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In 2005 all ECB publications will feature a motif taken from the €50 banknote.

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Abstract

This paper investigates whether output and inflation respond asymmetrically to credit shocks in the euro area. The methodology, based on a non-linear VAR system, follows work by Balke (2000) for the US. The results reveal evidence of threshold effects related to credit conditions in the economy. Consistent with this finding, the impulse responses show some signs of asymmetric responses over the lending cycle.

JEL classification: E51, C15, C32

Keywords: credit, euro area, asymmetric shocks, non-linearities

Executive Summary

A substantial amount of resources has been historically dedicated to the study of the relationship between money and output. Given that credit represents the main counterpart of money on the balance sheets of banks, a similar relationship between credit and output should also exist. There are reasons to believe that such relationship may be non-linear. In fact, due to the existence of frictions arising from informational asymmetries and contractual rigidities, credit markets may act as non-linear propagators of the impact of aggregate disturbances to the economy.

Several studies have used regime-dependent econometric models to investigate possible non-linearities arising from credit markets. In particular, a strand of these empirical studies has focused on threshold models which typically split the sample endogenously into regimes of 'tight' and 'loose' credit conditions and search for evidence of asymmetric responses to shocks. The intuition is that if economic agents are constrained in their access to credit, exogenous shocks – for instance, due to changes in credit supply conditions – should have relatively larger effects on spending decisions than in the case of abundant availability of credit. McCallum (1991), Galbraith (1996) and Balke (2000) have found evidence of threshold effects in the relationship between economic activity and money in the US. Similar results have been found for the UK by Atanasova (2003).

This paper uses aggregate data to investigate whether in the euro area there are asymmetries in the response of output and inflation to credit shocks (e.g. to lending conditions) over the lending cycle. The testing strategy follows Balke (2000) and consists of: (1) selecting and estimating a threshold VAR model; (2) testing formally for the presence of threshold effects; and (3) analysing whether (conditional linear and non-linear) impulse responses reveal signs of asymmetric propagation of shocks across the separate regimes identified by the threshold model.

The empirical exercise finds clear evidence of threshold effects related to credit conditions in the economy. The threshold critical value for the quarter-on-quarter growth of real loans is estimated at 0.78%. According to this estimate, a level of the transition variable above this threshold would indicate that the euro area economy is in a regime of relatively 'high' credit growth. By contrast, when the threshold variable falls below the estimated critical value, the economy would enter a regime of relatively 'low' credit growth. Given the identification issues involved by the use of aggregate credit data, it is not possible to unambiguously determine whether the switch to a low credit growth regime reflects the impact of credit supply constraints or weak demand (or both).

Consistent with this finding of threshold non-linearity, the conditional linear impulse responses reveal evidence of asymmetric responses of output and inflation to real credit growth shocks over the lending cycle. The response of real GDP to a positive shock in real loan growth is somewhat

bigger (but less persistent) in the low credit growth regime than in the high credit growth regime. The signs of amplification of credit shocks are clearer in the case of the inflation rate, with a shock to credit growth having larger effects when the economy is in the low regime.

We also use non-linear impulse responses that allow for regime switching throughout the duration of the response. The non-linear responses of output growth to credit shocks are larger and more volatile when the economy is initially in the low regime, but this asymmetry is not very pronounced. The non-linear impulse responses of inflation show more evident signs of asymmetry across regimes, with the response of inflation to the credit shock larger when the economy starts under the low regime.

Overall, the two sets of impulse responses analysed reveal evidence of asymmetric macroeconomic responses to shocks to real credit growth over the lending cycle. However, when one allows for regime switching throughout the duration of the response, this asymmetry seems to have a more limited impact.

1. Introduction

Following early work by Friedman and Schwartz (1963), substantial resources have been devoted to studying the relationship between money and output, using both linear methods and, more recently, non-linear techniques.[1] Given that credit represents the main counterpart of money on the balance sheets of banks, a similar relationship between credit and output should also exist. There are reasons to believe that such relationship may be non-linear. In fact, due to the existence of frictions arising from informational asymmetries and contractual rigidities, credit markets may act as non-linear propagators of the impact of aggregate disturbances to the economy.[2]

In an early contribution, Blinder (1987) suggested that the behaviour of the economy may be 'qualitatively different depending on whether or not the credit constraint is binding'. Several authors have subsequently formalized this notion in the context of general equilibrium models incorporating financial market imperfections in which temporary shocks can generate large and persistent fluctuations in output (e.g. Bernanke and Gertler, 1989; Kiyotaki and Moore, 1997; and Azariadis and Smith, 1998). In particular, Azariadis and Smith (1998) develop an overlapping generations model in which, assuming that investments in capital goods must be credit-financed and that credit markets are affected by adverse selection problems, the economy may lie either in a traditional Walrasian equilibrium regime or in an alternative regime characterised by binding credit constraints. One key feature of the model is that it allows for regime switching between the two possible equilibrium regimes depending on the value of transitional variables, such as 'an index of savers' expectations about credit market conditions' (p. 535).

The work by Azariadis and Smith (1998) provides a theoretical underpinning to those empirical studies at an aggregate level which have used regime-dependent models to investigate the behavior of credit markets. In particular, an important strand of these empirical studies has focused on threshold models. McCallum (1991) first argued that the effect of monetary growth on output in the US is greater when indicators of credit constraints exceed (predetermined) critical thresholds. Following McCallum's (1991) contribution, Galbraith (1996) and Balke (2000) have estimated

[1] See Walsh (1998, pp. 9-39) for a review of the empirical evidence on the relationship between money and output. Vilasuso (2000) and Rothman, van Dijk and Franses (2001) are recent examples of empirical studies modelling non-linearities in such relationship.

