in case of significant heterogeneity between the credit demand behaviour of firms and households.

The remainder of this paper is organised as follows: Section 2 describes stylised facts about loans to the private sector and other macroeconomic variables in the euro area. Section 3 focuses on the modelled long-run relationship and the main theoretical considerations underlying it. Section 4 is dedicated to the empirical analysis and, in particular, to the estimation and evaluation of the model. Section 5 investigates the information content on future inflation developments of the loan error-correction term. The main conclusions and suggestions for further research are in Section 6.

2. Stylised facts about loans to the private sector in the euro area

As Figure 1 shows, over the last twenty years the developments of loans to the private sector (deflated by the GDP deflator) have followed relatively well defined cycles, in turn positively correlated with the general business cycle.[5] In particular, whenever real GDP growth exceeded its average rate over this period (2.2%), real lending growth also tended to rise above its own average (4.0%). Nevertheless, while during the 1980s and the first half of the 1990s there was a close correlation between growth in real loans and real GDP, in recent years a sizeable gap between the two rates has seemed to develop. In Q2 2000 the annual percentage change in real loans reached a historical high of 8.9%, while the annual growth in real GDP stood at 4.0%. The growth differential between the two variables narrowed again only in 2001.

[4] An alternative approach – followed by Kakes (2000) and Hülsewig et al. (2001) – consists of modelling both demand and supply by considering, together with the conventional determinants of loan demand, some indicator of supply-related factors (changes in holdings of securities by banks in the former, the volume of banks' capital and reserves in the latter).

[5] An interesting feature of the historical behaviour of euro area lending cycles is that loans appear to follow a pattern of relatively prolonged periods of gradual acceleration followed by shorter and more abrupt periods of slowdown. One possible explanation for this non-linearity over the business cycle is that in times of slower economic growth, weaker demand for loans is compounded by supply constraints and credit rationing, thereby making the loan slowdown more pronounced.

Figure 2 shows the behaviour of real loans to the private sector and the real composite lending rate (i.e. the nominal composite lending interest rate less annual percentage changes in the GDP deflator). During the 1980s the real lending rate showed relatively limited fluctuations around its historical mean. However, in the 1990s the rate seemed to deviate consistently from its average. In particular, the real lending rate remained above its historical average during the first half of the 1990s (which may to a certain extent explain the low rate of growth of real loans in that period). By contrast, from 1997 onwards it has remained consistently below the sample average, reaching a twenty-year minimum of 4.2% in the third quarter of 2001. During this relatively protracted period of relatively low lending rates, growth in real loans has been consistently above its own average rate.

Sectoral data show that in recent years the behaviour of loans to non-financial corporations has somewhat differed from that of loans to households (see Figure 3). Between 1998 and 1999, annual growth in loans to non-financial corporations remained below that of loans to households, while in 2000 the picture was reversed. Whereas the developments of loans to households over the last few years can be to a large extent explained by the combination of historically low mortgage interest rates and rising housing prices, those of loans to non-financial corporations seem to have reflected an overall increase in the financing needs of euro area firms (as confirmed by the acceleration in debt securities issuance by non-MFIs, especially during 2000).

Some of the main factors that have been suggested to explain the strong demand for credit from non-financial corporations are the following: (1) intense mergers and acquisitions activity by euro area firms; (2) large capital outflows, notably to the US (to a large extent reflecting merger and acquisition activity abroad by euro area companies); and (3) the demand for funds from telecommunications companies to finance the UMTS license auctions, especially in the second and third quarters of 2000 (see ECB (2000)). Empirical findings by De Bondt (2002b) show that, while euro firms make increasing recourse to the corporate debt securities market to finance mergers and acquisitions both domestically and abroad, these tend to be initially financed primarily by means of other sources of corporate finance, such as bridge-financing bank loans.[6]

Overall, the analysis of past developments in loans to the private sector and their main determinants suggests that the growth of real loans over the last two decades was related positively with real GDP growth and negatively with the real composite lending rate. However, other factors, such as mergers and acquisitions activity, direct investment abroad and the interplay between rising property prices and credit or housing purchases, may have also contributed to the high rate of growth in loans in recent years.

[6] Regarding the third factor, Gadanecz (2000) notes that the majority of loans to telecommunications companies in correspondence with the UMTS license auctions were syndicated loans not necessarily involving euro area banks, which suggests that this factor may have influenced the developments of loans granted by euro area MFIs to non-financial corporations in 2000 to a lesser extent.

3. The model

Based on the previous observations and more general theoretical work on the determinants of the demand for loans, this study models the stock of loans to the private sector as a function of prices, a scale variable and the cost of loans. The long run relationship is specified in the following (semi-) log-linear form:

$$(loans - p)_t = \beta_0 + \beta_1 \cdot y_t + \beta_2 \cdot R_t + \beta_3 \cdot \pi_t \qquad (1)$$

where *loans, p* and *y* denote logs of nominal loans to the private sector, the GDP deflator and real GDP; *R* represents the nominal composite lending rate and $\pi = \Delta p * 4$ stands for the annualised quarterly inflation rate.[7] In this equation, loans to the private sector are deflated by the GDP deflator, based on the theoretically plausible hypothesis that in the long run nominal loans are homogeneous with respect to prices. Nevertheless, the inclusion of inflation as an additional variable in the long-run relationship implies that no common factor restriction of short-run price homogeneity is imposed in the model. A further advantage of such inclusion is that it allows to test for the hypothesis of homogeneity between the nominal lending rate and inflation (obtained by imposing the restriction $\beta_2 = -\beta_3$), which is consistent with the argument that economic decisions on investment and consumption are based on real – rather than nominal – interest rates.

Regarding the expected signs of the coefficients in (1), the traditional hypothesis (based on the principle that credit is demanded to finance transactions) suggests that the stock of loans should be positively related to real GDP. For instance, economic agents may require loans for liquidity reasons, or, in the case of firms, for the financing of working capital requirements (see Bernanke and Blinder (1988)). A further argument is that, via its impact on expectations of future activity and profitability, strong economic growth may lead to higher demand for credit to fund spending decisions.[8] Based on the assumption that the nominal interest rate on loans represents a reasonable proxy for the nominal cost of loans, it is rather uncontroversial to expect *R* to enter the long-run demand relationship with a negative sign (i.e. $\beta_2 < 0$).[9]

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[7] In principle, π should refer to the rate of expected inflation. However, since inflation expectations are not directly observable and there are several difficulties in their estimation, the conventional approach in empirical work consists of taking contemporaneous inflation as a proxy (implicitly relying on the assumption that expectations are - on average - in line with the out-turns).

[8] While these arguments point to the existence of a long-run positive relationship between credit and economic activity, some authors (e.g. Bernanke and Gertler (1995)) have suggested that this relationship may also include a counter-cyclical component related to the desire by economic agents to smooth the impact of business cycles on their spending in consumption and investment. In the case of firms, this component relates to the fact that corporations adjust their demand for external sources of finance to "cash flow" developments (see Friedman and Kuttner, 1993).

[9] It has pointed out that the nominal lending rate measures only the observable component of the "true" cost of loans, with the latter cost also depending on additional charges and, more generally, on the set of conditions and other terms (e.g. collateral) attached to the contract (cf. Bernanke (1993), Fase (1995) and Friedman and Kuttner (1993)).

It should be noted, though, that authors adhering to the "credit view" on monetary transmission have argued that the positive (negative) correlation between output (interest rates) and loans cannot in theory be exclusively attributed to factors operating on the demand side. In line with the "balance sheet channel" hypothesis, Bernanke (1993) and Bernanke and Gertler (1995) have suggested that, in the presence of frictions due to information asymmetries or other factors, changes in economic activity and/or policy rates may affect the amount of loans banks are willing to extend (as well as the terms at which these are supplied) via their impact on the balance sheets and financial positions of borrowers.[10] In addition, Bernanke and Blinder (1988) have argued that changes in interest rates may also generate loan supply effects via the "bank lending" channel.[11]

As for the sign of the long-run coefficient of inflation β_3, this cannot be determined a priori. The main argument in support of a positive relationship between inflation and real loans is that a rise in inflation should lead to a fall in the real cost of loans, thereby creating an incentive for households and firms to demand more loans. It has also been argued that if bank lending rates are somewhat sticky relative to changes in prices, increases (decreases) in inflation may prompt economic agents to anticipate (postpone) their investments, given expectations of future variations in lending rates (Howells and Hussein, 1999). On the other hand, in support of the existence of a negative relationship between the two variables, it can be argued (see for example Cuthbertson (1985)) that if the rise in inflation is also associated with large variability of the inflation rate, this may generate uncertainty about the future return on investments and discourage firms from undertaking investments (consequently, reducing their demand for real loans).

4. Empirical analysis

Cointegration analysis

As a preliminary step, the statistical properties of the data are examined using a number of unit root tests. The results (see Table 1) suggest that all the variables in the system should be modelled as I(1) variables over the sample period. The cointegrating properties of the data are subsequently tested by applying the Johansen maximum likelihood-based procedure (see Johansen (1995)) to a vector autoregression (VAR) model for $z_t = [(loans - p)_t, y_t, R_t, \pi_t]'$. On the basis of F-tests of system reduction, the number of lags (*k)* in the VAR is set at five, in correspondence to which the residuals are well-behaved.[12] Based on

[10] For instance, higher interest rates due to monetary policy tightening affect the financial positions and creditworthiness of borrowers by reducing the price of assets apt to be used as collateral, while increasing interest rate payments on their existing short-term or floating-rate debt.

