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Abstract
We examine the dynamics of bank lending to the private sector for countries of the Euro area by applying a Markov switching error correction model. We identify for Belgium, Germany, Ireland and Portugal stable, mean reverting regimes and unstable regimes with no tendency to return to the long term credit demand equation, whereas for some other countries there is only weak evidence. Furthermore, for these as well as for other countries we detect in the less stable regimes a strong co-movement with the development of the stock market. We interpret this as evidence for constraints in bank lending. In contrast, the banks' capital seems to have only marginal impact on the lending behaviour.

JEL classification: C32, G21

Keywords: Credit demand, credit rationing, asset prices, credit channel

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* Michael Frömmel, University of Hannover; Torsten Schmidt, RWI Essen. All correspondence to Torsten Schmidt, Rheinisch-Westfälisches Institut für Wirtschaftsforschung (RWI Essen), Hohenzollernstraße 1-3, 45128 Essen, Telefon: +49 201 / 81 49-287, Fax: +49 201 / 81 49-200. Email: tschmidt@rwi-essen.de.

1. Introduction

The strong decline in asset prices followed by an economic slowdown in major economies beginning in 2001 has brought new attention to the real economic effects of asset price bubbles once more (Bordo, Jeanne 2002; Borio, Lowe 2002; Detken, Smets 2004). The major findings are, first, that it is important to consider in which asset markets the bubble occurs: Real effects are particularly severe if a bubble occurs in the real estate market, but also stock prices can have substantial effects. Second, the consequences are the more severe the more private investment is involved. It is therefore important to understand the propagation of an asset price shock to private investment. Third, as a stylized fact asset price slumps and recessions are often accompanied by financial crises. In particular this last finding highlights the importance of financial factors for the transmission of asset price shocks.

For this reason the relation between asset prices and bank lending has often been analyzed for periods of severe economic and financial crisis. Popular examples are the great depression 1929 in the United States (Bernanke 1983, 1995; Eichengreen, Michener 2003), the collapse of real estate and stock prices in Japan in 1990 (Kim, Moreno 1994; Brunner, Kamin 1998), or the East Asia Crisis in 1997 (Stiglitz, Greenwald 2003; Carporale, Spagnolo 2003). These and other empirical studies of asset price bubbles provide details about how financial factors transmit asset price shocks to the real economy (Higgins, Osler 1997). Private investment can be affected at least through two channels: Directly by eroding private firms' value of equity capital, which might be required as collateral for a loan. Hence, firms lending opportunities shrink with subsequent effects on investment. Second, and more indirectly, the drop in asset prices may affect the balance sheets of banks and therefore their lending capacities. In addition, a decline in asset prices may also recluce firms' ability to issue new shares. These arguments contrast with the view that a decline in asset prices reduces firms' investment opportunities by lowering their profitability.

However, as specified by the literature asset price busts are not always accompanied by a financial crisis (Bordo, Jeanne 2002). It is, however, also possible that a restriction in bank lending has dampening effects on economic activity but do not lead to a real recession. Examples of such moderate real effects of credit lending are reported by the literature on the bank lending channel of monetary policy. Related empirical studies for the Euro Area find that changes in interest rates have effects on bank lending only in some countries and with different magnitudes (Altunbas et al. 2002; Angeloni et al. 2003; de Bondt 1999; Kakes and Sturm 2002). Countries therefore where the bank lending channel seems to play some role are Belgium, France, Germany, Greece, Italy, the Netherlands and Portugal. In these countries the link between bank lending and business investment may be important for the transmission of monetary policy (Angeloni et al. 2003). This result suggests that differences in the financial systems of Euro Area member countries also matter for other financial market shocks.

The aim of this paper therefore is to analyze the relation between asset prices and bank lending before and after the stock market crash in 2000 in member countries of the European monetary union. Our main hypothesis is that bank lending is normally related to demand side factors such as GDP and interest rates, but during some shorter periods is determined by other factors. In a first step, we therefore test for cointegration between these variables. This approach is roughly in line with an empirical study for Germany (Deutsche Bundesbank 2002). However, the recently established bank lending survey of the Eurosystem provides some evidence that bank lending at least in some countries was quite restrictive during recent years (Deutsche Bundesbank 2003). As suggested by the literature on asset price bubbles it is likely that bank lending in the Euro Area was affected by the recent decline in asset prices. To test this hypothesis we estimate an error correction equation including the long-run credit demand determinants and changes in share prices and banks' equity capital as short-term determinants. To account for the possibility that bank lending is affected by these short-term factors only during relatively short phases we allow the coefficients of this error correction equation to switch between two states by estimation of a Markov regime switching model. Using this approach it is possible to distinguish between a regime of credit market equilibrium determined by demand factors and a regime were bank lending is affected by asset prices or banks' equity capital.

The outline of this paper is as follows: In section 2 we give a brief overview on empirical research on determinants of bank lending. In section 3 we estimate aggregated credit demand equations for a sample of European countries. In a second step we investigate whether asset prices lead to a regime shift in determinants of credit demand during the asset price bust. In particular we test whether the drop of asset prices is related to a slowdown or decline of bank lending. Section 6 summarizes and concludes.