[2] See, for instance, Bernanke (1993) and Trautwein (2000) for overviews of the literature on the credit channel of monetary policy transmission. See also de Bondt (2000), Chatelain et al. (2003) and Ehrmann et al. (2003) for empirical evidence on the credit channel in the euro area.

threshold models in which the critical values are endogenously determined. These threshold models typically split the sample endogenously into regimes of 'tight' and 'loose' credit conditions and search for evidence of asymmetric responses to shocks. The intuition is that if economic agents are constrained in their access to credit, exogenous shocks – for instance, to credit supply – should have relatively larger effects on spending decisions than in the case of abundant availability of credit.

The studies by McCallum (1991), Galbraith (1996) and Balke (2000) consistently conclude that there is significant evidence of threshold effects in the relationship between economic activity and money in the US. Similar results have been recently found for the UK by Atanasova (2003).

The purpose of this paper is to use euro area aggregate data to test for the presence of possible threshold effects in the response of output and inflation to credit shocks (e.g. to lending supply conditions). In analogy with the money-output literature, we focus on shocks to quantities. The testing strategy follows Balke (2000) and consists of: (1) selecting and estimating a threshold VAR model; (2) testing formally for the presence of threshold effects; and (3) analysing whether (conditional linear and non-linear) impulse responses reveal signs of asymmetric propagation of shocks across the separate regimes identified by the threshold model. The paper is organised as follows. Second 2 provides a short description of the econometric methodology used in the different phases of the empirical investigation. Section 3 describes the results of estimating the threshold VAR and testing for threshold effects. Section 4 presents the impulse responses of output and inflation to a one-standard error shock to credit growth. Section 5 includes some conclusions and suggestions for further work.

2. Econometric methodology

The first step to study the potential role of credit in the non-linear propagation of shocks is to estimate a two-regime threshold VAR model following the specification:

$$Y_t = \mu^1 + A^1 Y_t + B^1(L) Y_{t-1} + (\mu^2 + A^2 Y_t + B^2(L) Y_{t-1}) I_t + \varepsilon_t \qquad (1)$$

where Y_t is a vector of endogenous variables, $I_t[.]$ is an indicator that takes the value 1 when the d-lagged threshold variable c_t is lower than the threshold critical value γ and 0 otherwise. The indicator $I_t[.]$ acts as a transitional variable identifying two separate regimes on the basis of the value of c_{t-d} relative to γ, an unknown parameter that must be endogenously estimated.[3]

Asymmetry is introduced by allowing for the coefficients of the VAR – the vector of constant terms μ, the matrix A and the matrix polinomial in the lag operator $B(L)$ – to vary across the two separate regimes: μ^1, A^1 and $B^1(L)$ represent the parameters of the VAR in the regime defined by $I_t[.]=0$, while $\mu^1+\mu^2$, A^1+A^2 and $B^1(L)+B^2(L)$ are the parameters in the regime identified by $I_t[.]=1$. By specifying c_t as a function of one of the variables in Y_t, it is possible to model regime switching as an endogenous process determined by movements in the variables forming the model. This implies that shocks to any of the variables in Y_t may - via their impact on the variable underlying c_t - induce a shift to a different regime.

While estimating Model (1), it is important to test formally for the presence of threshold effects (with a linear VAR under the null hypothesis and a threshold VAR under the alternative). One complication is that when the threshold value is unknown, parameter γ is identified only under the alternative, leading to a so-called nuisance parameter problem. A commonly used approach to testing consists of first conducting a grid search over c_t and the possible threshold values, estimating each time the selected specification of the threshold VAR model and computing test statistics on the restriction of equality between the linear and the non-linear models (see e.g. Hansen 1996, 1999; Galbraith, 1996 and Balke, 2000).[4] Because the distributions of the test statistics are non-standard, the p-values are usually derived using the simulation techique proposed by Hansen (1996) to calculate the unknown asymptotic distributions. If formal tests reject the linear model, the threshold critical value is estimated from the grid search as the value from the (trimmed) parameter space γ which minimises the log-determinant of the variance-covariance matrix of residuals.

Finally, it is possible to evaluate whether the dynamics of the economy, particularly the reaction of output and inflation to credit shocks, differ across regimes by means of two complementary sets of impulse response functions: (1) regime-dependent impulse responses and (2) non-linear impulse

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[3] Balke (2000) defines model (1) as a ‘structural’ threshold VAR because, in analogy with the literature on identification in standard VAR models, a block-recursive structure based on a standard Cholesky decomposition is imposed in order to orthogonalise the residuals.