[11] According to the "bank lending" channel, if the central bank implements its policy by draining reserves from the banking system through open market sales, a monetary policy tightening may lead to a contraction in deposits and, provided that banks have limited access to alternative sources of loanable funds, also to lower supply of loans.

[12] The conventional information criteria were not very useful as they suggested lag-lengths (3 for Akaike, 2 for Schwarz and Hannan-Quinn) in correspondence to which the residuals of some of the equations of the VAR were affected by autocorrelation or non-normality. By contrast, for $k = 5$, the results of standard diagnostic tests (not reported for the sake of brevity) fail to show any major signs of misspecification at the conventional significance level.

the trending behaviour of the series considered and the results of the application of the so-called Pantula principle, we select a VAR specification including linear trends in the levels of the data and an (unrestricted) intercept in the cointegrating vector.[13]

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + ... + \Gamma_{k-1} \Delta z_{t-k+1} + \alpha\beta' z_{t-1} + \mu + u_t \qquad (2)$$

where Δ stands for the first-difference operator; Γ_i is the (4×4) matrix of short-run coefficients associated with Δz_{t-i}; α and β represent the $(4 \times r)$ matrices of the loading factors and long-run coefficients, respectively; and μ denotes the vector of deterministic components.

Table 2 reports the results of applying the Johansen cointegration test to z_t over the period Q4 1981 – Q3 2001. While both the likelihood-ratio trace statistic and the maximum eigenvalue statistic reveal the existence of one cointegrating vector at the 1% significance level, the latter would also seem to suggest the presence of a second cointegrating vector at the 5% significance level. However, when the test statistics are adjusted for degrees of freedom to take account of potential sample small bias (which may be advisable given that the sample amounts to roughly eighty observations), they consistently indicate the presence of only one cointegrating vector at the 5% significance level.

Given the variables included in the model, it would have been plausible to expect a second cointegrating vector linking the nominal lending rate with the inflation rate or an equilibrium relation - analogous to an IS curve – between real GDP and the real interest rate. However, the restrictions needed to identify either relationship are clearly rejected by the data.[14] This, together with an additional consideration of a more technical nature,[15] suggests that it may be preferable to assume a cointegrating rank of one (see estimated parameters in Table 3).

Testing for long-run restrictions

Having selected the cointegrating rank, we perform conditional Chi-square tests on restrictions for stationarity, long-run exclusion and weak-exogeneity (see Panel B of Table 2). The test of stationarity by Johansen and Juselius (1992) reveals that the null of stationarity can be rejected at the 1% significance level for all the variables of the system, confirming the findings of the univariate tests for unit roots. The results of the long-run exclusion tests show that none of the variables can be excluded from the

[13] Because of the specification of linear trends in the data, the intercept of the cointegrating vector is not be restricted to the cointegrating space. In the estimation of the model this intercept is combined with those reported for the short-run equations.

[14] In particular, while a relationship linking real GDP, the lending interest rate and inflation is not rejected by the data, it is not possible to find an economically-grounded interpretation for it. As mentioned in Section 2, one possible explanation for this outcome is that there may have been a shift in the mean of the real lending interest rate in the run-up to the start of EMU.

[15] Cheung and Lai (1993) show that the trace statistic is overall more robust to skewness and excess kurtosis of the residuals (which are not uncommon when modelling variables such as inflation and interest rates) than the maximum eigenvalue test.

cointegrating vector at the conventional significance levels. The tests for weak exogeneity reveal that, while y and R can be treated as weakly exogenous for the parameters of interest β , this is not the case for π . This finding implies that the system cannot be reduced in the spirit of the general-to-specific approach to a single equation without incurring a loss of information. One important implication of this result is that it cannot be assumed that, in case real loans deviate from their equilibrium level, the process of return to equilibrium will be necessarily driven by adjustments in loans. Instead, the return to equilibrium will involve also a feedback reaction from inflation.

We also test for the hypothesis of homogeneity between the long-run coefficients of the nominal lending rate and inflation (i.e. $\beta_2 = -\beta_3$) in equation (1). Figure 4 shows that the time path of the recursive test on the restriction lies consistently below the horizontal line denoting the 95% critical value, implying that the restriction holds across different sample periods. In the rest of the paper, the homogeneity restriction on the coefficients of the composite lending rate and inflation is imposed, implying that the lending rate enters the long-run relation in real terms.

Estimating the VEC model

The results of the estimation of the full VEC model are reported in Table 4. Panel A of the table shows the estimated long-run parameters (normalised with respect to real loans). The signs of the coefficients, which are highly significant, suggest that the relationship may describe a demand behavioural relationship. In the long run real loans are positively related to real GDP (with an elasticity of 1.48) and negatively to the real lending rate (with a semi-elasticity of –5.08).[16] The hypothesis of a unitary long-run income elasticity is clearly rejected by the data (the test statistic yields $Chi^2(1)=12.97$ [0.00]). One possible explanation for this is that the coefficient of income might capture the effect of omitted variables, e.g. wealth or non-GDP transactions such as financial transactions and housing purchases, which may also be relevant to explain credit demand.[17]

The estimated semi-elasticity of the real interest rate is somewhat higher than the coefficients on interest rates previously estimated in Calza, Gartner and Sousa (2001).[18] One possible reason for this difference relates to the fact that in this study the cost of loans is measured by a synthetic indicator of bank lending rates instead of market interest rates. While the composite lending rate follows the same long-run trends as market interest rates, these series differ to some extent in terms of their higher frequency dynamics. In

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[16] The estimated long-run income elasticity is fairly close to a previous estimate of 1.34 provided by Calza, Gartner and Sousa (2001) for the euro area. It is also not distant from values recently estimated for individual euro area countries (e.g. 1.76 for the Netherlands by Kakes (2000); 1.11 for Germany in Hülsewig et al. (2001); and values ranging between 1.1 and 1.4 for most euro area countries in Hofmann (2001)).

[17] Some authors recommend that, in order to avoid measurement errors, studies on credit should replace real GDP with variables able to capture spending for items such as financial transactions (including those related to mergers and acquisitions), existing real estate assets, second-hand goods and intermediate output, which may also be financed through loans. See, for example, Howells and Hussein (1999) who conduct an empirical study of the relationship between demand for bank loans and a measure of total transactions derived from data on UK payments clearing systems.

particular, the composite lending rate tends to be steadier than market interest rates, reflecting the sluggish adjustment of retail lending interest rates to fluctuations in market interest rates (see Mojon (2000) and De Bondt (2002a)). An alternative explanation would be that, because of the increased availability of alternative sources of financing over the last few years, the demand for loans to non-financial corporations may have become more interest-rate elastic.[19]

Figure 5 plots the developments of the cointegrating vector $\beta' z$ used as an error-correction term in the model. In general terms, the cointegrating vector represents the deviations of the endogenous variable from its long run equilibrium level. The figure suggests that, while the deviations of the stock of real loans from its equilibrium were mean-reverting, in some occasions they were rather large. For instance, it is possible to identify some points in time in which the four-quarter moving average of the deviations from equilibrium ranged between 6% and 8% of the equilibrium stock of real loans (both in the direction of an overhang and in that of a shortfall) and there is one episode – at the beginning of the 1990s – in which the value was even higher.[20] Overall, the behaviour of the cointegrating vector suggests that, while disequilibria in the market for private loans are eventually corrected, the return to equilibrium may be very slow. This probably reflects the existence of significant transaction costs in the market for loans (see Borio (1996) for a survey on redemption penalties and restrictions on interest rate adjustment in the largest euro area countries).

The coefficient of the error-correction term (ECT) in the equation for real loans is highly significant and negative, providing support to the interpretation of the cointegrating vector as a long-run loan demand relationship (see Panel B of Table 4). The small size of the coefficient (-0.075) is consistent with the evidence of sluggish correction of loan disequilibria. The loading factor in the inflation equation (0.136) is statistically significant, which could be interpreted as an indication that the deviations of real loans from their equilibrium level may contain information about the future developments of inflation (this hypothesis is investigated more formally in section 5).

Evaluating the estimated VEC model

In this section we evaluate the statistical properties of the residuals of the model by means of several standard misspecification tests. On the basis of these tests (reported in Table 5), we can conclude that the model is adequately specified at the conventional significance levels. Parameter constancy tests are also performed to examine the stability of the system. An informal method of investigating the stability of

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[18] The long-run semi-elasticities of the short-term and long-term interest rates were estimated in that study at –1.01 and –1.79, respectively, over the sample period Q1 1980 – Q2 1999.

[19] This argument would imply, though, that the long-run coefficient of the interest rate should have permanently risen during the last few years. However, as shown in the next section, this hypothesis is not confirmed by the time path of the recursively estimated coefficient (which only shows evidence of a temporarily larger coefficient in 2000).