2. The link between asset prices and bank lending

Empirical studies on the determinants of bank lending find that measures of economic activity, like GDP or industrial production, and interest rates influence the outstanding credit volume. The relation between both variables and loans are usually interpreted as credit demand (Barajas, Steiner 2002; Calza et al. 2003; Ghosh, Ghosh 1999; Pazarbasioglu 1997). A credit demand equation of this form can be derived under simplifying assumptions from a private firm balance sheet constraint (Friedman, Kuttner 1993: 211). The reasoning behind the relation between GDP and loans is as follows: If a firm decides to increase investment outlays over its net revenues it can do it by either raising equity capital or the demand for loans. In the second case, the economic activity indicator can be seen as a broad measure of firms' investment prospects and therefore in this case GDP is a proxy for credit demand. Instead, a reduction of expected revenues also influences firms' access to bank loans. In addition, if interest rates – the price for loans – increase, this will reduce credit demand.

However, most of these empirical studies include variables which are related to credit supply and can be derived from a bank balance sheet (Friedman, Kuttner 1993: 214). Here we find two sets of variables: The first set indicates banks' ability to lend. In this context the most important variables are equity capital of banks and bank deposits. In addition, the portfolio approach suggests that banks aim at holding assets in specific relations to each other. Thus, if the value of one asset changes substantially, the bank will adjust other assets, too, to maintain these relations. Second, the theory of credit rationing highlights the importance of banks' willingness to lend and stresses the importance of non-price variables. These ideas are in line with the findings of empirical studies that interest rates are highly important for credit demand. Indicators of the banks' willingness to lend are related measures for the riskiness of assets and assets which can be used as collateral (Jaffee, Stiglitz 1990).

Recently, the interaction of asset prices, bank lending and investment was analyzed in an extension of the Kiyotaki/Moore model (1997), in which bank capital as well as firms' net worth serve as collateral (Chen 2001). In this model, a reduction in return on investments reduces the net worth of firms as well as banks and therefore constraints the sum of bank loans and investment. In addition, this reduction in investment lowers the prices of collateralized assets and again erodes firms' net worth and banks' lending ability. This line of argument highlights the interaction of asset prices and bank lending reinforcing each other and therefore amplifying the initial shock. The model offers an explanation for the fact that depressions in asset markets are often accompanied by banking crises. Nevertheless, it is not necessary to focus only on financial crises or economic recessions because the strong effect in the Kiyotaki and Moore model relies on some extreme assumptions. Under more general specifications collateral constraints may still amplify unexpected shocks in the economy but real effects are then much smaller (Cordoba, Ripoll 2004; Kocherlakota 2000). Hence, this model already provides some insights for the investigation of the phase after the asset price bust in 2000.

Moreover the approach can be extended by credit rationing, allowing credit market conditions to change over time. Azariadis/Smith (1998) allow credit market conditions to switch between a competitive market allocation and credit rationing. Both regimes depend on private information about loan repayment probabilities. In the following we use this model as a theoretical basing point. It provides a link between theoretical models of credit rationing and empirical studies by allowing demand factors to determine the outstanding credit volume in the regime of credit market equilibrium. On the other hand in the credit rationing regime also supply factors may affect bank lending. This model of switching credit market conditions depicts changes in credit growth over the business cycle but it is also possible to use it as a theoretical basis for analyzing singular events.

Periods in which unusual events cause credit markets to be temporarily in a disequilibrium and restricted from the supply side are often called a credit crunch or capital crunch. The ability or willingness of banks to lend is then reduced: changes in the volume of credits are not in line with equilibrium determinants such as overall economic activity and interest rates. This definition of a credit crunch is closely related to the regime switching model described above, emphasizing the supply side of the loan market. So it is not surprising that empirical studies of credit crunches try to detect changing determinants of the outstanding credit volume (recent examples are Barajas, Steiner 2002; Ghosh, Ghosh 1999; Nehls, Schmidt 2004; Pazarbasioglu 1997). While credit demand is usually assumed to be related to interest rates and economic activity, the main problem is to identify economic variables which only affect credit supply and are related to the source of the credit crunch.

As mentioned above, a first source of a credit crunch can be a burst of an asset price bubble. An often cited example is the banking crisis in Japan after the plunge of asset prices in the stock market and in the housing market in the early nineties (Brunner, Kamin 1998; Kim, Moreno 1994). The case of Japan stresses the crucial role of property as collateral for loans. A significant relation between asset prices and bank lending was found for a number of countries (Hofmann 2004). However, the erosion of collateral was aggravated by the fact that japanese banks were allowed to hold shares. This means that the stock valuations affect bank balance sheets and therefore their ability to lend. The reduction of outstanding credit in the Euro Area after the strong decline in asset prices could be an indicator for a credit crunch at least in some EU countries like Germany (Nehls, Schmidt 2004). Moreover, these factors are likely to play a role in Austria and Belgium, too.

A second trigger of a credit crunch is seen in major reforms of regulation standards for the banking sector. The first Basle Accord for example may have caused the credit crunch in the United States in the early nineties (Bernanke, Lown 1991). In this case the crunch occurred because banks had to improve their balance sheet positions and therefore had to reduce their lending activity for fulfilling the demands by the new regulation. With regard to the importance of banks' equity capital, some authors prefer to call this situation a *capital* crunch. As is stressed in the literature on the bank lending channel the magnitude of the effects of asset price changes on bank balance sheets depend on the structure of the banking system. The more recent discussion of a reform of regulation standards (Basle II) starting in 2000 may have already forced banks to adjust their capital positions (Estrella 2004). Again, the discussion about new capital standards might have led to a new credit crunch. This possibility is in particular important for Germany because it is often argued that the equity capital position of German banks is weak compared to international standards of banks. But it is also likely that the preparation for the new capital standards has lead to adjustments in bank balance sheets in other countries.

