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One size does not fit all. A non-linear analysis of European monetary transmission

Giulio Cifarelli° and Giovanna Paladino*

Abstract

This paper investigates the interest rate pass-through in eight European countries analyzing their short-run and long-run monetary transmission mechanisms. We investigate the relationship between the Euribor and the long-run interest rate on loans to non-financial corporations and allow for a mark-up which can be affected by country specific funding conditions and/or stochastic structural breaks. We detect significant differences across countries. Cointegration between the Euribor and the long-term bank loan interest rates holds for Germany, France, and the Netherlands, where banks seem to apply a constant mark-up. In the remaining countries of the sample the long-run pass-through is directly affected by changes in banks' cost of funding, due to shifts in the spread between domestic and German long-term government bond interest rates.

The selection of the country specific ESTAR/LSTAR parameterization of the short-run dynamics detects a high degree of heterogeneity. The transition variables vary from the government bond spreads, in countries which were involved in the European debt crisis via sovereign bond market contagion, to the VXO index and to the Euribor monthly volatility.

Jel Classification E43, E52, F36, C32

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Introduction

The creation of the European Monetary Union has prompted a renewed interest for the pass-through mechanism. A proper understanding of the monetary policy transmission process, i.e. of how changes in the policy rate bring about corresponding shifts in the interest rates set by banks, is of great significance in the case of the Euro Area, where different national banking systems interact with a single common central bank. The recent financial crisis, compounded by the subsequent sovereign debt crisis, has struck the European monetary system in its infancy and has altered the newly established pass-through channels. Our research focuses on the importance of the distortion due to banking fragmentation in the transmission of a common monetary policy.

A burgeoning literature attempts to investigate the working of the pass-through during a crisis, using both linear and non-linear procedures. These studies usually analyze the impact of interest rate volatility changes on the transmission mechanism and identify regime shifts that are associated with them (see, e.g. Humala, 2005, among others).

This paper goes further and directly takes into account the impact of the sovereign debt crisis on the short/long-run monetary policy transmission mechanism by introducing, as transition variables in the non-linear short-run parameterization, various indicators of financial stress which may differ from country to country. We deal with the interest rate setting of national banks which could transfer the common monetary policy across the Euro Area in a heterogeneous way. Indeed, different reactions, both in terms of degree and of speed of adjustment, could distort the desired effect of the monetary stance.

Harmonized monthly data from January 2003 to March 2013 are used to investigate the determinants of the (short-run and long-run) pass-through in eight economies of the EMU: Germany, France, the Netherlands, Italy, Ireland, Greece, Spain and Portugal. The non stationary nature of the data dictates the choice of estimation procedure. The long-run pass-through is quantified by cointegration relationships while the short-run dynamics is modeled with an error correction parameterization. The time interval includes, from October 2008 onwards, an unprecedented financial disruption originating in the US with deep repercussions on the EMU banking sector. As a consequence the empirical investigation has to account for regime shifts and relevant non-linearities.

At first, cointegration long-run pass-through relationships are estimated for each country. Secondly, non-linear ESTAR/LSTAR parameterizations of the short-run pass-through dynamics, conditional on the first step estimates, are investigated. The selection both of the non-linear model and of the corresponding transition variable is data driven and differs across countries.

The paper improves upon previous research in the following respects:

- the long-run pass-through analysis allows for the presence of a time varying mark-up, which takes into account funding problems of national banks in countries in which the financial turmoil was related to the clearance sale of sovereign debt;
- the impact of exogenous factors on the short-run pass-through and on the speed of adjustment is investigated with the help of a highly flexible non-linear approach, which allows for country specific idiosyncrasies both in the dynamic parameterization and in the selection of the transition variables. The latter can be interpreted as latent variables that emerge in periods of turmoil and bring about coefficient shifts if the estimated thresholds are crossed.