[4] One important testing issue is to restrict the parameter space of γ during the grid search to ensure that each regime contains a minimum number of observations. Hansen (1999) recommends 10% of the total number of observations.

responses. The first set of impulse responses describe the reaction of the system to a shock within each of the regimes identified by the estimated threshold.[5] These impulse responses are regime-dependent in that they are conditional on the system remaining in the regime prevailing at the time of the shock throughout the horizon of the response. While these conditional impulse responses are linear within regimes, they may reveal asymmetries if the responses to analogous shocks differ in terms of size, sign and persistence across regimes. Regime-dependent impulse responses are, therefore, effective tools to characterise the behaviour of the system within each regime. However, they may not be suitable to assess the ultimate macroeconomic impact of a shock if the likelihood of regime switching over the duration of the response is non-negligible. In this case, it may be preferable to consider non-linear impulse responses that do not bind the system to remain within the regime prevailing at the time of the initial shock (see Gallant, Rossi and Tauchen, 1993; Koop Pesaran and Potter, 1996; and Potter, 2000). For instance, a sufficiently large shock to a variable may lead to the economy switching away from the starting regime once its direct and indirect effect feed through and, over time, responses may potentially switch repeteadly between the two regimes.

More generally, the non-linear impulse responses differ from their conventional linear correspondents in that, with regime-dependent coefficients, they depend on the history of the time series as well as on the size and magnitude of the shock. In this paper, generalised impulse responses (GI) under alternative regimes are computed numerically using bootstrap simulations as suggested by Balke (2000). The computation of these GI relies on the definition of an impulse response as a revision in conditional expectations. Namely, the response at horizon k (k=1 to h) of a variable Y to a shock at time t (u_t) is given by the difference between (1) the expected value of variable Y given the shock and conditional on a particular history (Ω_{t-1}) of the shocks at time t-1 and (2) the expected value of Y in the absence of such shock:

$$GI_k = E(Y_{t+k} \mid u_t, \Omega_{t-1}) - E(Y_{t+k} \mid \Omega_{t-1}) \qquad (2)$$

In order to compute each of the expectations an iterative procedure is used. The simulation is performed by assuming that at the time of the shock the model is either in the high or in the low regime. In a first step, a sequence of initial values of the actual and contiguous lagged values of the endogenous variables is chosen, corresponding to a particular history (Ω_{t-1}) falling under one of the two regimes. There are as many sets of initial conditions as observations in the regime for which

5 See Ehrmann, Ellison and Valla (2003) for a discussion of regime-dependent impulse response in the context of Markov-switching VAR models.

the impulse responses are being computed (with the regimes identified on the basis of the results of the estimation of the model). A series of shocks is then randomly selected from the residuals of the system. For each sequence of shocks, a path of the variables of interest is simulated with the model conditional on the particular history under consideration. The model is allowed to change regimes during the simulation period. The simulated path provides one estimate of $E(Y_{t+k} \mid \Omega_{t-1})$. In a second step, the same series of random shocks is used but in this case an extra shock (u_t), equal to a one standard deviation shock to the variable in question, is added at time t to each shock sequence. The simulation of the model with this series of shocks conditional on the specified initial history provides one estimate of $E(Y_{t+k} \mid u_t, \Omega_{t-1})$. The difference between the two simulations provides one simulated value of GI_k. This procedure is repeated 500 times for each set of initial observations. The average of the simulated GI_k provides the final estimate of the generalized impulse responses for horizon k and for a given regime. The confidence bands for each horizon k are then derived from the standard errors of the simulated GI_k assuming that the shocks follow a normal distribution. The procedure is subsequently applied to produce the impulse responses under the other regime.

3. Results

In practical terms, the specification of Model (1) requires several choices: (a) the list of variables to be included in Y_t (and whether in levels or first differences); (b) the threshold variable c_t; (c) the delay *d* of the threshold variable; (d) the lag length of the VAR; and (e) the recursive ordering.[6]

The choice of variables to include in Y_t is based on Calza, Manrique and Sousa (2003), who estimate a vector error correction model of euro area loans to the private sector $(loan)$ deflated by the GDP deflator (p), real GDP (y), annualised quarterly inflation $(\pi = \Delta p * 4)$ and a measure of the average nominal lending interest rate (R) defined as a weighted average of retail bank lending rates to households and firms.[7] The choice between levels and differences depends on the stationarity properties of the data. Based on evidence of non-stationarity in levels (not reported for the sake of brevity), these variables are considered in first differences (see Figure 1). The choice of a specification in first differences instead of levels is important because non-stationarity may induce spurious non-linearity in the estimated system (particularly when it affects the variable underlying

[6] The results presented in this section and the non-linear impulse responses in the next one were mainly obtained using Prof. N. Balke's RATS codes. The conditional impulse responses were produced using Malcolm.

[7] See data appendix for a description of the variables used.

the threshold indicator) and may also affect the regularity conditions required for the application of Hansen's (1996) simulation technique. Thus, the threshold VAR is estimated using quarterly data over the sample period 1981:2-2002:3 for the vector of variables:

$$Y_t \equiv [\Delta(y_t), \Delta(\pi_t), \Delta(loan - p), \Delta(R_t)] \tag{3}$$

Regarding the choice of threshold variable, this is typically specified as a moving average of one of the variables forming Y_t. Since we are interested in studying non-linearities in output and inflation over the different stages of the lending cycle, we consider a moving average of the rate of growth in real loans $\Delta(loan - p)$ as the threshold variable.[8] Because $\Delta(loan - p)$ is endogenous to the system, also movements in output, prices or the lending rate may – via their impact on real loans – determine changes in the threshold variable and possibly induce regime switches.