In what follows we investigate whether we can identify periods during which outstanding bank credit in several European countries deviates from its long run equilibrium. In addition, we study whether these deviations are related to factors like asset prices or banks equity capital positions which are usually seen as supply side determinants of bank lending.

3. Data description

In this analysis we use quarterly data for the period 1993 to 2004. Credit volume is measured by MFI lending to non-financial corporations which is drawn from national central banks as well as equity capital of banks. Several structural breaks in these series are removed. Nominal GDP and ten year government bond yield are from the EUROSTAT database, except for Portugal where we use data from the OECD database. The share price indices used are those of Morgan Stanley. All variables except interest rates are in logarithms[1]. While the sample period covers the first quarter of 1993 till the fourth quarter of 2004 for most countries, due to reasons of data availability the series are shorter for Greece (1998:3), Ireland (1997:1), Italy (1998:2) and Portugal (1995:1). Furthermore, for Greece and Ireland no data on banks' equities were available to us.

4. The Long-Term Credit Equations

To get an error correction equation for credit demand we use the two step procedure by Engle/Granger (1987). In a first step we estimate the long run credit demand equation. We follow the literature, modeling the demand for loans as a function of economic activity, GDP, and financing costs, which are given by the interest rate. The GDP is expected to have a positive impact on the demand for loans. This textbook effect can be explained by the influence of economic activity on income and profits, and therefore finally on an increased level of investment and consumption (inter alia Kashyap et al.). The interest rate on the other hand reveals a negative impact on the demand of loans, as the demand for loans depends on their price. Our analysis therefore starts with a long-term credit demand equation of the form

[1] Unit root tests for the used series can be found in the appendix.

Table 1

Estimation of long-term credit equations

	ω	α	β
Austria	–0.027 (0.163)	1.090*** (0.103)	–0.029*** (0.010)
Belgium	5.137*** (1.335)	0.570*** (0.117)	–0.021* (0.012)
Finland	5.698*** (0.951)	0.427*** (0.088)	–0.001 (0.010)
France	2.558** (1.119)	0.812*** (0.085)	0.028*** (0.009)
Germany	–11.564*** (2.113)	1.434*** (0.158)	–0.047*** (0.011)
Greece, from 1998:3	–31.932*** (2.510)	4.320*** (0.238)	–0.032** (0.015)
Ireland, from 1997:1	–5.279*** (0.471)	1.654*** (0.042)	–2.31E-05 (0.014)
Italy, from 1998:2	–10.250*** (0.484)	1.850*** (0.038)	–0.001 (0.005)
Netherland	0.826 (0.741)	0.964*** (0.060)	–0.009 (0.011)
Portugal, from 1995:1	–22.703*** (0.550)	3.246*** (0.052)	0.015*** (0.004)
Spain	–1.949*** (0.579)	1.781*** (0.047)	0.004 (0.004)

Significance is given in parentheses. Asterisks refer to level of significance: ***: 1 per cent, **: 5 per cent, *: 10 per cent.

$$k_t = \omega + \alpha \cdot y_t + \beta \cdot i_t + \varepsilon_t \qquad (1)$$

where k_t is the log of the credit volume, y_t the log of GDP, i_t the long-term interest rate and ε_t is an error term.

The estimation results for eq. 1 are shown in Table 1. All estimated coefficients α for the impact of output on credit volume are significant and have the expected sign. For most countries the size of α covers the range between 0.427 (Finland) and 1.850 (Spain). For Portugal and Greece the elasticity of the credit volume with respect to the output are higher and exceed 3.246 and 4.320. All coefficients for the influence of the output, however, are in line with theory, i.e. they have a significantly positive sign, and they are close to the results of other empirical studies[2].

Regarding the influence of interest rates the results are less convincing. The coefficients show the expected negative sign in 8 out of 11 cases, only four of them being significant. For three countries (France, Portugal and Spain) the coefficient is positive, in the cases of Portugal and France even significant. This

[2] For instance Calza et al. (2003) find a coefficient for the whole Euro area of 1.457 for GDP, Bundesbank (2003) find for Germany, depending on the type of loans, between 1.14 and 1.67. Hofmann (2004) finds for various European countries values between 1.269 and 2.169.

observation, however, may be explained by the steadily declining real and nominal interest rates in the Euro area prior to and following the launch of EMU and is in line with other recent studies (for instance Calza et al., 2003). We have also tried other specifications of the long term credit equation, including real variables and GMM estimators with various sets of instrument variables[3]. However, as the results do not substantially change, we rely on the OLS estimation as given in Table 1.