Finally, to test the efficacy of a single monetary policy for such a diverse sample of countries, the fitted national loan interest rate shifts are associated to the corresponding industrial production rates of change. It turns out that the heterogeneous behavior of the former is not fully explained by the lack of synchronization of the national business cycles, especially in the second half of the sample. As for the counter-cyclical effort, the ECB does not seem to have been very successful during the hectic years of the financial crisis. This is not to say that the ECB was not reading the economic context correctly but that the majority of the national banking systems were hindering, for various reasons, the proper pass-through of the expansionary monetary policy to the real economy.

Only in the case of Germany, is the simulated change in the loan interest rate negatively correlated with the industrial production rate of change.
The paper is structured as follows: Section 1 discusses the tenets of the pass-through model and of the related literature; Section 2 investigates the properties of the data in a sample period deeply affected by the financial crisis; Section 3 investigates how the monetary policy impinges on bank lending rates; Section 4 focuses on the impact of banks' interest rate settings, associated with a common ECB monetary policy, on industrial output behavior. Section 5 concludes the paper.

1. Bank interest rate pass-through: theory and econometric methodology

1.1 A short survey of the literature

The pass-through model deals with two concurrent transmission mechanisms. The first one refers to the impact of changes in the monetary policy rate across the yield curve, i.e. the term structure of interest rates. The second mechanism addresses the effects of changes in market rates on shifts in bank deposit and lending rates. This is loosely based on the Klein-Monti monopolistic bank model, in which the lending rate is determined as a mark-up on a marginal cost quantified by the market rate at which funds can be obtained.[1]
Most empirical studies focus on the extent to which shifts on monetary interest rates are transmitted to bank lending rates (i.e. the extent of the pass-through), and with the speed of this transmission process.
Two main strands can be distinguished according to the selection of the exogenous market rate. The "cost of funds version" posits a connection between bank lending/deposit rates and market rates of comparable maturity, the latter being selected in order to account for the marginal cost considerations that may affect the price setting behavior of banks (see Sander and Kleimeier, 2004, and De Bondt, 2005, among others). In the "monetary policy version", under the assumption of a stable yield

[1] See Klein (1971) and Monti (1972).

curve, the short-term money market rate is used to gauge the pass-through to bank interest rates (see Belke et al., 2013, and the studies quoted therein).

Several papers analyze this theme, either in a single country context (Coffinet, 2005, and Chionis and Leon, 2006, among others) or in a multiple country one (De Bondt, 2005, Ègert et al., 2006, among many others), typically using single or multiple equation autoregression/error correction models to quantify the short-run dynamics of the interest rates. A stylized finding, common to most studies, is that the pass-through tends to be incomplete as bank interest rates are sticky, in the short-run, and tend to be higher or even complete in the long-run.

Another strand of research focuses on micro bank data and employs panel techniques to examine the interest rate policies of banks in individual countries (see e.g. Weth, 2002, and Gambacorta, 2008). Differences in the degree of pass-through can also be attributed to factors - discussed in the analysis of banks' interest margins by Ho and Saunders (1981) and Maudos and de Guevara (2004) - that are mostly related to idiosyncratic characteristics of the national banking systems. Among these, bank competition and ownership, rigidities and size of the operating costs, dimension and rating of the public debt, and finally differences in supervisory approaches as supervisory fragmentation may affect the transmission via the credit channel (ECB 2012, chart 34, p. 50).

The literature also spans different time periods and data sources. Studies based on data that precede the introduction of the Euro, as Mojon (2000), tend to conclude that the degree and speed of pass-through differ considerably across countries, especially in the short-run, and find scattered evidence of a full pass-through in the long-run. Sørensen and Werner (2006), among others, detect a partial reduction in pass-through heterogeneity in the first years of the Euro. The process of convergence, as pointed out by Blot and Labondance (2011) and Illes and Lombardi (2013), was interrupted by the financial crisis.