The length of this moving average is determined jointly with the delay of the threshold variable and the lag structure of the VAR by applying standard information criteria to the models arising from the various possible combinations. In particular, we examine the VAR models derived from all the possible combinations obtained with a lag-length of between 1 and 4, a threshold variable including between 2 and 4 moving average terms and a delay of between 1 and 4 quarters. Tables 1 and 2 present the results for the Akaike and the Schwartz criteria, respectively. Based on the Akaike criterion, the optimal specification should be a 4-lag VAR with the threshold variable consisting of a 4-quarter moving average of the quarterly rate of growth in real loans, delayed by 1 quarter. The Schwartz criterion would suggest to use a 1-lag VAR and to construct the threshold variable as a 2-quarter moving average of $\Delta(loan - p)$, delayed by 2 quarters. For the sake of parsimony, we opt for the more conservative specification suggested by the Schwartz criterion.

Table 3 reports the estimated threshold critical value and the results of formally testing for non-linearity using three Wald test statistics (mean-Wald, exp-Wald, sup-Wald). The threshold critical value for the quarter-on-quarter growth of real loans is estimated at 0.78%. According to this estimate, a level of the threshold variable above 0.78% would indicate that the euro area economy is in a regime of 'high' credit growth. By contrast, when the threshold variable falls below the

[8] Measures of credit conditions used in US threshold models include monetary growth (McCallum, 1991 and Galbraith, 1996), the spread between commercial paper and the Treasury Bill rate, the mix of loans and commercial paper in external finance and the difference between the growth rates in the short-term debt of small and large manufacturing firms (Balke, 2000). With the exception of monetary growth, historical data on these variables are not available for the euro area as a whole.

estimated critical value, the economy would enter a regime of ‘low’ credit growth. On the basis of the estimated γ, the sample period is split between the high and low credit growth regimes according to the ratio 60% to 40%.

It should be noted that, given the identification issues involved by the use of aggregate credit data, it is not possible to unambiguously determine whether the switch to a low credit growth regime reflects the impact of credit supply constraints or weak demand (or both).[9] In this sense, this model cannot be used to provide precise answers to specific questions about whether or not at given points in time there are binding credit constraints in the euro area. However, the model is a useful tool to investigate more formally if there is evidence of asymmetric dynamic behaviour of the euro area economy across different credit growth regimes. Indeed, the results of the Wald tests reported in Table 3 provide strong support to the hypothesis of threshold effects (p-values are nil for all tests), suggesting that linear models may not correctly represent the response of key macroeconomic variables, such as output and inflation, to shocks (this issue is further investigated in the next section by means of non-linear impulse responses).

For illustrative purposes, Figure 2 plots the deviations of the threshold variable from the estimated threshold critical value γ against a measure of the output gap derived from a production function approach model by Proietti, Musso and Westermann (2002). As the figure shows, there appears to be a positive correlation between the two variables: the emergence of a negative output gap is typically associated with a negative deviation of the threshold variable from its critical value. By contrast, when the output gap is positive, the threshold variable tends to be above its critical value. Given the definition of the threshold variable in terms of smoothed growth in real loans, this correlation is likely to reflect a more general relation linking the business cycles to lending cycles.[10]

4. Impulse responses

Having found evidence of asymmetry over the lending cycle in the dynamics of the model, it is important to characterise each regime and to assess the macroeconomic significance of this non-linearity by examining the responses of key macroeconomic variables – aggregate output and inflation – to shocks to loans. Larger responses of output and inflation to shocks under the low

[9] See Cecchetti (1995) for a discussion on identification of loan supply and demand shocks in reduced-form models based on aggregate data.

[10] See Bernanke (1993) and Gertler (1988) on the link between financial structure and macroeconomic activity.

regime than under the high regime would provide evidence in support of the existence of threshold effects related to credit conditions. The rationale for this is that in the high regime, firms are likely to be less dependent on the availability of bank lending and may possibly adjust more easily to a shock to lending by substituting it with other means of financing such as internal finance or issuance of debt securities. By contrast, if the low regime corresponds to a situation where firms are more generally restricted in their access to financing (and this is likely to be the case as the low regime broadly corresponds to periods of negative output gap) then a positive shock to bank lending should have a stronger impact on both growth and inflation.

One important issue for this exercise regards the ordering scheme. We consider the following order: $\Delta(y_t) \to \Delta(\pi_t) \to \Delta(loan - p)_t \to \Delta(R_t)$. This implies that shocks to real GDP affect the other variables in the system contemporaneously, but real output reacts sluggishly to shocks to the other variables. Inflation reacts contemporaneously to shocks to real GDP, but only with a lag to those to credit and interest rates. Credit reacts contemporaneously to movements in all variables, with the exception of interest rates; but shocks to credit affect the real economy indicators only after one quarter. Finally, lending interest rates respond contemporaneously to unanticipated changes in all the other variables of the system, but shocks to lending rates have a delayed impact on the other variables.

The positioning of the real sector variables before those related to the credit market is standard in the empirical literature (see, for instance, Walsh and Wilcox, 1995 and Lown and Morgan, 2002).[11] It reflects the more general assumption that financial markets adjust simultaneously to macroeconomic shocks, but that the real sector reacts only sluggishly to shocks to financial variables.