[20] This specific episode is probably related to the developments following German reunification when, a combination of weak economic activity and high levels of interest rates drove the area-wide equilibrium level of loans down, while actual loan growth was supported by large subsidies granted to investments in East Germany.

coefficients consists of examining their recursive estimates to see whether they change significantly over time.

Figure 6 shows that the time paths described by the recursively-estimated coefficients of adjustment α are rather stable, with no major signs of instability over the sample period. Figure 7 shows the recursive estimates of the coefficients of the cointegrating vector (normalised with respect to loans) under the assumption of constant short-run parameters Γ_i. The long-run coefficients of real GDP and the real lending interest rate seem to be fairly stable, with the exception of 2000, when some perturbation seems to take place. As a result of this, the coefficient of the real lending interest rate declines somewhat at the beginning of that year, though during 2001 it appears to revert to its historical level. Similar signs of apparent instability are visible in the income coefficient, though with a smaller effect. The fact that, after the end of 2000, the long-run coefficients appear to return to their historical levels would seem to suggest that the fluctuations experienced by these parameters during that year were caused by factors of a temporary nature.

A formal assessment of whether such fluctuations reflected temporary shocks or more structural ones is presented in Figure 8. The figure plots two alternative statistics (labelled "beta_z" and "beta_r") of the test of parameter constancy by Hansen and Johansen (1993).[21] It should be noted that this is a recursive test of the joint stability of the long-run parameters (as opposed to that of specific coefficients). Both test statistics seem to exclude structural breaks in the parameters of the cointegrating vector during the period analysed, confirming that the fluctuations of the long-run coefficients were probably caused by one or more temporary shocks rather than by factors permanently altering the long-run relationship between real loans and its determinants.

The stability of the short-run coefficients Γ_i of the VEC model is assessed by means of recursive Chow tests for parameter constancy. According to the Chow 1-step ahead test there seems to be evidence of one outlier in the equation for real loans in the second quarter of 2000 (see Panel A of Figure 9). However, the Chow breakpoint and predictive failure tests fail to reveal any evidence of structural instability in the equation in that quarter. The 1-step ahead test also reveals an outlier for the second quarter of 1995 in the equation for inflation (see Panel D of Figure 9). The Chow tests for the system as a whole point to the overall stability at the 5% significance level of the parameters of the model.

To sum up, the estimated loan demand system performs satisfactorily in terms of the statistical properties of its residuals. Nevertheless, the recursive estimates of the long-term parameters reveal that in 2000 some disturbance, probably of a temporary nature, affected the system. One possible explanation

[21] Beta_z corresponds to the case when both the long-run coefficients and the short-run dynamics are re-estimated in each sub-sample period, while in the case of beta_r only the cointegrating vector is re-estimated recursively, while keeping the short-run coefficients at their full sample estimates.

advanced in this work is that these disturbances were related to some factors, probably of a temporary nature, that might have played a non-negligible role in determining the loan developments at the end of the sample period. The main candidates are intense mergers and acquisitions activity (both domestically and abroad), rising real estate prices and increased need by telecom companies to finance the UMTS license auctions (see ECB (2000)).

5. Information content of the loan overhang/shortfall on future inflation

As mentioned above, one specific finding of the empirical analysis in this paper is that the error-correction term of the model enters the inflation equation with a statistically significant and positive coefficient. This suggests that a positive deviation of the loan stock from its equilibrium level (a loan overhang) should directly lead to a higher rate of change in inflation, while a negative disequilibrium (loan shortfall) should have a dampening effect on changes in inflation.

This section aims to investigate more formally whether the estimated error-correction term contains reliable information on future developments in inflation. Two hypotheses are specifically investigated: (1) whether the loan overhang/shortfall has information content on the future *level* of inflation; and (2) whether, consistent with the findings from the model estimation, the loan overhang/shortfall is useful to predict future *changes* in inflation. The empirical investigation presented in this section is based on an approach commonly used in the empirical literature to study the usefulness of the yield spread for predicting future economic real GDP growth (see for instance Hamilton and Kim (2002)).

Hamilton and Kim (2002) examine the information content of the yield spread x_t on y_t^h, the annualised cumulative rate of growth in real GDP (Y_t) over the next h quarters, on the basis of the following regression:

$$y_t^h = \alpha + \beta \cdot x_t + \varepsilon_t \tag{3}$$

where $y_t^h = (4/h) \cdot \ln(Y_{t+h} / Y_t)$. Based on the observation that past growth in a variable can be useful to predict future developments in the same variable, the authors also specify an equation including lagged annualised quarterly growth rates of real GDP:

$$y_t^h = \alpha + \beta \cdot x_t + \gamma(L) \cdot y_t + \varepsilon_t \tag{4}$$

where $\gamma(L)$ is a lag polynomial of order L and $y_t = 4 \cdot \ln(Y_t / Y_{t-1})$.

Applying the methodology to our problem, the indicator variable under investigation is the loan overhang/shortfall from the VECM (i.e. $x_t = ECT_t$).

Under the *first* hypothesis, which suggests that the loan overhang/shortfall has information content on the future inflation rate, equations (3) and (4) can be re-written as:

$$\pi_t^h = \alpha + \beta \cdot ECT_t + \varepsilon_t \tag{3'}$$

$$\pi_t^h = \alpha + \beta \cdot ECT_t + \gamma(L) \cdot \pi_t + \varepsilon_t \tag{4'}$$

where π_t^h denotes the annualised cumulative rate of inflation in the GDP deflator over the next *h* quarters, *ECT* is the estimated loan error-correction term and π_t is the annualised quarterly inflation rate.

Under the *second* hypothesis, which suggests that the loan overhang/shortfall has information content on the future change in the inflation rate, the equations are re-specified as:

$$\pi_t^h - \pi_t = \alpha + \beta \cdot ECT_t + \varepsilon_t \tag{3''}$$

$$\pi_t^h - \pi_t = \alpha + \beta \cdot ECT_t + \gamma(L) \cdot \Delta\pi_t + \varepsilon_t \tag{4''}$$

where Δ is the first-difference operator.

It should be noted that equations (3') and (4') are correctly specified only if inflation can be assumed to be stationary. As mentioned in Section 4, the results of several unit root and stationarity tests (see Table 1 and Panel B of Table 2) tend to suggest that, over the sample period considered, inflation should be treated as I(1). However, it could be argued that these results may not be entirely reliable because of factors such as possible sample size distortions, the strongly downward trending behaviour of inflation over the period considered and the well-known low power of these tests in the presence of structural breaks (see Nicoletti-Altimari (2001) for a discussion).

Nevertheless, if inflation cannot be safely assumed to be stationary over the sample period, the equations corresponding to the second hypothesis – namely, equations (3'') and (4'') – are to be preferred. This would also be more in line with the assumptions made in the construction of the VEC model in previous sections and the empirical finding that one-lagged deviations of the stock of real loans from its equilibrium level have a statistically significant influence on *changes* in the quarterly inflation rate.

Table 6 presents the results of the estimation of equations (3') and (4') for *h* from 1 to 12. As Panel A shows, the coefficient of the loan overhang/shortfall is statistically significant and positive in equation (3') for all values of h, except 12. These results would seem to support the hypothesis that the loan overhang/shortfall has information content on the future level of inflation. However, the estimates of model (4’) reported in Panel B reveal that these results are not robust to the inclusion of own lags of inflation in the model. Indeed, in the case of the augmented model, the loan error-correction term is only significant at the 10% level for *h* equal to one and, somewhat surprisingly, twelve.

The tests of the hypothesis that the loan error-correction term is useful to predict future changes in (as opposed to the level of) the inflation rate are presented in Table 7. As Panel A shows, the coefficient of the loan error-correction term is statistically significant and positive for all values of *h* in equation (3’’). According to the estimates, if loans are initially in equilibrium, the emergence of a loan overhang of 1% should be associated with an increase of just under 0.1% in the annualised cumulative rate of inflation over the next four quarters. These findings are robust to the inclusion in the model of lagged changes in the inflation rate (see Panel B), since also in this case the coefficient of the loan error-correction term is statistically significant at the conventional significance levels for all values of *h*. Another noteworthy observation is the fact that the estimates for the constant term in Table 7 are all statistically significant and negative. This could reflect the downward trend in the inflation rate during the sample period.

The above results for the loan overhang/shortfall complement previous findings by Nicoletti-Altimari (2001) who investigated the leading indicator properties for price developments of changes in loans to the private sector (though there are significant differences in both methodology and sample periods covered). Nicoletti-Altimari (2001) found evidence that changes in (log) nominal loans add predictive power to an univariate model of the inflation *level*, with the stronger gains in predictability occurring at longer forecast horizons. This study also showed that the predictive performance of a univariate model for *changes* in the rate of inflation deteriorates when the second difference in (log) loans is added, with the loss in predictability increasing with the forecast horizon. By contrast, our empirical results suggest that, while the loan overhang/shortfall does not seem to have significant information content as regards the future *level* of inflation, loan disequilibria should help to predict future *changes* in inflation. In addition, the coefficient on the loan overhang declines as the forecasting horizon increases, which could signal a weakening of the effect of the overhang on inflation over time. One possible explanation for the poor performance of the overhang in predicting the *level* of inflation is that, being the overhang a stationary variable, it cannot be expected to be a good forecaster of the level of inflation since the latter seems to behave as a non-stationary variable over the sample periods used.