This crisis has been characterized by bouts of severe financial turbulence and by events of liquidity shortage which introduce non-linearities in the data. They have to be accounted for in the parameterization of the pass-through dynamics to avoid model misspecification. Some studies rely on a split sample estimation which is assumed to coincide with the breakpoints of the interest rate time series under investigation (see Panagopoulos and Spiliotis, 2012). More recently, Aristei and Gallo (2014) follow Humala (2005), and analyze the stochastic nature of the pass-through regime shifts brought about by the Euro Area financial crisis with the help of the two-stage Markov-switching VECM procedure of Krolzig (1997).

1.2 A model of non linear pass-through.

The price setting reaction of banks reads as follows

$$lr_t = \beta_0 + \beta_1 eur_t + \beta_2 x_t + \varepsilon_t \tag{1}$$

lr_t and eur_t are, respectively, a bank lending rate and a monetary policy rate. x_t is an additional factor which may impact on the bank retail rates, depending upon the economic context. β_0 provides a measure of a constant bank mark-up and β_1 quantifies the pass-through. It is assumed that $0 \leq \beta_1 \leq 1$, the pass-through being complete only if $\beta_1 = 1$. If $\beta_1 < 1$, banks do not transfer the entire market interest rate shift on their retail rates for various reasons, which range from market power, switching costs and loan/deposit markets liquidity. It is assumed that $\beta_0 > 0$, $\beta_1 > 0$, whilst no a priori assumption is made on the sign of β_2.

If, as is the case in this paper, the variables entering equation (1) are non stationary and cointegrated, equation (1) measures the long-run pass-through. The short-run dynamics about the long-run equilibrium can be modeled by the following error correction relationship

$$\Delta lr_t = b_0 + \sum_{j=1}^{n} a_j \Delta lr_{t-j} + \sum_{i=1}^{n} b_i \Delta eur_{t-i} + \sum_{l=1}^{n} d_l \Delta x_{t-l} + \alpha \varepsilon_{t-1} + v_t \tag{2}$$

where ε_{t-1} is the lagged residual of the cointegration relationship (1), the b_i coefficients quantify the short-run pass-through dynamics and α, the error correction term, measures the speed of adjustment of banking rates to the long-run equilibrium.

The pass-through dynamics, in order to avoid substantial model misspecification, has to account for non-linearities in the data. We model stochastic regime switches using the LSTAR/ESTAR procedure of Granger and Teräsvirta (1993). The error correction equation used to parameterize the non-linear short-run pass-through reads as follows[2]

$$\Delta lr_t = (b_{10} + \sum_{j=1}^{n} a_{1j}\Delta lr_{t-j} + \sum_{i=1}^{n} b_{1i}\Delta eur_{t-i} + \sum_{l=1}^{n} d_{1l}\Delta x_{t-l} + \alpha_1\varepsilon_{t-1})(1 - S_t) +$$

$$\left(b_{20} + \sum_{j=1}^{n} a_{2j}\Delta lr_{t-j} + \sum_{i=1}^{n} b_{2i}\Delta eur_{t-i} + \sum_{l=1}^{n} d_{2l}\Delta x_{t-l} + \alpha_2\varepsilon_{t-1}\right)S_t + u_t \qquad (3)$$

S_t (S_t^*) is the smooth transition function, which drives the motion from one regime to the other.
It is assumed that

$$S_t = \frac{1}{1+e^{-\gamma(z_{t-d}-c)}} \qquad (4)$$

in the case of a LSTAR model and

$$S_t^* = 1 - e^{(-\gamma(z_{t-d}-c)^2)} \qquad (4')$$

in the case of an ESTAR parameterization. z_t is the transition variable associated with the regime switching process, c is the estimated threshold value, and γ denotes the speed and the smoothness of the transition from one regime to the other.[3] The delay, d, with which z_t's crossing of the threshold brings about a regime shift is assessed empirically and depends upon the structural characteristics of the national banking systems. At any point in time, the dependent variable is generated by a combination of both regimes.