5. Analysis of Deviations from the Long-Term Credit Demand

Assuming that the credit equations from section 4 form a long run equilibrium, one may model the short term dynamic as an error correction model (ECM):

$$\text{(2)} \qquad \begin{aligned}\Delta k_t &= a + b\cdot\left[k_{t-1} - (\omega + \alpha\cdot y_{t-1} + \beta\cdot i_{t-1})\right] + c\cdot\Delta k_{t-1} + u_t \\ &= a + b\cdot\varepsilon_{t-1} + c\cdot\Delta k_{t-1} + u_t\end{aligned}$$

where a, b and c are real numbers, and t is the residual from eq. 1, that is the deviation of credit volume from the long run equation. Furthermore, to take some autocorrelation in the error terms into account, we add the lagged change in credit volume Δk_{t-1} to the equation. While it is common to assume that the error correction coefficient b is always present and constant over time, we use a Markov switching error correction model (MS-ECM)[4], in which the presence or the speed of adjustment may differ depending on the unobservable regime s_t. This state variable s_t may take the values 1 and 2, therefore equation (2) emerges to:

$$\text{(3)} \qquad \Delta k_t = \begin{cases} a_1 + b_1\cdot\varepsilon_{t-1} + c_1\cdot\Delta k_{t-1} + u_t,\ s_t = 1 \\ a_2 + b_2\cdot\varepsilon_{t-1} + c_2\cdot\Delta k_{t-1} + u_t,\ s_t = 2\end{cases}$$

where the state variable st follows a first order Markov process, characterized by the transition probabilities:

$$\text{(4)} \qquad \begin{aligned} p &= P(s_t = 1 | s_{t-1} = 1) \\ 1-p &= P(s_t = 1 | s_{t-1} = 2) \\ q &= P(s_t = 2 | s_{t-1} = 2) \\ 1-q &= P(s_t = 2 | s_{t-1} = 1)\end{aligned}$$

We refer to this model as the basic model. The specification of eq. (3) and (4) provides much flexibility to the estimation: As we do not place any prior as-

[3] The estimation results are available from the authors on request.

[4] This model has been – for instance – applied for estimating bubbles in British house prices (Hall, Psaradakis, Sola 1997), for exchange rates in the European Monetary System (Bessec 2002).

Table 2

Estimation of basic MS-ECM

	Austria	Belgium	Finland[1]	France	Germany
a_1	–0.008*** (0.001)	–0.015*** (0.003)	0.024*** (0.000)	0.008*** (0.003)	0.023*** (0.000)
b_1	–0.597*** (0.144)	–0.090** (0.040)	–0.141** (0.055)	–0.226*** (0.071)	–0.067** (0.025)
c_1	-0.215 (0.236)	–0.419** (0.171)	–	0.361** (0.154)	–0.501*** (0.000)
a_2	0.005*** (0.002)	0.017*** (0.003)	0.000 (0.395)	–0.003* (0.002)	–0.003* (0.002)
b_2	–0.050 (0.031)	0.081 (0.052)	0.010 (0.393)	–0.110*** (0.040)	0.068* (0.035)
c_2	0.539*** (0.125)	–0.246 (0.147)	–	–0.035 (0.195)	0.039 (0.309)
p	0.881	0.835	0.853	0.971	0.981
q	0.980	0.902	0.932	0.972	0.972
Wald 2 ($b_1 = b_2$)	13.833	7.443	3.100	2.079	10.410
Wald 3 ($c_1 = c_2$)	9.553	0.534	–	2.520	2.641

	Greece	Ireland	Italy	Netherland	Portugal	Spain
a_1	0.030*** (0.000)	0.062*** (0.017)	0.062*** (0.017)	–0.038*** (0.009)	0.048*** (0.013)	0.005* (0.003)
b_1	–0.247** (0.035)	–0.685*** (0.193)	–0.684*** (0.119)	–0.335 (0.345)	–0.495*** (0.007)	–0.046 (0.038)
c_1	0.352** (0.011)	0.448 (0.280)	0.448 (0.278)	–0.919** (0.476)	0.269 (0.205)	0.833*** (0.103)
a_2	0.049* (0.067)	0.045*** (0.009)	0.043*** (0.000)	0.013*** (0.004)	0.008* (0.004)	0.068*** (0.009)
b_2	0.007 (0.199)	0.077 (0.121)	0.068 (0.119)	–0.003 (0.055)	–0.102 (0.089)	0.593*** (0.213)
c_2	–0.201 (0.194)	–0.154 (0.185)	–0.232 (0.186)	0.145 (0.163)	0.433** (0.203)	–1.003*** (0.229)
p	0.961	0.770	0.760	0.574	0.894	0.948
q	0.889	0.912	0.899	0.951	0.938	0.876
Wald 2 ($b_1 = b_2$)	1.052	10.259	10.402	0.906	3.976	8.874
Wald 3 ($c_1 = c_2$)	0.427	3.142	4.007	4.277	0.333	52.465

Standard errors are given in parentheses. Asterisks refer to level of significance. – ***: 1 per cent, **: 5 per cent, *: 10 per cent. – Wald 1 (2, 3): Wald test against the null hypothesis that the constant (coefficient for error correction/lagged credit volume) is equal in both regimes. – [1]Model with lagged change of credit volume does not converge for Finland.

sumption on the adjustment process to the long run equation, it may be present or not, or simply differ in speed between the states. Even regarding the number of regime switches we do not need the prerequisite that there is more than one switch. The model is flexible enough to deal with a permanent change[5].