6. Conclusions

This study shows that a long-run relationship between real loans, real GDP and a newly constructed real composite lending interest rate can be estimated for the euro area. Based on identification techniques conventionally applied in money demand studies, the estimated relationship can be interpreted as a long-run loan demand equation, though it cannot be excluded that this relationship may also capture to some extent supply effects. The study reports the results of the estimation of a vector error-correction model. On the basis of the results of standard misspecification tests, the residuals of this model appear statistically well behaved, while formal parameter constancy tests fail to reveal any evidence of structural breaks at the conventional significance level.

While the results are overall encouraging, the recursive estimates of the long-run parameters (particularly, that of the real composite lending rate) reveal signs of some temporary shocks in the year 2000. One explanation advanced in the study is that developments in loans in 2000 may have been affected upwards by factors not considered in this investigation. These include strong domestic mergers and acquisitions activity, large direct investment outflows from the euro area (especially towards the US), rising property prices and the extra funding needs related to the financing of the UMTS licence auctions.

The analysis of the leading indicator properties of the loan overhang/shortfall on inflation based on the Hamilton and Kim (2002) methodology suggests that this indicator should provide relevant information about future changes in inflation but not on its level. This result may be due to the apparently different statistical properties of inflation and the loan overhang/shortfall. Indeed, while the loan error-correction term exhibits a stationary behavior, the inflation rate seems to behave as a non-stationary variable over the sample period.

It should be noted that the conclusions of this study apply to total loans to the private sector, which implies a high level of aggregation, in particular taking into consideration the likely differences in behavior of the sub-sectors households and non-financial corporations. Therefore, follow-up studies should aim at an investigation of loan demand by sub-sector, seeking a more detailed explanation of the behavior of loans. In addition, supply considerations need to be incorporated into the analysis in order to better discriminate between supply and demand effects. The issue of a potential change in the mean of the real lending rate and its implications for modeling work would also deserve attention.

Data Annex

Data description

This study is based on quarterly data for the euro area – defined on the basis of the principle of current composition (the 11 original countries up to Q4 2000; these plus Greece, thereafter) - over the period Q1 1980 to Q3 2001. Nominal loans are measured by the logs of quarterly averages of end-of-month outstanding amounts of loans to the private sector (seasonally adjusted, EUR billions). Until Q3 1997 the series for loans is based on stock data, from Q4 1997 on flow statistics. The real GDP series (as well as the nominal GDP data used for the calculation of the GDP deflator) are constructed by aggregating logs of seasonally adjusted national accounts data (ESA95 whenever available). The composite lending rate is calculated using both area-wide data on retail interest rates on loans to households and non-financial corporations compiled by the ECB and non-fully harmonised country data on commercial and mortgage loans sourced from BIS and IMF (see below). The inflation rate is given by the annualised quarter-on-quarter change in the GDP deflator. The quarterly values of the inflation rate and the lending interest rate are period averages expressed in decimal points.

Like in all studies using reconstructed historical series for the euro area, one important issue concerns the choice of the method for the aggregation of the national data for the period prior to the adoption of the single currency. In this paper, the national contributions to area-wide data on nominal loans, GDP and prices before 1 January 1999 are aggregated using the irrevocable conversion rates announced on 31 December 1998 (the historical data on the composite lending rate are more difficult to compute and require the use of two distinct aggregation methods). The main reason for using this approach is that it corresponds to the methodology officially adopted by the ECB for the computation of historical series on monetary statistics.[22]

The composite lending rate [23]

The composite lending rate is constructed by aggregating data from different sources. For the recent years, the composite lending rate is computed with area-wide interest rate data compiled by the ECB weighing retail interest rates on different types of loans. These consist of retail interest rates on loans to enterprises up to one year, loans to enterprises over one year, consumer loans to households and loans to households for housing purchases at a monthly frequency.[24] The weights are given by the relative importance of the corresponding loan category and derived from the end-of-quarter outstanding amounts of loans provided by the MFI sector to non-financial corporations and households. Since the quarterly

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[22] See Brand, Gerdesmeier and Roffia (2002) for a discussion on aggregation issues.

[23] The construction of this series is partly based on previous work by one of the authors with C. Martin-Moss and D. Marqués Ibañez (both at the ECB).

[24] Detailed information on the methodological aspects and time coverage of these data are available on the ECB's website at *http://www.ecb.int.*

outstanding amounts of loans necessary for the calculation of the variable weights are only available from the third quarter of 1997, for the period prior to this we assume fixed weights as at the Q3 1997 levels.

As the ECB data do not cover the entire sample period, we extend the composite lending rate backwards using an aggregation of the statistics on national interest rates available on the BIS and IMF's International Financial Statistics databases. These mainly refer to data on interest rates on commercial loans and mortgage loans reported by national sources. As for the aggregation method adopted, the area-wide historical rates have been computed using fixed GDP weights at exchange rates given by PPP in 1995. As a caveat, it is important to note that the procedure used for the compilation of the composite lending rate (based on the combination of ECB area-wide statistics and country historical data on retail lending rates) implies some potential inconsistency. Indeed, the two underlying sets of interest rate data differ in some important aspects, notably the degree of harmonisation and the aggregation method employed for the computation of the historical data.

Statistical properties of the data

The time series properties of the variables included in the system are formally investigated by means of standard unit root tests and a stationarity test. The unit root tests performed – Dickey-Fuller (DF), its Augmented variant (ADF) and Phillips-Perron (PP) – rely on the null hypothesis of non-stationarity. By contrast, the KPSS test suggested by Kwiatkowski et al. (1992) tests the null of stationarity against the alternative hypothesis of non-stationarity. The unit root tests suggest that all the variables in the system should be modelled as I(1); and this is confirmed by the outcome of the KPSS test (see Table 1).

The dataset

Date	Nominal loans to the private sector	GDP deflator	Real GDP	Composite lending rate	Inflation
80Q1	3.17447	3.90036	15.18348	13.06529	NA
80Q2	3.20483	3.91937	15.17627	14.00052	7.60103
80Q3	3.22603	3.93685	15.17262	13.64640	6.99188
80Q4	3.25799	3.95345	15.17403	13.81892	6.64358
81Q1	3.29260	3.97198	15.17725	14.02776	7.41023
81Q2	3.31629	3.99227	15.18151	15.58041	8.11639
81Q3	3.33766	4.01585	15.18189	15.89821	9.43173
81Q4	3.36176	4.03943	15.18478	15.54409	9.43278
82Q1	3.38720	4.05974	15.18896	15.10826	8.12342
82Q2	3.40616	4.07762	15.18898	14.82678	7.15358
82Q3	3.42689	4.09355	15.18456	14.45451	6.37033
82Q4	3.45120	4.11026	15.18545	13.68272	6.68215
83Q1	3.47397	4.13017	15.19331	13.07437	7.96580
83Q2	3.49348	4.14362	15.19823	12.70210	5.37845
83Q3	3.51466	4.16039	15.20116	12.71118	6.71117
83Q4	3.54136	4.17609	15.21315	12.75658	6.27825
84Q1	3.56119	4.18983	15.22190	13.21965	5.49596
84Q2	3.58039	4.20090	15.21582	12.79290	4.42980
84Q3	3.59928	4.21452	15.22821	12.53866	5.44735
84Q4	3.62173	4.22291	15.23207	12.25719	3.35381
85Q1	3.64095	4.23496	15.23257	11.73056	4.81983
85Q2	3.65868	4.24467	15.24335	11.58528	3.88507
85Q3	3.67116	4.25775	15.25371	11.22208	5.23065
85Q4	3.69086	4.26825	15.25944	10.57741	4.20217
86Q1	3.70460	4.28255	15.25817	10.55925	5.71933
86Q2	3.71981	4.29430	15.26888	10.14158	4.70003
86Q3	3.73434	4.30347	15.27651	9.67851	3.66731
86Q4	3.75290	4.31094	15.28148	9.52415	2.98935
87Q1	3.76987	4.31487	15.27523	9.49691	1.57107
87Q2	3.78809	4.32356	15.29356	9.82379	3.47430
87Q3	3.80274	4.32758	15.30082	9.83287	1.61011
87Q4	3.82312	4.33839	15.31357	9.66943	4.32385
88Q1	3.84443	4.34491	15.32068	9.16096	2.60933
88Q2	3.86700	4.35416	15.32897	8.91580	3.69934
88Q3	3.88846	4.36062	15.34060	9.15188	2.58137
88Q4	3.91380	4.37355	15.35113	9.40611	5.17488
89Q1	3.94205	4.38254	15.36358	9.90551	3.59386
89Q2	3.97019	4.39049	15.36858	10.33226	3.18187
89Q3	3.99369	4.40006	15.37614	10.65005	3.82827
89Q4	4.02095	4.41238	15.38735	11.17668	4.92748
90Q1	4.04645	4.42182	15.40171	11.44000	3.77704
90Q2	4.06560	4.43463	15.40757	11.58000	5.12216
90Q3	4.08255	4.44507	15.41553	11.47000	4.17610
90Q4	4.10001	4.45154	15.42291	11.57000	2.58818