[2] The selection of this parameterization is justified, according to Teräsvirta (1994), by the plurality of the economic agents that are involved in the decision process. Even if a single investor takes a dichotomous decision, it is unlikely that all agents act simultaneously. Since the interest rate time series provide information on the aggregate decision process only, the regime shift will be smooth rather than discrete.
[3] The larger is γ the stronger is the synchronization of banks' reaction to Euribor changes.

In the case of the LSTAR model, if z_{t-d} is large and positive, S_t will tend to 1 and the weighted coefficients will be close in value to those of regime 2; if z_{t-d} is large and negative, S_t will tend to 0 and the weighted coefficients will be similar to the coefficients of regime 1. If $z_{t-d} = c$, $S_t = 0.5$ and the weighted coefficients will be the arithmetic average of the coefficients of both regimes. In the case of the ESTAR model if $z_{t-d} = c$, $S_t^* = 0$ and the weighted coefficients will be equal to the coefficients of regime 1, whereas, for values of z_{t-d} that differ from c, $0 \leq S_t^* \leq 1$ and the weighted coefficients will be a combination of the coefficients of both regimes, the weight of the coefficients of regime 2 growing will the increase in the divergence between z_{t-d} and c.

In the stochastic regime shift studies quoted in the previous section, the regimes are typically associated with differing levels of interest rate volatility, which are assumed to model differing pass-through policies in normal versus turbulent market contexts. The approach introduced in this paper is more general and investigates alternative potential sources of regime shifts during the recent financial crisis. The behavior of banks, in countries where financial distress is linked to the Lehman Bank collapse, differs from the reaction of banks in the periphery countries, which are severely affected by the sovereign bond sell offs initiated in Greece.

We have tested, for each national banking system, alternative potential transition variables. z_{t-d} ranges from the Euribor monthly volatility, to the change in the VXO index, and to the change in the spread between the national long-term government bond and the corresponding German bond. An additional dimension is added in this way to the identification process of the stochastic regime shifts of the short-term pass-through.

2. Preliminary statistical analysis

2.1 Monetary policy in Europe across a stormy decade.

Figure 1 shows monthly interest rates set by banks in Germany, France, Greece, Ireland, Italy, the Netherlands, Spain and Portugal for new loans to non-financial firms with maturity of over 5 years plus the 3 month

Euribor. A graphical inspection reveals at least seven critical points in the trend of these time series.

[Insert figure 1 here]

The first break, in 2003, occurs when the European Central Bank cut official interest rates by 0.25% in March and by 0.5% in June. As a result the minimum bid rate on main refinancing operations tumbled to 2.0%. These measures were taken in a context characterized by low inflationary pressures, productivity stagnation, uncertain prospects for recovery, and rising international political tensions due to the war in Iraq and to terrorist acts in Europe and the Middle East.

The second critical point takes place at the end of 2005 and early 2006, when, after a period of substantial stability, interest rates start to rise. Actually, the ECB kept official interest rates unchanged for almost two years, in a context of uncertainty about the strength of the economic recovery and stable inflationary expectations. From Autumn 2005 onwards, a resumption in growth and higher oil prices brought about an increase both in consumer prices and in medium term inflation expectations. As a result, the ECB raised the official interest rates by 0.25% in December 2005, and again in March 2006.

The two subsequent years witnessed continuous increases in official interest rates and, consequently, in interbank rates.

The third point is in September 2007. It corresponds to the spreading of the US sub-prime mortgage crisis in the European financial system and to the related increase in money market tensions. The European Central Bank implemented unconventional monetary policy measures to overcome the mutual lack of trust on the interbank market.