[5] Hamilton (1993) states: “Some might object that a change in regime could be represented as a *permanent* change…, rather than the cycling and back and forth between states 1 and 2 that seems to be implicit in (1.2) [i.e. the Markov chain]. However the specification (1.2) allows the possibility of a permanent change as a special case if p_{21}=0.” (p. 235).

Estimation results of eq. (3) and (4) are presented in Table 2. In the estimation the error correction coefficient b1 respectively b_2 is of main interest. Most countries show one regime which is stabilizing, that is b_1 (we regard to the stabilizing regime as regime 1 without any loss of generality) is significantly smaller than zero. The only exceptions are the Netherlands and Spain, for which b1 is negative without being significant. The results, however, differ for b_2 : For France it is significantly negative, too, indicating that both regimes are characterized by an adjustment to the equilibrium, with different speeds. For Spain and Germany b_2 takes a significantly positive value, whereas for the majority of countries b_2 does not significantly differ from zero at all. It is therefore informative to look at standard Wald tests against the null of $b_1 = b_2$. The results are given in the rows at the bottom of Table 2. Indeed, for all countries but France we find significantly different values of b_1 and b_2 in both regimes. Summing up so far, there is evidence that most countries of the Euro area, France being the only exception, show a stable regime with mean reversion to the long-term credit equation and an unstable regime, which is at its best inconclusive. The differences between b_1 and b_2 are for these countries statistically significant.

It is now straightforward to introduce additional explanatory variables. The rationale is that the credit volume may react in different ways to other factors depending from whether we are in the stable (mean reversion to the long term relationship) or in the unstable regime (no tendency to return). Therefore we extend equation (3) to equation (5)

$$(5) \qquad \Delta k_t = \begin{cases} a + b_1 \cdot \varepsilon_{t-1} + c_1 \cdot \Delta \mathrm{B} C_t + d_1 \cdot shares_t + e_1 \cdot \Delta k_{t-1} + u_t,\; s_t = 1 \\ a + b_2 \cdot \varepsilon_{t-1} + c_2 \cdot \Delta \mathrm{B} C_t + d_2 \cdot shares_t + e_2 \cdot \Delta k_{t-1} + u_t,\; s_t = 2 \end{cases}$$

which additionally takes into account the availability of banks' capital and the influence of asset prices, representing the bank lending and the balance sheet channel of transmission. Whilst the variable ΔBC is defined as the change in logs of the bank capital, the variable *shares* measures the deviation of the respective national stock index from its long term trend[6]. This reflects ongoing under- or overvaluation of companies, which may perpetually prevent banks from expanding their credit offers to companies.

We apply this extended MS-ECM to the countries from table 2. The estimation results are given in table 3. Indeed the estimations provide some additional insights. Results for *Belgium*, *Germany*, *Ireland* and *Portugal* show significant differences (in terms of the Wald test) between the significant ECM

[6] We assume a constant long term rate of return for the stock market. For both variables we tried different specifications, among them a bank capital to credit volume ratio instead of ΔBC and log changes of the stock market index for *shares*. These variations do not substantially change results, but in the case of *shares* the significance suffers.

Table 3

Estimation of the extended MS-ECM

	Austria	Belgium	Finland	France	Germany
Regime 1:					
const.	0.005***	0.009***	0.015***	0.001	0.023***
	(0.002)	(0.002)	(0.003)	(0.002)	(0.001)
b_1 (ECM)	–0.032	0.042	0.032	–0.237**	0.331***
	(0.031)	(0.041)	(0.047)	(0.099)	(0.108)
c_1 (ΔBC)	0.079	0.034*	0.311	0.462	0.406
	(0.057)	(0.017)	(0.452)	(0.299)	(0.256)
d_1 (Shares)	0.013	0.069***	0.051***	0.016	0.067***
	(0.011)	(0.011)	(0.014)	(0.013)	(0.017)
e_1 (ΔK_{t-1})	0.437***	–0.356**	0.215	–0.062*	–2.120
	(0.129)	(0.154)	(0.233)	(0.033)	(1.706)
Regime 2:					
const.	0.005***	0.009***	0.015***	0.001	0.023***
	(0.002)	(0.002)	(0.003)	(0.002)	(0.001)
b_2 (ECM)	–1.111	–0.329***	-0.083	–0.103***	–0.188***
	(1.496)	(0.052)	(0.091)	(0.034)	(0.039)
c_2 (ΔBC)	–0.648	0.017	–0.076**	–0.042	–0.340***
	(2.168)	(0.021)	(0.029)	(0.057)	(0.011)
d_2 (Shares)	–0.297	0.014	–0.007	0.036***	0.017***
	(0.311)	(0.010)	(0.005)	(0.011)	(0.005)
e_2 (ΔK_{t-1})	0.275	–0.324**	0.156	–0.029	–0.260*
	(1.037)	(0.012)	(0.149)	(0.233)	(0.132)
p	0.980	0.805	0.974	0.952	0.961
q	0.870	0.768	0.974	0.976	0.982
Wald 2 ($b_1 = b_2$)	0.520	32.668***	1.143	1.564	19.003***
Wald 3 ($c_1 = c_2$)	0.112	0.427	0.728	2.937	6.605**
Wald 3 ($d_1 = d_2$)	0.994	15.055***	14.257***	1.172	8.683***
Wald 3 ($e_1 = e_2$)	0.024	0.029	0.045	0.018	1.196