continued on the next page

Date	Nominal loans to the private sector	GDP deflator	Real GDP	Composite lending rate	Inflation
91Q1	4.12033	4.46323	15.43462	11.74000	4.67561
91Q2	4.14001	4.47604	15.43649	11.63000	5.12417
91Q3	4.15955	4.48624	15.43705	11.64000	4.08044
91Q4	4.17886	4.50154	15.44557	11.70000	6.11961
92Q1	4.19893	4.50930	15.45868	11.79000	3.10568
92Q2	4.21754	4.51840	15.45195	11.77000	3.63840
92Q3	4.23306	4.52614	15.44985	12.23000	3.09659
92Q4	4.24913	4.53337	15.44669	12.27000	2.89071
93Q1	4.26165	4.54361	15.44041	11.65000	4.09771
93Q2	4.27278	4.55285	15.44063	10.99000	3.69614
93Q3	4.28468	4.55876	15.44484	10.28000	2.36428
93Q4	4.29410	4.56585	15.44865	9.64000	2.83357
94Q1	4.30213	4.57225	15.45743	9.19000	2.56075
94Q2	4.31005	4.57775	15.46216	9.03000	2.20188
94Q3	4.31922	4.58374	15.46933	9.06000	2.39379
94Q4	4.33190	4.59103	15.47752	9.13000	2.91558
95Q1	4.33940	4.59623	15.48352	9.29000	2.08187
95Q2	4.35522	4.60507	15.49003	9.31000	3.53441
95Q3	4.36663	4.61443	15.49142	9.16000	3.74457
95Q4	4.37870	4.61920	15.49284	8.91000	1.90992
96Q1	4.39262	4.62424	15.49594	8.59000	2.01286
96Q2	4.40598	4.62807	15.50186	8.36000	1.53193
96Q3	4.41833	4.63151	15.50700	8.17000	1.37860
96Q4	4.43266	4.63522	15.50884	7.82000	1.48404
97Q1	4.45185	4.63923	15.51313	7.46000	1.60448
97Q2	4.46797	4.64183	15.52381	7.27000	1.03679
97Q3	4.48463	4.64582	15.53007	7.09000	1.59658
97Q4	4.50392	4.65108	15.53956	7.00000	2.10486
98Q1	4.52612	4.65484	15.54854	6.79000	1.50532
98Q2	4.54607	4.65986	15.55229	6.55000	2.00671
98Q3	4.57002	4.66268	15.55691	6.34000	1.12956
98Q4	4.59287	4.66585	15.55804	6.00000	1.26677
99Q1	4.62219	4.66883	15.56759	5.66000	1.19250
99Q2	4.64157	4.67125	15.57307	5.42000	0.96844
99Q3	4.66616	4.67325	15.58361	5.68000	0.79965
99Q4	4.68794	4.67579	15.59349	6.03000	1.01676
00Q1	4.71203	4.67891	15.60381	6.31000	1.24736
00Q2	4.73711	4.68123	15.61245	6.57000	0.92829
00Q3	4.75736	4.68585	15.61715	6.87000	1.84497
00Q4	4.77919	4.68961	15.62287	7.00000	1.50449
01Q1	4.79821	4.69902	15.62917	6.83000	3.76527
01Q2	4.81501	4.70458	15.63031	6.75000	2.22545
01Q3	4.82675	4.70956	15.63131	6.62888	1.99141

Note: Nominal loans to the private sector, the GDP deflator and real GDP are in natural logs; composite lending rate and inflation in percentages per annum.

Tables and Figures

TABLE 1. TESTING THE STATISTICAL PROPERTIES OF THE DATA

Variable	DF		ADF		PP		KPSS (l=9)	
	Constant	Constant and trend	Constant	Constant and trend	Constant	Constant and trend	Constant	Constant and trend
$(loans\text{-}p)$	4.92	0.09	0.63	-3.88*	-1.13	-1.71	0.91**	0.10
$\Delta(loans\text{-}p)$	-3.52**	-3.90*	-2.57†	-2.38	-2.90*	-2.99	0.33	0.09
y	0.21	-1.50	-0.44	-2.67	0.14	-1.79	0.91**	0.12†
Δy	-7.30**	-7.34**	-3.65**	-3.60*	-7.46**	-7.43**	0.09	0.09
R	-1.90	-1.47	-2.82†	-3.08	-1.04	-1.83	0.73*	0.09
ΔR	-6.11**	-6.20**	-4.17**	-4.20**	-4.61**	-4.58**	0.13	0.10
π	-3.16*	-5.15**	-2.43	-2.09	-2.50	-5.06**	0.77**	0.10
$\Delta\pi$	-16.63**	-16.61**	-6.80**	-7.06**	-17.57**	-17.67**	0.17	0.07

*Note: ** Denotes rejection of null hypothesis at 1% significance level, * at 5% significance level, † at 10% significance level, based on critical values by MacKinnon (1991) for DF, ADF and PP, by Kwiatkowski et al. (1992) for KPSS. DF denotes the Dickey Fuller test, ADF the Augmented Dickey Fuller test (with lags up to and including the highest lag statistically significant at least at the 5% level), PP the Phillips Perron test (with 3 truncation lags, as dictated by the Newey-West criterion), KPSS the stationarity test by Kwiatkowski et al. (with l representing the parameter of the Bartlett window). The KPSS test has been run in EViews using a program written by H. Hansen).*

TABLE 2. COINTEGRATION ANALYSIS

A. COINTEGRATION TESTS

Eigenvalue	H_0: rank	Trace	Trace (adjusted for d.o.f.) [a]	95% Critical value [b]	Maximum eigenvalue	Maximum eigenvalue (adjusted for d.o.f.) [a]	95% Critical value [b]
0.392	= 0	68.8**	51.6*	47.2	39.8**	29.8*	27.1
0.237	≤ 1	29.0	21.7	29.7	21.6*	16.2	21.0
0.073	≤ 2	7.4	5.5	15.4	6.1	4.6	14.1
0.016	≤ 3	1.3	1.0	3.8	1.3	1.0	3.8

B. CHI [2] TEST OF RESTRICTIONS (CONDITIONAL ON RANK = 1) [c]

	(loans-p)	y	R	π
Stationarity	32.63**	32.82**	35.86**	34.67**
Long-run exclusion	14.49**	17.66**	17.28**	16.02**
Weak-exogeneity	16.51**	0.15	0.58	13.35**

*Note: ** Denotes rejection of null hypothesis at 1% significance level, * at 5% significance level, † at 10% significance level.*
[a] Small-sample correction by Reimers (1992).
[b] Critical values from Osterwald-Lenum (1992).
[c] See Hansen and Juselius (1995).

TABLE 3. ESTIMATED LONG-RUN PARAMETERS β AND COEFFICIENTS OF ADJUSTMENT α

	A. Normalised β	B. α
(loans-p)	1.000 (-)	-0.071 (0.01)
y	-1.589 (0.09)	0.008 (0.02)
R	5.047 (0.74)	-0.006 (0.01)
π	-5.862 (0.86)	0.143 (0.03)

Note: standard errors in parentheses.

TABLE 4. ESTIMATING THE VEC MODEL

A. COINTEGRATING EQUATION

	(loans-p)	*y*	*(R-π)*
	1.000	-1.485 (34.97)	5.083 (7.03)

B. DYNAMIC EQUATIONS

	Δ(*loans-p*)	$\Delta(y)$	$\Delta(R)$	$\Delta(\pi)$
ECT_{t-1}	-0.075 (-5.93)	-0.003 (-0.14)	-0.004 (-0.46)	0.136 (3.87)
$\Delta(loans\text{-}p)_{t-1}$	0.144 (1.15)	-0.144 (-0.73)	0.104 (1.15)	0.675 (1.96)
$\Delta(loans\text{-}p)_{t-2}$	0.294 (2.12)	0.080 (0.37)	0.024 (0.24)	-0.256 (-0.67)
$\Delta(loans\text{-}p)_{t-3}$	0.114 (0.81)	0.120 (0.54)	-0.092 (-0.91)	-0.017 (-0.04)
$\Delta(loans\text{-}p)_{t-4}$	0.335 (2.56)	0.176 (0.86)	0.050 (0.53)	0.003 (0.01)
$\Delta(y)_{t-1}$	-0.073 (-0.79)	0.074 (0.51)	-0.008 (-0.12)	-0.093 (-0.36)
$\Delta(y)_{t-2}$	-0.078 (-0.84)	0.016 (0.11)	0.102 (1.54)	0.462 (1.83)
$\Delta(y)_{t-3}$	0.005 (-0.06)	0.130 (0.86)	0.106 (1.53)	0.120 (0.46)
$\Delta(y)_{t-4}$	-0.104 (-1.13)	0.346 (2.41)	-0.009 (-0.14)	0.503 (2.01)
$\Delta(R)_{t-1}$	0.231 (1.18)	-0.329 (-1.07)	0.577 (4.09)	-0.572 (-1.07)
$\Delta(R)_{t-2}$	-0.069 (-0.38)	0.009 (0.03)	-0.101 (-0.79)	0.567 (1.17)
$\Delta(R)_{t-3}$	0.208 (1.24)	-0.223 (-0.84)	0.012 (0.09)	-0.747 (-1.62)
$\Delta(R)_{t-4}$	0.045 (0.31)	-0.343 (-1.51)	-0.101 (-0.97)	-0.782 (-1.97)
$\Delta(\pi)_{t-1}$	-0.318 (-4.10)	0.018 (0.15)	-0.020 (-0.37)	-0.015 (-0.07)
$\Delta(\pi)_{t-2}$	-0.304 (-4.02)	-0.045 (-0.38)	-0.007 (-0.13)	0.234 (1.13)
$\Delta(\pi)_{t-3}$	-0.258 (-3.59)	0.033 (0.29)	-0.008 (-0.16)	0.290 (1.48)
$\Delta(\pi)_{t-4}$	-0.106 (-2.31)	0.037 (0.51)	0.039 (1.18)	0.179 (1.42)
Constant	-1.715 (-5.93)	-0.064 (-0.14)	-0.099 (-0.47)	3.057 (3.86)
R^2 adjusted	0.75	0.10	0.49	0.50
S. E. of regression	0.31%	0.48%	0.22%	0.84%
Sum sq. residuals	0.06%	0.15%	0.03%	0.44%

Note: t-statistics in parentheses.