The fourth critical point is at the end of 2008, when the international financial crisis triggered by the Lehman bankruptcy forced the ECB to persist in an expansive monetary action. The uncertainty about possible defaults of counterparts impaired the correct functioning of the wholesale markets on which banks do their fundraising. Central banks tried to compensate the paralysis of national interbank markets with substantial liquidity injections. On 8 October 2008, the ECB, the Federal Reserve, the Bank of England, the Bank of Canada, the Bank of Sweden and the Swiss National Bank, with the support of the Bank of Japan, carried out a coordinated reduction in interest rates. Further cuts also occurred in the following months when it became clear that the Euro Area was in recession. The ECB continued to cut interest rates until May 2009. In June of the same year, it launched a first covered bond program with the aims of encouraging banks to maintain and expand their lending to clients; improving market liquidity in important segments of the private debt security market; and easing funding conditions for banks and enterprises. In April 2010, Greece sought financial support and in May the ECB reacted with a new program, aiming to ensure depth and liquidity in those market segments which were dysfunctional. The objective of this program was to address the malfunctioning of securities markets and restore an appropriate monetary policy transmission mechanism. On 21 November 2010, Ireland sought financial support too.

The fifth break-point highlights the change in monetary policy towards a more restrictive stance, introduced on 7 April 2011, despite the request of aid made the day before by Portugal. After two years of historically low levels, the Governing Council decided a 0.25% increase on the basis of an upside risk to price stability and in order to keep inflationary expectations anchored at 2%. A further increase followed in July, fully compensated, however, by a corresponding cut in November: a clear sign of the difficulties encountered by the Central Bank in reading the economic picture.

In the sixth phase, additional key interest rate cuts and other monetary expansion measures were adopted to cope with the destabilizing effects of the sovereign debt crises.

2.2 The data

This study uses monthly purely harmonized data from the ECB MFI interest rate data base for eight EMU countries: Germany, France, Greece, Ireland, Italy, the Netherlands, Spain, Portugal. Bank interest rates are applied to new loans to non-financial corporations with maturity over 5 years. Long-term (10 years) government bond interest rates are used to compute spreads between national and German rates. The monthly VXO is the S&P100 implied volatility index and comes from the CBOE whereas the industrial production index is taken from the OECD data base.

The analysis ranges from January 2003, which corresponds to the beginning of the MFI interest rate statistics, to March 2013. We use the 3 month Euribor as money market rate and as proxy of the stance of monetary policy. The official interest rate cannot be used directly, since the ECB interest rate on refinancing operation (MRO) changes only when the ECB modifies it.

In the literature some researchers use the Eonia, as they assume that it reflects relatively well official rate decisions, is closer to the MRO, and is less related to liquidity issues. Our decision to use the Euribor is due to the fact that it influences the monetary policy transmission by incorporating the expectation of short-run interest rate changes (Abbassi and Linzert, 2012). Moreover, the Euribor is used as the basis of most of the floating rate loans and represents a good indicator of the cost of money for the real economy. This rate is also important since, in theory, it measures the cost of interbank funding and depends on the expectation of bank solvency.

Both the interest rate and spread time series are I(1). Since the graphs in figure 1 suggest that the time series may be affected by structural breaks, with subsequent reduction in the power of the standard unit root tests, we use the Lee and Strazicich (2003, 2004) procedure.[4] The test statistics are set out in table A.2 of the appendix. The monthly changes in the Euribor and in the bank loan interest rates (see table A.1 of the appendix) are always serially correlated and, as expected in such a turbulent time period, conditionally heteroskedastic and non-normally distributed.

[4] The Lee and Strazicich Lagrange multiplier unit root test allows for one or two breaks in level and/or trend under both the null and alternative hypotheses, and rejection of the null implies unambiguous trend stationarity.

3. Non-linear pass-through model estimates

To answer our research question we consider the relationship between the policy interest rate and the interest rate set by banks on loans to non-financial corporations. We implement, for each country, a two stage estimation procedure which builds on Krolzig (1997).[5]

At first, the long-run co-movement of the time series is investigated, using the maximum likelihood cointegration analysis of Johansen (1995). Secondly, conditional on the presence of cointegration, and on the estimates of a cointegration vector, the non-linearities of the short-run dynamics of the corresponding error correction process are parameterized as LSTAR/ESTAR models and estimated using the complex maximum likelihood procedure set out by Teräsvirta (1994).