Starting with the characterisation of the regimes, we estimate a linear structural VAR model in each of the regimes and then compute the corresponding regime-dependent impulse responses. Figure 3 shows the impulse responses of the various variables forming the system to shocks to real loan growth. The response of real GDP to a positive shock in real loan growth is somewhat bigger (but less persistent) in the low regime than in the high regime. The signs of amplification of credit shocks are clearer in the case of the inflation rate, with a shock to credit growth having larger effects when the economy is in the low regime. In order to facilitate comparisons, the responses of loan growth to its own shocks have been normalised across regimes by rescaling the size of the low

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[11] In order to check the sensitivity of the results to the relative positioning of quantities and prices of loans, we also consider the alternative scheme: $\Delta(y_t) \to \Delta(\pi_t) \to \Delta R_t \to \Delta(loan - p)_t$. The results proved robust to the ordering change.

regime shock (0.20%) to that of its high regime counterpart (0.31%). Thus, the responses now represent the impact across regimes of shocks of equal size. The responses of loan growth to its own shocks are roughly of the same magnitude in both regimes, but more persistent in the high regime than in the low regime. There are signs of asymmetries also in the case of the lending rate, with the impact of the loan growth shock being significant only in the low regime.

More formally, we can test whether the responses are statistically significant across the identified regimes.[12] Given that the effects of the shocks are relatively short-lived, we restrict the test to differences in the short-term responses, that is over the period following the shock. The results show that the short-term responses are statistically significant across the two regimes for all the variables.[13]

Figure 4 shows the non-linear responses of changes in real GDP and inflation to a positive shock to real loan growth starting under different regimes. In this case, shifts between the regimes are allowed throughout the duration of the response. The figure provides some evidence of asymmetry across regimes with the response of output growth to shocks larger (and somewhat more volatile) when the economy is initially in the low regime, though this is not very pronounced. In addition, it should be noted that the output response is not statistically significantly different from zero when the credit shock occurs under the high regime. By contrast, the figure reveals that when the initial state is the low regime, credit shocks have a statistically significant positive impact on real GDP, though rather small and short-lived.

Figure 4 also includes the impulse responses of inflation to credit shocks. In this case the asymmetry across regimes is more pronounced. Starting under the high regime, a real credit shock is followed by a positive deviation of inflation from the baseline but this is never statistically significant. However, when the economy starts under the low regime the response of inflation to the credit shock is larger in absolute terms and statistically significant in the short-term.[14]

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[12] The test statistics is given by the ratio between (1) the square of the difference between the means of the bootstrapped responses in the two regimes and (2) the sum of the variances of the responses in the two regimes. This test is distributed as a Chi-square with one degree of freedom.

[13] The results of the tests are as follows: real GDP, 5.64 [p-value=0.02]; inflation, 20.07 [p-value=0.00]; loans, 29.83 [p-value=0.00]; lending rate, 4.70 [p-value=0.03].

[14] It should be noted that we have tested whether there was evidence of sign or size asymmetries by computing non-linear impulse responses to a negative one standard deviation shock and to positive and negative two standard deviation shocks. The results, available upon request, show no signs of these types of asymmetry.

By contrast, there are no major signs of asymmetry in the responses of loan growth and lending rates to credit shocks, with the exception of a slightly larger persistence in the low regime. The responses of loan growth and the lending rate to real credit shocks are statistically significant and fairly persistent in both regimes. The sign of the lending rate responses suggests that the credit shock may be characterised as predominantly demand-driven.

Overall, this impulse response analysis suggests that there is some evidence of asymmetric response of key euro area macroeconomic variables to credit shocks over the lending cycle, but this asymmetry is more limited if we allow for regime switching throughout the horizon of the response. Indeed, while the non-linear response of output and inflation to shocks to real credit is larger when the economy is initially in the low credit regime, it is short-lived and of a relatively small magnitude. This evidence of a limited macroeconomic impact of credit shocks would be consistent with the empirical findings by Lown and Morgan (2002), who – using data at the aggregate level for the US – argue that, while loans tend to be highly sensitive to output shocks, the opposite is not necessarily true.

5. Conclusions

This paper uses aggregate data to investigate whether in the euro area there are asymmetries in the response of output and inflation to credit shocks (e.g. to lending conditions) over the lending cycle. Following the approach in Balke (2000), the paper estimates a threshold VAR model over the period 1981:2 – 2002:3 and analyses the dynamic response of the model to credit shocks by means of impulse responses.

The exercise finds clear evidence of threshold effects related to credit conditions in the economy, suggesting that linear models may not correctly represent the dynamic response of output and inflation to shocks. Consistent with this finding, the regime-dependent impulse response analysis reveals evidence of asymmetric responses to shocks to real credit growth over the lending cycle, with shocks having larger effects - particularly on inflation - when the economy is in the low lending growth regime. However, if one allows for regime switching throughout the duration of the response by using non-linear impulse response functions, this asymmetry seems to have a more limited impact.

Overall, the results of the paper suggest that in the euro area non-linearities arising from credit market imperfections may be less relevant than found for the US in similar studies at the aggregate level (as e.g. in McCallum, 1991; Galbraith, 1996 and Balke, 2000).

One possible explanation for the results is that the level of aggregation of the data used in this analysis may be too high and that, in order to uncover asymmetries arising from capital market imperfections, it may be necessary to consider less aggregated data. In order to shed light on these issues, further research should be undertaken on the asymmetries in the response of the components of GDP to credit shocks as in Angeloni et al. (2003). It may also be important to explore differences across sectors and to replicate the analysis distinguishing between households and firms.