TABLE 5. PROPERTIES OF RESIDUALS OF THE VEC MODEL

A. SINGLE EQUATION STATISTICS

	$\Delta(loans\text{-}p)$	$\Delta(y)$	$\Delta(R)$	$\Delta(\pi)$
LM(1) F(1, 61)	0.03 [0.87]	2.05 [0.16]	0.28 [0.61]	0.24 [0.62]
LM(4) F(1, 61)	0.56 [0.46]	0.02 [0.90]	1.22 [0.27]	0.48 [0.49]
LM(1-5) F(5, 57)	0.13 [0.99]	0.48 [0.79]	0.49 [0.78]	0.19 [0.97]
NORM $Chi^2(2)$	0.02 [0.99]	2.99 [0.22]	1.93 [0.38]	3.83 [0.15]
ARCH(4) F(4, 54)	0.38 [0.82]	0.62 [0.65]	0.19 [0.94]	2.38 [0.06]
HET F(34, 27)	0.39 [0.99]	0.75 [0.79]	0.63 [0.90]	0.84 [0.69]

B. MULTIVARIATE STATISTICS

LM(1) F(16, 168) = 0.84 [0.64]; LM(4) F(16, 168) = 0.93 [0.53]; LM(1-5) F(80, 156) = 0.94 [0.61]; NORM $Chi^2(8)$ = 10.98 [0.20]; HET F(340, 201) = 0.65 [1.00]

Note: p-values in square brackets. LM denotes the Lagrange-multiplier Godfrey test for autocorrelation, NORM the Doornik and Hansen normality test, ARCH a test for autoregressive conditional heteroscedasticity, HET heteroscedasticity tests (White tests for the individual equations and the Doornik and Hendry test for the system as a whole). For details on these tests see Doornik and Hendry (1994) and the references cited herein.

TABLE 6. INFORMATION CONTENT OF LOAN OVERHANG/SHORTFALL ON FUTURE INFLATION

PANEL A

Equation (3'): $\pi_t^h = \alpha + \beta \cdot ECT_t + \varepsilon_t$

h	α	β	R^2adjusted
1	0.032**	0.077*	0.065
2	0.032**	0.071*	0.064
3	0.033**	0.077*	0.078
4	0.033**	0.081*	0.087
5	0.033**	0.079*	0.080
6	0.034**	0.076*	0.068
7	0.034**	0.083*	0.081
8	0.035**	0.080*	0.071
12	0.036**	0.046	0.014

PANEL B

Equation (4'): $\pi_t^h = \alpha + \beta \cdot ECT_t + \gamma(L) \cdot \pi_t + \varepsilon_t$

h	α	β	γ_1	γ_2	γ_3	γ_4	R^2adjusted
1	0.008**	0.029†	0.510**	0.237**	0.014	-0.053	0.528
2	0.008**	0.008	0.412**	0.168*	-0.016	0.119†	0.589
3	0.008**	0.012	0.333**	0.164**	0.099	0.078	0.598
4	0.008**	0.012	0.277**	0.227**	0.085	0.074	0.628
5	0.008**	0.008	0.299**	0.169**	0.078	0.124†	0.661
6	0.007**	-0.002	0.251**	0.156*	0.127*	0.143*	0.682
7	0.007**	-0.003	0.239**	0.171**	0.120*	0.138*	0.703
8	0.007**	-0.006	0.248**	0.161**	0.120*	0.127*	0.719
12	0.007*	-0.003†	0.296**	0.126*	0.085†	0.145*	0.716

*Note: **, * and † denote statistically significant at the 1%, 5% and 10% critical levels, respectively. The order L of the lag polynomial $\gamma(L)$ is set at 4 following Hamilton and Kim (2002).*

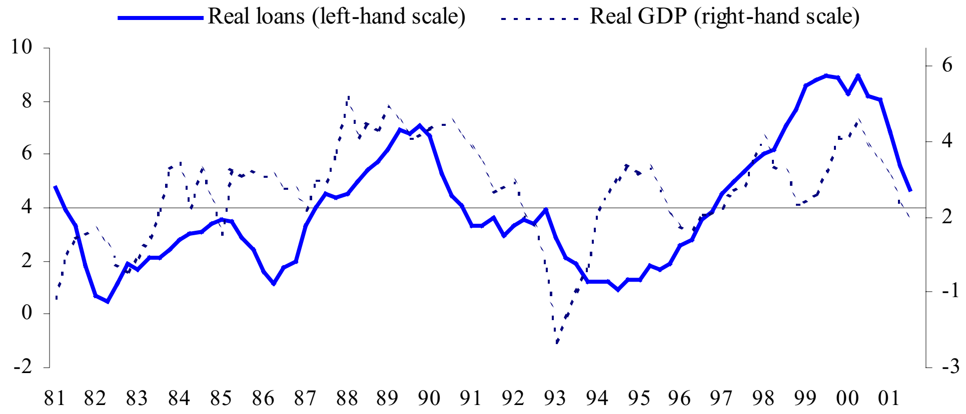
TABLE 7. INFORMATION CONTENT OF LOAN OVERHANG/SHORTFALL ON FUTURE CHANGES IN INFLATION

Equation (3''): $\pi_t^h - \pi_t = \alpha + \beta \cdot ECT_t + \varepsilon_t$			
h	α	β	R^2adjusted
1	-0.000	0.113**	0.286
2	-0.001	0.094**	0.318
3	-0.001	0.091**	0.297
4	-0.002	0.087**	0.263
5	-0.002†	0.088**	0.242
6	-0.003*	0.079**	0.192
7	-0.003*	0.075**	0.167
8	-0.004*	0.070**	0.140
12	-0.006**	0.054**	0.066

Equation (4''): $\pi_t^h - \pi_t = \alpha + \beta \cdot ECT_t + \gamma(L) \cdot \Delta\pi_t + \varepsilon_t$							
h	α	β	γ_1	γ_2	γ_3	γ_4	R^2adjusted
1	0.000**	0.105*	0.328	0.197	-0.020*	-0.180*	0.339
2	-0.000**	0.084**	0.244	0.061	-0.074*	-0.071	0.325
3	-0.000**	0.086*	0.186	0.045	-0.014	-0.052	0.286
4	-0.001**	0.084*	0.152	0.046	-0.046	-0.083	0.254
5	-0.001**	0.083**	0.169	0.021	-0.067†	-0.064	0.250
6	-0.002**	0.082**	0.138	0.034	0.015	-0.004	0.206
7	-0.002**	0.082**	0.144	0.073	0.060	0.028	0.176
8	-0.003**	0.076**	0.170	0.099	0.078	0.028	0.152
12	-0.006**	0.048*	0.048	-0.137	-0.111†	-0.054	0.039

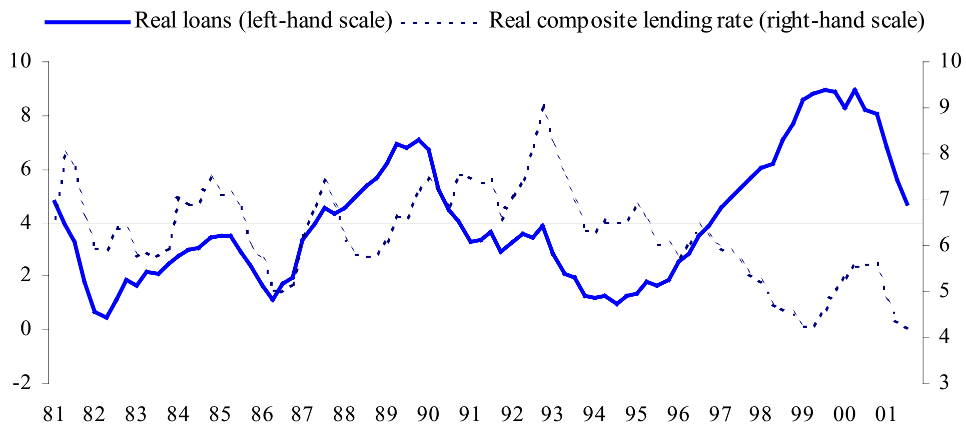
*Note: **, * and † denote statistically significant at the 1%, 5% and 10% critical levels, respectively. The order L of the lag polynomial $\gamma(L)$ is set at 4 following Hamilton and Kim (2002).*