3.1 Cointegration analysis and the long-run pass-through

The Johansen trace statistics detect the presence of cointegration between the 3 month Euribor and the non-financial firms loan rate in Germany, France and the Netherlands (see table 1, panel A).

[Insert table 1 here]

In the case of Greece, Italy and Portugal cointegration holds when to the loan rate and to the Euribor we add the spread between the domestic and German long-term bonds plus a dummy, which accounts for a structural break in the constant term, in the case of Spain and Ireland (see table 1, panel B).

The estimates of the long-run relationship set out in table 2 are surprisingly heterogeneous across countries and conducive to the

[5] In the two-stage maximum likelihood approach of Krolzig (1997) a first stage Johansen cointegration analysis is followed by a second stage Markov regime switching investigation of the short-run dynamics (see Clarida et al. (2006) for further details).

hypothesis of a time varying mark-up due to changes in market conditions during the financial crisis. The estimated cointegration equation reads as

$$lr_t - \beta_0 - \beta_1 eur_t - \beta_2 spr_t - \beta_3 dum_t = \varepsilon_t \tag{5}$$

where $dum_t = \begin{cases} 0, \ if \ t \leq 2008{:}09 \\ 1, if \ t > 2008{:}09 \end{cases}$

The empirical findings are satisfactory and in line with the specification of the model. The long-run pass-through seems to be complete in the case of Germany only. In the case of France, the Netherlands and Ireland, the long-run pass-through declines, a symptom of the autonomy of national banks, which are able to exert some market power. Banks raise their lending rates, whenever the spread between the national government bonds and the corresponding German bonds rises, in Spain, Greece, Italy, Ireland, and Portugal, countries affected by the sovereign debt crisis via contagion. A decline in the price of the government bonds, i.e. an increase in the spread, reduces the value of bank portfolios, lowers their creditworthiness, and raises their cost of funding which they pass-through to the borrowers by raising the lending rates.

Small (and not significantly different from zero) in Germany, the constant component of the mark-up, β_0, is larger in the remaining countries of the sample and ranges from 1.287 in Spain to 3.461 in Greece. The dummy variable coefficients suggest, moreover, that it shifted because of the financial crisis; Spanish banks were raising their constant mark-up and Irish banks were reducing it.

In order to assess whether the estimates are stable over time, we have performed the recursive $supQ_T^{(t)}$ test of parameter constancy of Hansen and Johansen (1999) and have found that, with the exception of Ireland, the null of parameter constancy cannot be rejected at the 5% level of significance.[6] This finding supports the use of a two stage approach, a

[6] $supQ_T^{(t)}$ is the *supremum* of the Nyblom statistic $Q_T^{(t)}$ defined in Hansen and Johansen (1999, p. 315, eq. 20).

stable long-run pass-through relationship being associated with a stochastic regime switching parameterization of the short-run dynamics.

[Insert table 2 here]

3.2 ESTAR/LSTAR parameterization of the short-run pass-through

Given the unprecedented turbulence of the time period under examination, the question on non-linearities rises spontaneously to researchers. Conditional on the cointegration analysis set out in the previous section, a test of linearity against the non-linear parameterization of the error correction relationship, equation (3), is then performed adopting the procedure of Luukkonen et al. (1988).

As a first step, we select the transition variable, z_t, among the following set of candidates: Δlr_t, monthly change in the long-run lending rate to non-financial corporations, Δspr_t, monthly variation in the spread between the 10 years domestic goverment bond interest rate and the 10 years Bund interest rate, $eursd_t$, the monthly average of the daily standard deviation of the 3 month Euribor change, and Δvxo_t, the month to month shift in the monthly VXO index divided by 100. As a second step, the transition function S_t is replaced in the error correction relationship (3), by third order Taylor series approximations, and after some algebraic manipulation, the following auxiliary equation is estimated, where $x_t = spr_t$, in the case of Spain, Greece, Italy, Ireland, and Portugal, and $x_t = 0_t$ in the remaining countries of the sample.[7]