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DATA APPENDIX

This study is based on quarterly data for the euro area – the 11 original countries up to 2000:4; these plus Greece, thereafter - over the period 1981:2 to 2002:3. Nominal loans to the private sector are quarterly averages of end-of-month seasonally-adjusted (s.a.) notional stocks computed by the ECB. GDP data are based on the aggregation of logs of s.a. national accounts (ESA95 whenever available) up to Q4 1998; hereafter, on area-wide Eurostat statistics. Inflation is the annualised quarter-on-quarter change in the GDP deflator. Historical national data on loans and GDP have been aggregated using the irrevocable conversion rates announced on 31 December 1998. The composite lending rate is calculated using area-wide data on bank lending rates on loans to households and non-financial corporations compiled by the ECB and non-fully harmonised country data on commercial and mortgage loans from BIS and IMF. The quarterly values of the inflation rate and the lending interest rate are period averages expressed in decimal points.

Table 1. Akaike information criterion for model selection

Lag-length L = 1	Moving average terms for threshold variables		
Delay *d*	2	3	4
1	-43.8075	-43.8272	-43.7934
2	-43.9287	-43.8834	-43.8648
3	-43.7355	-43.6914	-43.6575
4	-43.6531	-43.6481	-43.6006
L = 2	Moving average terms for threshold variables		
Delay *d*	2	3	4
1	-43.9917	-43.6917	-43.8424
2	-43.9259	-43.9184	-43.9587
3	-43.5706	-43.6061	-43.555
4	-43.6289	-43.5777	-43.5353
L = 3	Moving average terms for threshold variables		
Delay *d*	2	3	4
1	-43.965	-43.7043	-43.883
2	-43.8369	-43.8399	-43.8927
3	-43.8197	-43.7702	-43.762
4	-43.6002	-43.7805	-43.7502
L = 4	Moving average terms for threshold variables		
Delay *d*	2	3	4
1	-44.1434	-43.6458	**-44.1618**
2	-43.887	-44.0413	-43.9046
3	-43.8935	-43.9556	-43.8397
4	-43.6594	-43.7606	-43.814

Note: The AIC for a two-regime threshold VAR with four variables and L lags like Model (1) is given by: $[T\ln(\det(\hat{\Sigma}))+2(2(4(4L+1)+6))]/T$ where $\ln(\det(\hat{\Sigma}))$ denotes the log-determinant of the estimated variance-covariance matrix of residuals and T stands for the number of observations.

Table 2. Schwartz information criterion for model selection

Lag-length L = 1	Moving average terms for threshold variables		
Delay *d*	2	3	4
1	-42.2921	-42.3118	-42.278
2	**-42.4133**	-42.368	-42.3494
3	-42.2201	-42.176	-42.1421
4	-42.1377	-42.1327	-42.0852
L = 2	Moving average terms for threshold variables		
Delay *d*	2	3	4
1	-41.5437	-41.2437	-41.3944
2	-41.4779	-41.4704	-41.5107
3	-41.1226	-41.1581	-41.107
4	-41.1809	-41.1297	-41.0873
L = 3	Moving average terms for threshold variables		
Delay *d*	2	3	4
1	-40.5845	-40.3238	-40.5025
2	-40.4564	-40.4594	-40.5122
3	-40.4392	-40.3897	-40.3815
4	-40.2197	-40.4	-40.3697
L = 4	Moving average terms for threshold variables		
Delay *d*	2	3	4
1	-39.8303	-39.3327	-39.8487
2	-39.5739	-39.7282	-39.5915
3	-39.5804	-39.6425	-39.5266
4	-39.3463	-39.4475	-39.5009

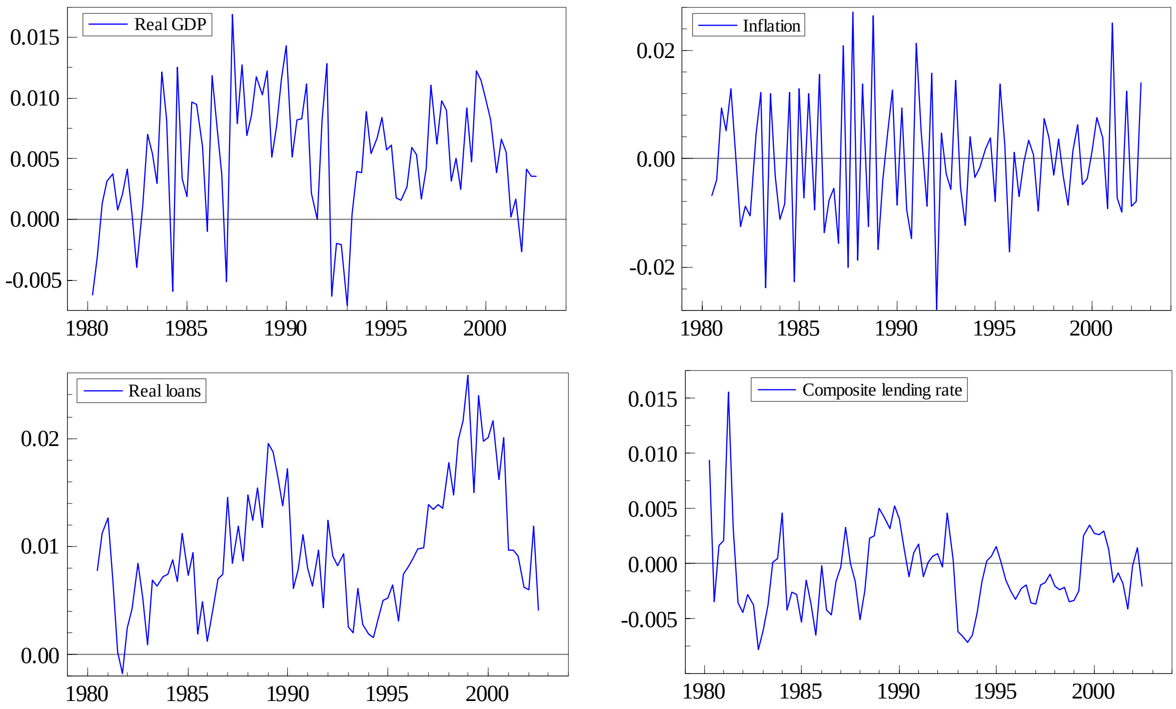
Note: The SIC of a two-regime threshold VAR with four variables and lag-length L like Model (1) is given by: $[T\ln(\det(\hat{\Sigma})) + 2\ln(T)(4(4L+1)+6)]/T$ where $\ln(\det(\hat{\Sigma}))$ denotes the log-determinant .of the estimated variance covariance matrix of residuals and T stands for the number of observations.