	Greece	Irland	Italien	Netherland	Portugal	Spain
Regime 1:						
const.	0.041***	0.046***	–0.001***	0.010**	0.022***	0.023***
	(0.000)	(0.008)	(0.000)	(0.004)	(0.004)	(0.003)
b_1 (ECM)	–0.297**	0.084	–0.018	–0.374**	–0.660**	0.045
	(0.038)	(0.126)	(0.437)	(0.170)	(0.304)	(0.185)
c_1 (ΔBC)	–	–	–0.042	0.004	0.249*	0.019
			(0.459)	(0.107)	(0.145)	(0.033)
d_1 (Shares)	0.008	0.017	–0.067	0.019	0.011	0.002
	(0.241)	(0.017)	(1.257)	(0.016)	(0.019)	(0.013)
e_1 (ΔK_{t-1})	0.321***	–0.223	–0.001	0.031	–0.158	0.354***
	(0.000)	(0.222)	(0.485)	(0.178)	(0.260)	(0.119)
Regime 2:						
const.	0.041***	0.046***	–0.001***	0.010**	0.022***	0.023***
	(0.000)	(0.008)	(0.000)	(0.004)	(0.004)	(0.000)
b_2 (ECM)	–0.170***	–0.578***	–0.005	–0.305**	–0.001	–0.163***
	(0.003)	(0.155)	(0.011)	(0.135)	(0.066)	(0.042)
c_2 (ΔBC)	–	–	0.011	–0.047	0.045	–0.085***
			(0.008)	(0.105)	(0.049)	(0.025)
d_2 (Shares)	–0.006**	0.083***	–0.001*	0.066***	0.051***	0.048***
	(0.046)	(0.030)	(0.000)	(0.021)	(0.009)	(0.008)
e_2 (ΔK_{t-1})	–0.097	0.245	1.045***	0.086	0.314**	–0.219
	(0.180)	(0.177)	(0.018)	(0.195)	(0.128)	(0.140)
p	0.631	0.853	0.915	0.968	0.517	0.971
q	0.783	0.722	0.999	0.962	0.903	0.978
Wald 2 ($b_1 = b_2$)	0.740	10.646***	0.001	0.107	4.280**	1.172
Wald 3 ($c_1 = c_2$)	–	–	0.013	0.130	1.880	6.216**
Wald 3 ($d_1 = d_2$)	2.632*	2.970*	0.003	4.036**	3.356*	8.612***
Wald 3 ($e_1 = e_2$)	52.603***	3.961**	4.647**	0.050	3.359*	18.243***

Standard errors given in parentheses. Asterisks refer to level of significance. – ***: 1 per cent, **: 5 per cent, *: 10 per cent.

coefficient b_1 of the stable regime and the respective coefficient b_2 of the unstable regime. Furthermore, for these countries there are significant differences in the influence of the stock market development between the regimes. For Belgium, Germany and Portugal there is a clear pattern: During unstable periods there is a strong positive influence of stock market movements on the credit volume. This means that deviations from the long-term development of the credit volume go ahead with movements in the stock market: the credit volume tends to decline (increase) when shares prices are below (above) average. During the stable regime, however, a similar behavior for banks' capital seems not to exist, except for Germany. In contrast to these three countries (i.e. Belgium, Germany and Portugal) we find a relation in the opposite direction for Ireland: here the positive relation is in the stable regime, meaning that an increase in stock prices brings the credit volume closer to the long term trend.

For *Austria*, *Finland*, *Italy* and *Spain* the results remain inconclusive. Although there are different degrees of stability between the regimes, most coefficients are not statistically significant and the Wald tests do not indicate significant differences between regimes.

France, *Greece* and the *Netherlands* finally, show two stable regimes with significantly negative ECM coefficients. The two regimes only differ in the speed of adjustment. However, France and the Netherlands show a higher influence of shares prices in the unstable than in the stable regime (significantly different only for the Netherlands). We interpret this as a slightly dampening effect of the stock market on the credit growth.

Summing up so far, we find evidence for the existence of a stable (tendency to return to the credit demand equation) and an unstable (no tendency to return) regime for Belgium, Germany, Ireland and Portugal. Except for Ireland the movement away from the demand equation shoes a positive relation to the movement of the stock market. For Germany we even find some evidence for a relation to the growth bank capital. Furthermore, we even find for France and the Netherlands, for which we can find no unstable regime, that the stock market development affects the speed of adjustment to the long term equation. For the other countries (Austria, Finland, Greece, Italy and Spain) the evidence is inconclusive.

6. Conclusions

This paper analyses the development of the credit volume in the countries of the Euro area. The analysis is based on the construction of a long-term credit demand equation. The actual credit volume is then compared to this long-term trend and deviations are analyzed in detail by applying a Markov switching error corrction model to the deviations. We find for several European countries (Belgium, Germany, Ireland, Portugal and Spain) strong evidence for switching between a stable, mean reverting regime, and one unstable, bubble-like regime, which does not show a tendency to return to the long-run demand equation. While France, Greek and the Netherlands show two regimes with different speeds of adjustments only, the three remaining countries Austria, Italy and Spain show only weak evidence of regime switches.