$$\Delta lr_t = (\delta_0 + \sum_{j=1}^{n} \rho_{0j} \Delta lr_{t-j} + \sum_{i=1}^{n} \theta_{0i} \Delta eur_{t-i} + \sum_{l=1}^{n} \vartheta_{0l} \Delta x_{t-l} + \varphi_0 \varepsilon_{t-1}) +$$

$$(\delta_1 z_{t-d} + \sum_{j=1}^{n} \rho_{1j} \Delta lr_{t-j} z_{t-d} + \sum_{i=1}^{n} \theta_{1i} \Delta eur_{t-i} z_{t-d} + \sum_{l=1}^{n} \vartheta_{1l} \Delta x_{t-l}\, z_{t-d} + \varphi_1 \varepsilon_{t-1} z_{t-d}) +$$

[7] The number of lags n in the non-linear error correction model estimated in this section coincides in the case of Germany, France, Spain, Italy and Ireland with the number of lags used in the Johansen cointegration analysis and estimation of the long-run pass-through. In the remaining countries it is arbitrarily set to 3.

$$(\delta_2 z_{t-d}^2 + \sum_{j=1}^{n} \rho_{2j} \Delta lr_{t-j} z_{t-d}^2 + \sum_{i=1}^{n} \theta_{2i} \Delta eur_{t-i} z_{t-d}^2 + \sum_{l=1}^{n} \vartheta_{2l} \Delta x_{t-l}\, z_{t-d}^2 + \varphi_2 \varepsilon_{t-1} z_{t-d}^2) + \quad (6)$$

$$(\delta_3 z_{t-d}^3 + \sum_{j=1}^{n} \rho_{3j} \Delta lr_{t-j} z_{t-d}^3 + \sum_{i=1}^{n} \theta_{3i} \Delta eur_{t-i} z_{t-d}^3 + \sum_{l=1}^{n} \vartheta_{3l} \Delta x_{t-l}\, z_{t-d}^3 + \varphi_3 \varepsilon_{t-1} z_{t-d}^3) + e_t$$

We test linearity against LSTAR modeling – for various values of the delay parameter d – performing F tests of the null hypothesis

$$H_0:\ \delta_1 = \delta_2 = \delta_3 = \rho_{11} = \rho_{21} = \rho_{31} = \cdots = \rho_{1n} = \rho_{2n} = \rho_{3n} = \theta_{11} = \theta_{21} = \theta_{31} = \cdots = \theta_{1n} = \theta_{2n} = \theta_{3n} = \vartheta_{11} = \vartheta_{21} = \vartheta_{31} = \cdots = \vartheta_{1n} = \vartheta_{2n} = \vartheta_{3n} = \varphi_1 = \varphi_2 = \varphi_3 = 0$$

As can be seen from the statistics set out in the LM-NLT row of table 3, the null is uniformly rejected which justifies the non-linear parameterizations of the short-run pass-through.[8]

[Insert table 3 here]

In table 3 are set out, for the sake of parsimony, besides the logistic or exponential function parameters, only the short-run pass-through and the error correction coefficients, over the two regimes. The quality of the estimates is here highly satisfactory. The residuals are serially uncorrelated and conditionally homoscedastic, a relevant result given the turbulence of the sample period. In the same way, the Jarque Bera test statistics suggest that the model has captured most of the non-normality of the data, originally detected in table A.1. As a final check of the validity of the specification of our model, we performed likelihood ratio tests of the hypothesis that the non-linear coefficients of the equation (3) estimates are nil. As can be seen from the LR(NL) statistics, set out in the last row of the table, the null is strongly and uniformly rejected.

[8] Following Escribano and Jordá (1999), we perform a test of linearity against an ESTAR non-linear parameterization by adding an additional component, corresponding to the fourth power of the transition variable z_{t-d}^4, in the linear auxiliary regression, and testing the null hypothesis that all the coefficients corresponding to the Taylor approximation of the non-linear function are nil. In the case of Portugal the null is rejected, which leads to the selection of an ESTAR model.