Table 3. Estimated threshold critical value and non-linearity tests

$\hat{\gamma}$	Wald tests		
	mean	exp	sup
0.78%	43.71 [0.00]	32.10 [0.00]	73.01 [0.00]

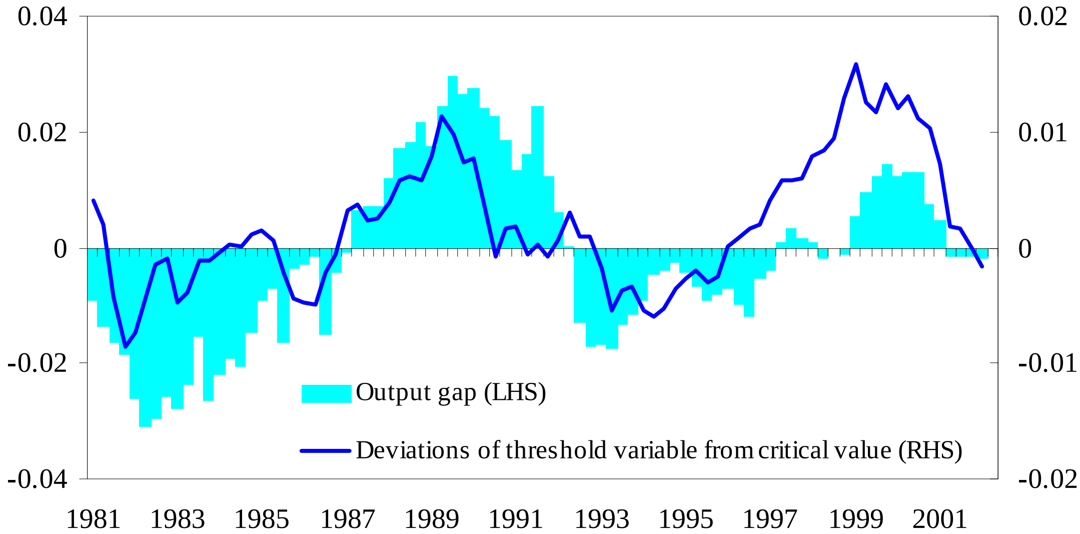
Note: The sample period is 1981:2-2002:3. The grid search over γ is restricted to ensure that each regime contains at least 10% of total observations as suggsted by Hansen (1999). P-values (in brackets) are computed following Hansen (1996), using 1000 replications.

Figure 1. The data (in first differences)



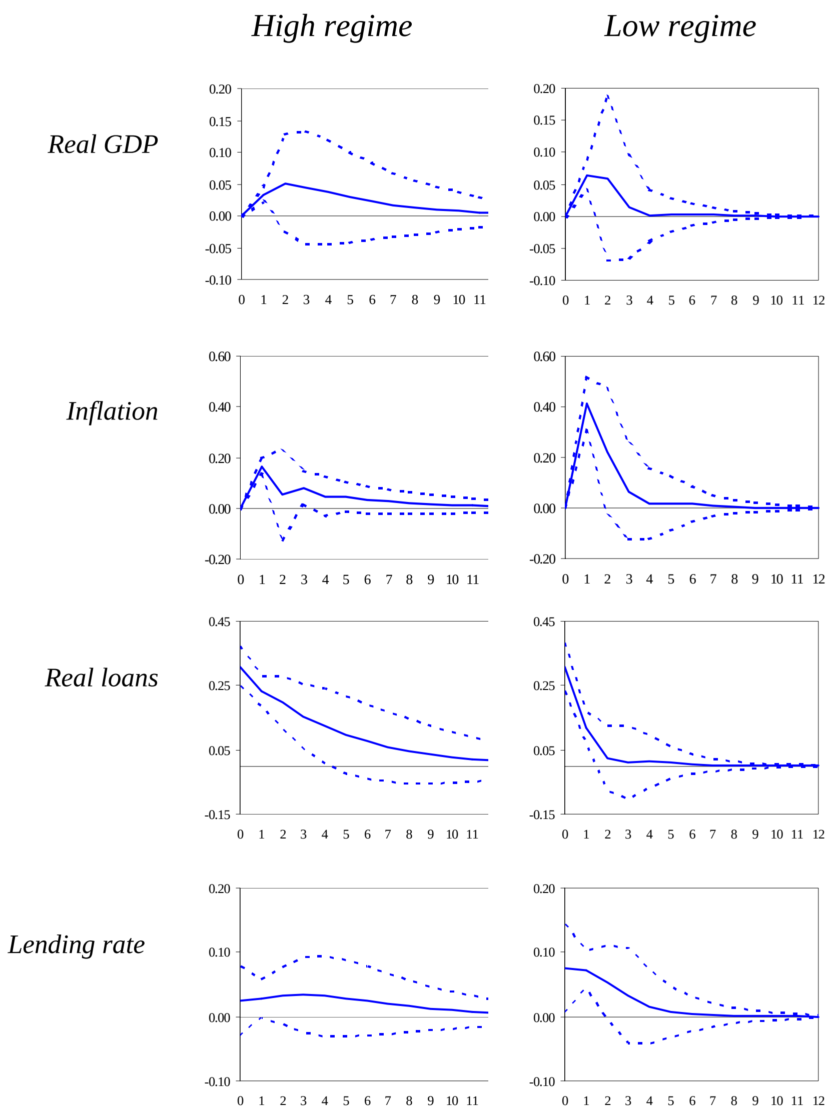
Note: Real GDP and real loans are in natural logs; the price level is the log of the implicit GDP deflator; inflation is the annualised first difference in the price level; the composite lending rate is in levels.