Furthermore, we find a positive relation between the credit volume and the development of the stock market in the unstable regime for Belgium, Germany and Portugal, as well as for Finland, France and the Netherlands in the regime with lower adjustment speed, which is significantly higher than in the other regime. A significant influence of the banks' capital, however, is only visible in the case of Germany.

From these results we conclude, that there are constraints, which may temporarily dampen lending of banks to the private sector, and even drive away the credit volume from the long term demand equation. Whether these constraints lead to a real lack of credits, or only affect the adjustment speed, seems to depend on country-specific characteristics. A crucial role is attributed to the development of stock markets and therefore our analysis is linked to the literature on the credit channel. The results, however, do not clearly indicate, whether the impact of the asset price growth works via the supply side or the demand side. In contrast, the availability of bank capital seems to play only a minor role in most countries, except of Germany. Our findings even affect monetary policy, which needs to take into account the development of asset prices when conducting monetary policy, because there seems to be a substantial, and between countries of the Euro area asymmetric effect on the transmission mechanism.

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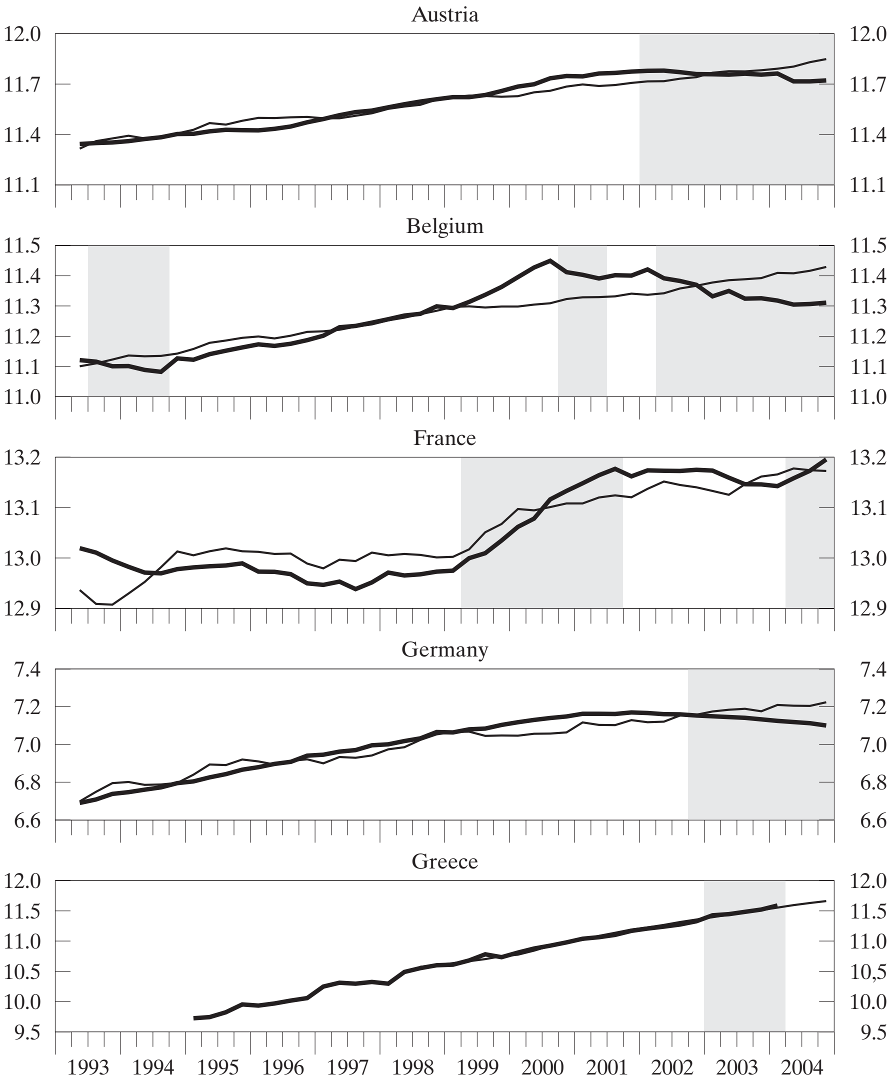
Figure 1