Figure 1. Growth Rates of real Loans to the Private Sector and Real GDP in the Euro Area (Annual Percentage Changes)



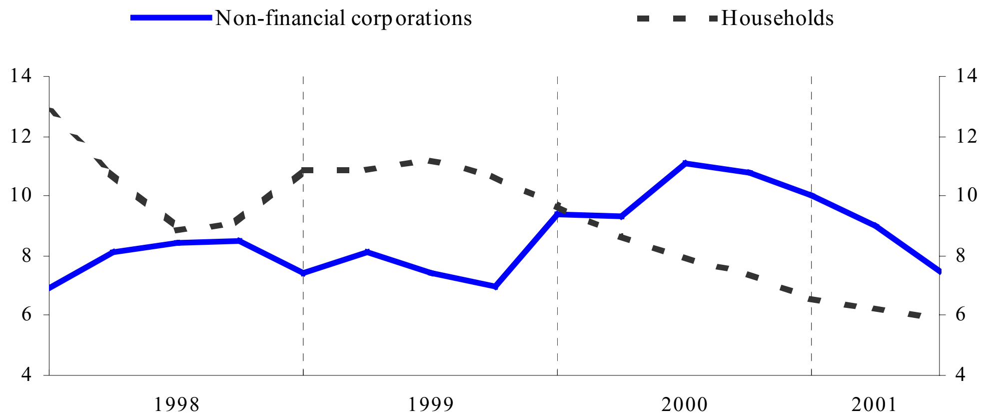
Note: Horizontal line corresponds to sample-period average of plotted variables. GDP deflator used as a measure of the price level.

Figure 2. Real Loans to the Private Sector and the Real Composite Lending Rate in the Euro Area (annual Percentage Changes, Percentages per Annum)



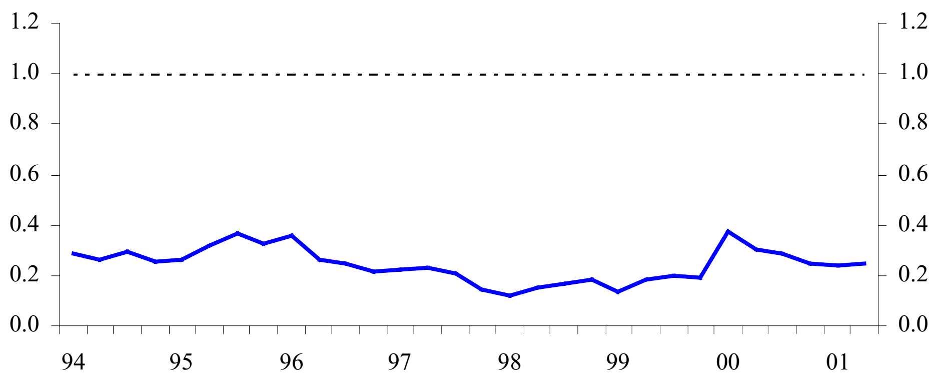
Note: Horizontal line corresponds to sample-period average of plotted variables. GDP deflator used as a measure of the price level.

FIGURE 3. GROWTH IN LOANS TO HOUSEHOLDS AND NON-FINANCIAL CORPORATIONS IN THE EURO AREA (ANNUAL PERCENTAGE CHANGES)



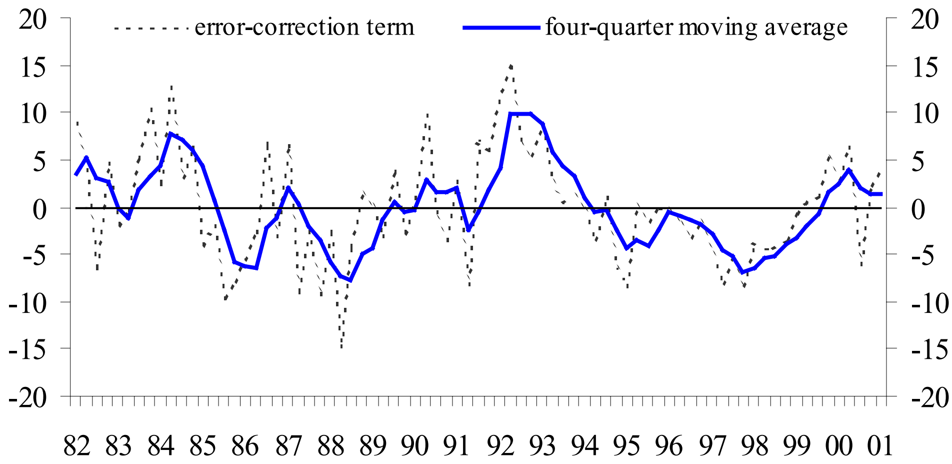
Note: The household sector also includes non-profit institutions serving households.

FIGURE 4. RECURSIVE CHI-SQUARE TEST FOR RESTRICTIONS ON COEFFICIENTS



Note: the dotted line denotes the 95% critical value; a value of the test statistic higher than 1 would imply the rejection of the null hypothesis at the 5% significance level. The test statistic is distributed as Chi^2(1).

FIGURE 5. COINTEGRATING VECTOR USED AS A LOAN-DEMAND ERROR CORRECTION TERM



Note: Deviations are measured as a percentage of the equilibrium stock. The error-correction term has been de-meaned.

FIGURE 6. RECURSIVE ESTIMATES OF THE COEFFICIENTS OF ADJUSTMENT (INITIALIZATION = 50 OBS.)

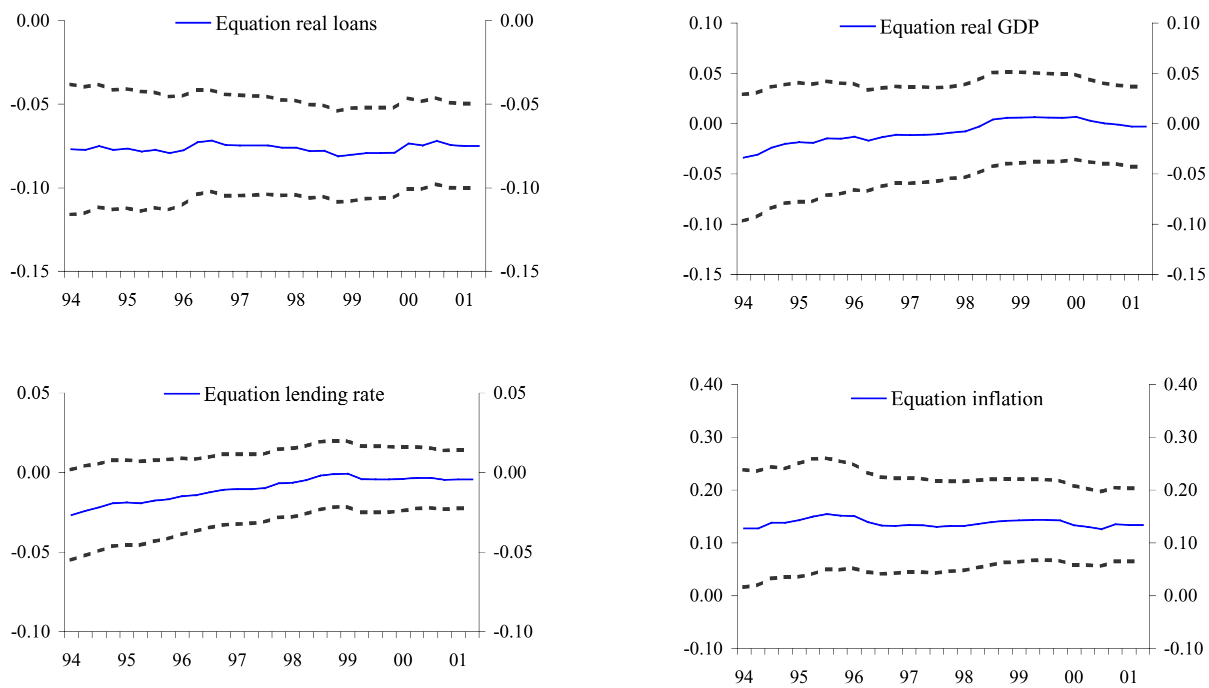


FIGURE 7. RECURSIVE ESTIMATES OF THE LONG-RUN COEFFICIENTS (INITIALIZATION = 50 OBS.)