Figure 2. Deviations of the threshold variable from its estimated critical value and output gap



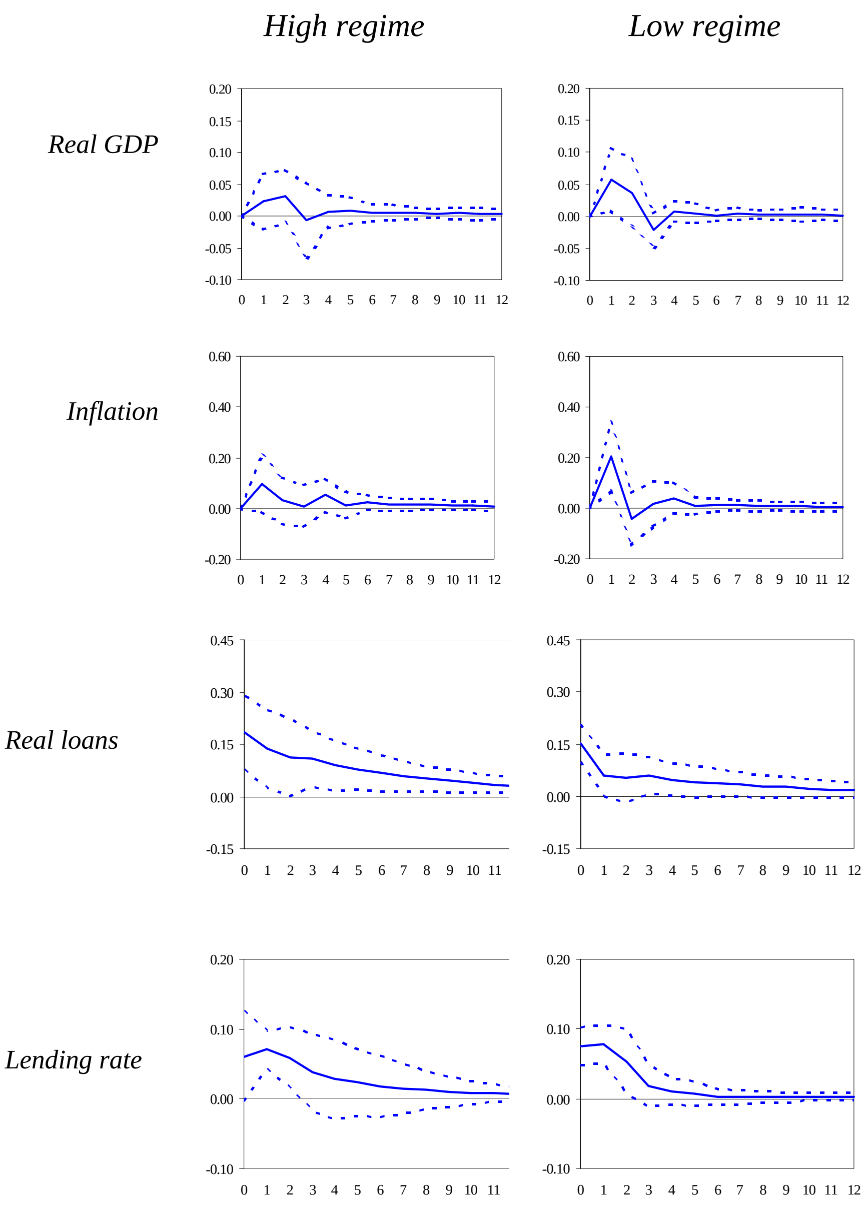
Note: Output gap is derived from the production function approach model featuring pseudo-integrated cycles by Proietti, Musso and Westermann (2002).

Figure 3. Conditional linear impulse responses



Note: The solid lines represent responses to one-standard deviation positive shock to real loan growth. Dotted lines correspond to two-standard error confidence bands. The responses of loan growth to their own shocks are normalised (the low regime shock has been re-scaled to equate the size of the shock in the high regime).

Figure 4. Non-linear impulse responses



Note: The solid lines represent responses to one-standard deviation positive shock to real loan growth. Dotted lines correspond to two-standard error confidence bands.

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