Estimated Long-run Credit Equations, Credit Volume and Regimes
1993 to 2004

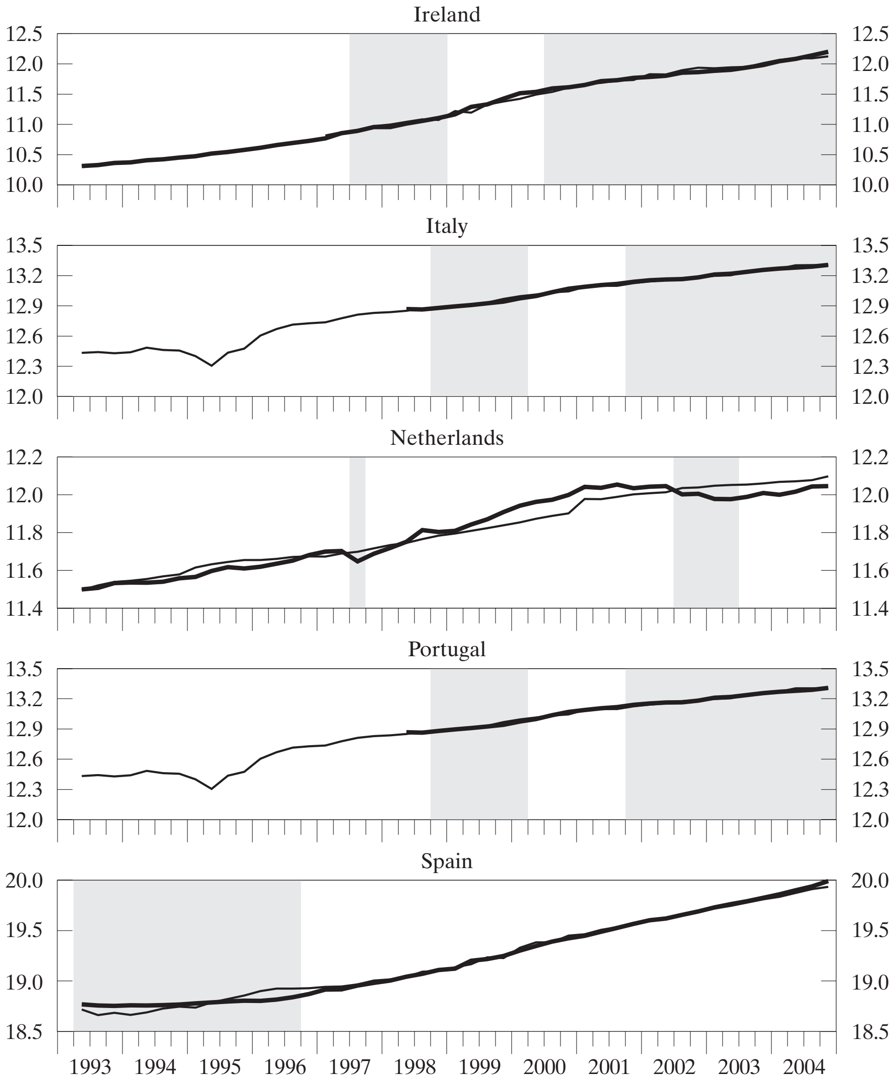


credit volume — estimated long-run relation

periods when the real exchange rate is identified as being in the bubble regime, that is when $P(s_t = 1 \mid \Phi_\tau) > 0.5$

Figure 1

Estimated Long-run Credit Equations, Credit Volume and Regimes, Continued
1993 to 2004



credit volume — estimated long-run relation

periods when the real exchange rate is identified as being in the bubble regime, that is when $P(s_t = 1 \mid \Phi_\tau) > 0.5$

Unit Root test results for time series used in the empirical analysis

Table 4

ADF Unit Root tests

	GDP		Interest Rate		Share prices		Equity Capital	
	Level (c,t)	Change (c)	Level (c,t)	Change (c)	Level (c,t)	Change (c)	Level (c,t)	Change (c)
Austria	–-2.63	–5.44***	–2.99	–4.57***	–1.20	–6.67***	–1.27	–4.03***
Belgium	–2.00	–4.38***	–2.89	–4.31***	–1.36	–7.65***	–1.25	–6.66***
Finland	–2.39	–4.58***	–2.25	–4.31***	–1.04	–6.70***	–1.74	–5.41***
France	–3.14	–9.45***	–3.37*	–4.12***	–1.17	–6.73***	–1.47	–7.71***
Germany	–3.02	–6.33***	–3.22*	–4.44***	–1.43	–6.96***	0.34	–4.20***
Ireland	–0.08	–8.87***	–3.10	–4.95***	–1.53	–7.61***	n.a.	n.a.
Italy	–2.00	–4.38***	–2.16	–3.68**	–1.58	–6.66***	–3.75**	–6.24***
Netherland	–1.31	–6.49***	–3.29*	–4.45***	–1.00	–6.97***	–0.53	–6.07***
Portugal	n.a.	n.a.	–2.69	–4.28***	–1.46	–6.39***	–1.01	–4.53***
Spain	–5.30***	–7.05***	–2.85	–3.44**	–1.36	–7.14***	–2.15	–6.81***

Lag selection by SIC. Lag length are not reported. – ***, **, * indicates significance at the 1, 5 and 10 percent level using Mac Kinnon (1996) one-sided p-values. – c, t: a constant and a trend is included. – c: only a constant is included.

Table 5

ADF Unit Root Tests for Credits

	Sample	Level (c,t)	Change (c)
Austria	93:1 – 04:4	1.26	–4.58***
Belgium	93:1 – 04:4	0.03	–3.14**
Finland	93:1 – 04:4	–2.13	–4.81***
France	93:2 – 04:2	–2.55	–2.29*
Germany	93:1 – 04:4	0.74	0.48
Ireland	93:1 – 04:4	–2.13	–2.99**
Italy	98:2 – 04:4	–0.89	–3.86***
Netherlands	93:1 – 04:4	–0.95	–5.97***
Portugal	93:1 – 04:4	–3.65**	–2.43
Spain	93:1 – 04:4	–4.92***	–3.46**

Zivot Andrews Unit Root test for series with unknown break points for German credit
Modell A

stat.	Break point
–5.0985***	2001:3

The ZA-test is performed using the GAUSS code provided by Junsoo Lee. Zivot Andrews (1992).