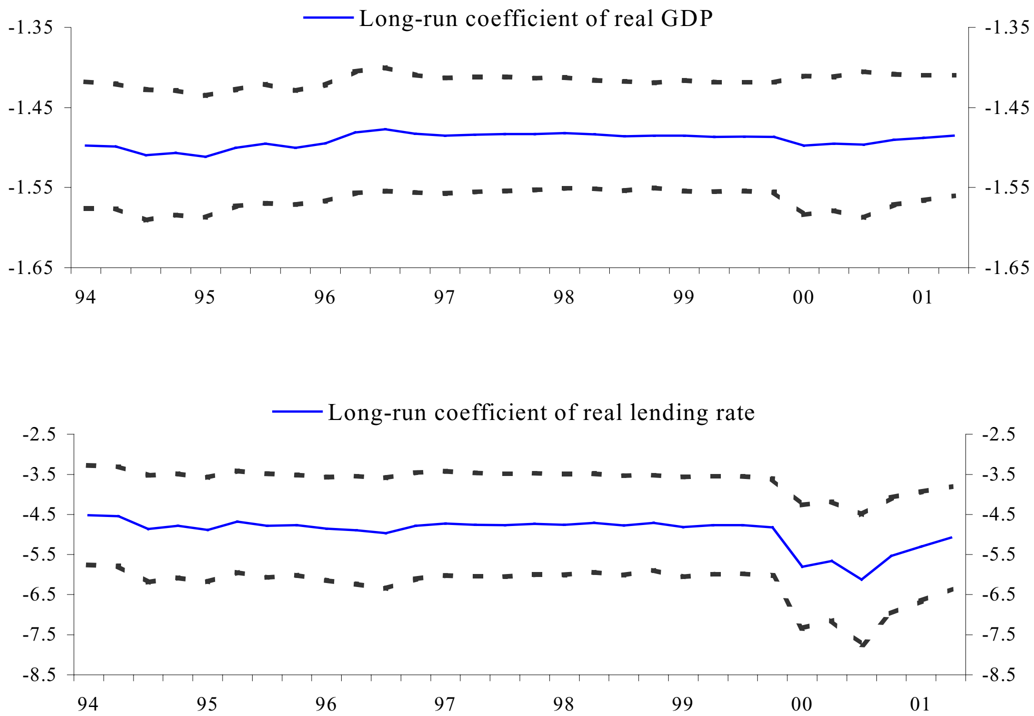
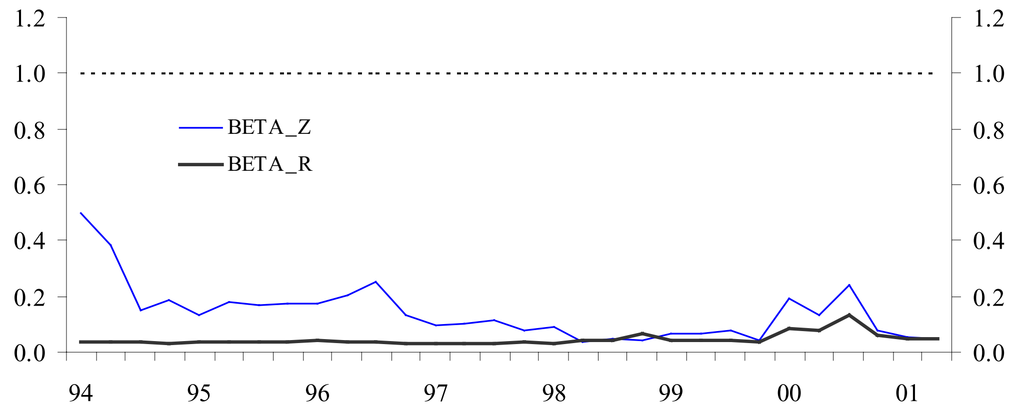


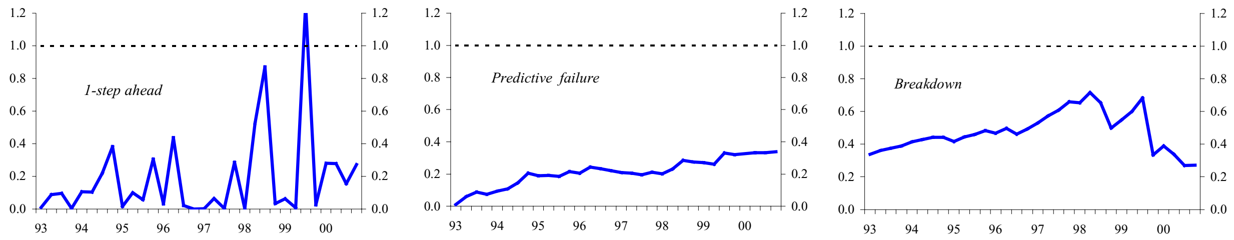
FIGURE 8. HANSEN AND JOHANSEN PARAMETER (1993) CONSTANCY TEST (CONDITIONAL ON $r = 1$)



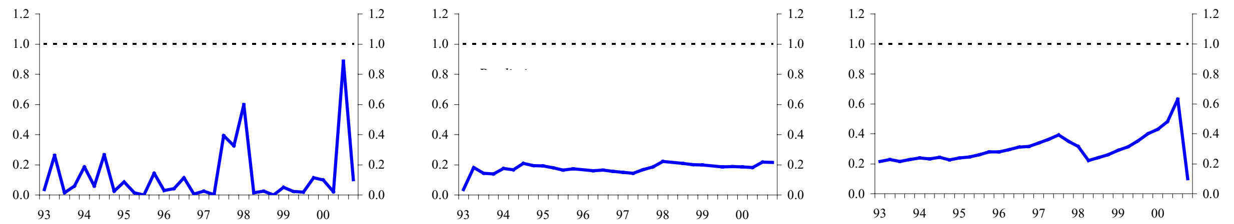
Note: The null hypothesis of the test is that the full sample estimate of the cointegrating vector lies in the space spanned by the cointegrating vectors obtained in each sub-sample. If this holds for all estimated sub-samples we can conclude that the cointegrating vector is constant over time. The dotted line denotes the 95% critical value; a value of the test statistic higher than 1 would imply the rejection of the null hypothesis at the 5% significance level. The test statistic is distributed as Chi^2(4).

FIGURE 9. CHOW TESTS FOR PARAMETER CONSTANCY (INITIALISATION = 44 OBS.)

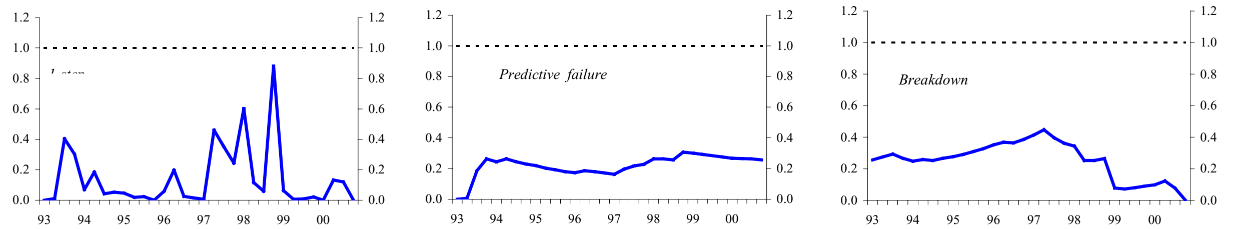
A. Real loans equation



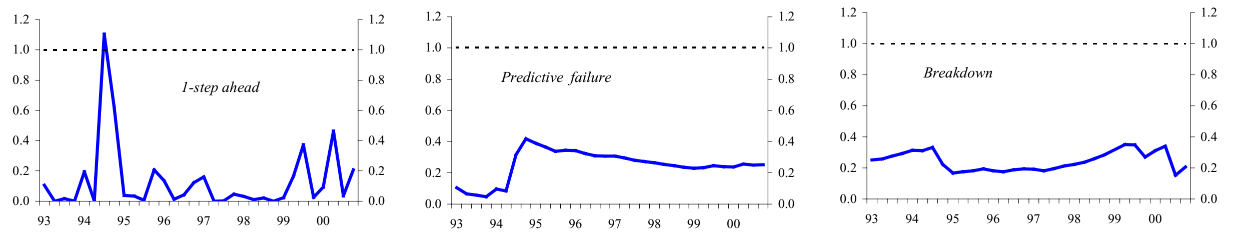
B. Real GDP equation



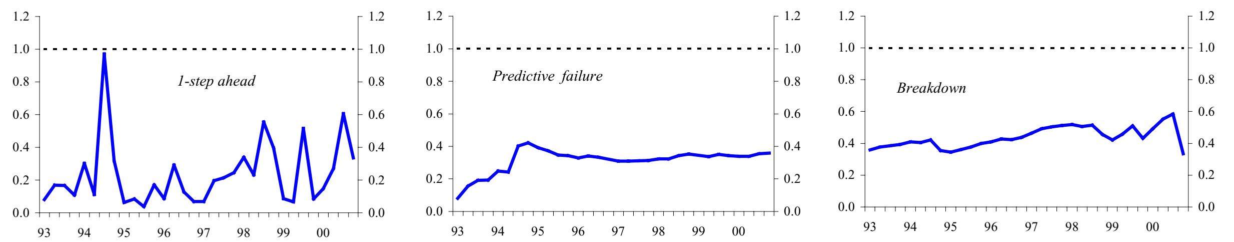
C. Composite lending rate equation



D. Inflation equation



E. System



Note: the dotted line denotes the 95% critical value; a value of the test statistic higher than 1 would imply the rejection of the null hypothesis at the 5% significance level. The test statistics are F-distributed.

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