The short-run pass-through parameters change over the two regimes and interact with the shifts in the error correction coefficient. Coefficient γ estimates are always rather large. The speed of the transition from one regime to the other is relatively high, which denotes a synchronized reaction of banks to shifts in the transition variable z_t. In the case of Germany and Ireland the speed of adjustment to deviations from the long-run pass-through α rises in regime 2, when uncertainty, measured by increases either in $eursd_t$, or in Δvxo_t, rises above a given threshold. In the estimates of Greece, epicenter of the crisis, the transition variable is the (lagged) change in the loan rate to non-financial corporations Δlr_t. The speed of adjustment accelerates after a sharp drop in Δlr_t, when the transition variable lies below a negative threshold, possibly in a period of turmoil, associated here with regime 1.

In countries that were involved in the sovereign debt crisis, we find an opposite reaction, due probably to the greater intensity of the banks' financial distress. With values of the spread change, Δspr_t, exceeding a positive threshold, the (regime 2) error correction process either becomes nil (Italy and the Netherlands), or fails to converge (Spain), since the error correction coefficient becomes positive. A stylized finding, common to these country estimates, is that the absolute value of the short-run pass-through coefficients rises in the turbulent regime, which denotes a greater responsiveness (or even overreaction) of banks to Euribor shifts.[9]

The analysis of the French estimates identifies a regime shift from 2003 to 2006, which brings about a decline in the speed of adjustment to deviations from the long-run relation and a decline in the absolute value of the short-run pass-through. The short-run non-linear dynamics, in the case of Portugal, are parameterized by an ESTAR, which is less informative. It turns out that as the spread change (Δspr_t) deviates from an equilibrium threshold value, a symptom of public finance stress, the speed of adjustment to the long-run relation increases. No clear cut change can be detected in the other components of the short-run pass-through parameterization.

[9] Indeed, in the case of Greece and Ireland some of the pass-through coefficients in the distress regime are larger than one in absolute value.

[Insert figure 2 here]

[Insert figure 3 here]

[Insert figure 4 here]

[Insert figure 5 here]

Figures 2 to 5 provide valuable information on the timing of the regime shifts. In them are to be found, besides the month to month changes in the Euribor and in the fitted bank interest rate on loans to non-financial corporations, the evolution over time of logistic or exponential transition functions, and of the transition variables. The graphs corroborate the hypotheses set out in our model, viz. that banks react to an increase in financial turmoil, identified here by the crossing of a threshold by the transition variable, and switch to a stress regime, proxied in all countries (with the exception of Greece, where the opposite is the case) by regime 2.

Distinct national patterns are identified. In Germany a shift to regime 2 is brought about by the Lehman collapse of 2008, a common feature to all the banking systems of the sample, with the notable exception of France. No regime shift occurs afterwards. In the remaining countries, the renewed financial fragility, due to the deterioration of the public finances, affects the banking system and brings about multiple regimes shifts from 2010 onwards. Bank switches to the stress regime correspond to large increases in the interest rate spread between the domestic and German government bonds in Italy, Spain, Portugal and the Netherlands, in the VXO index in Ireland and in its own loan interest rate in Greece.

The interest rate setting behavior of French banks is not affected, according to our model, by the recent financial crisis. Banks have slowly adapted to the Euro from 2003 to 2006 and no regime shift is detected afterwards.

It should finally be noticed that, with the exception of the Lehman collapse, (which is followed by a large decline in the Euribor and by concomitant regime shifts in the national banking systems), during this long financial crisis domestic banking systems react, by a shift in regime, to stress factors that are not under the ECB's control. This compounds the difficulty of providing a common monetary policy which is adapted to the requirements of each country.

4. The effectiveness of the pass-through on the real economy

In this section, we try to provide an answer to the question of whether the changes in the long-term interest rate on loans to non-financial firms, $\widehat{\Delta lr_t}$ estimated according to equation (3), have a relevant impact on the real business cycle.[10] The assumption is that monetary policy, especially after the crisis, is used for counter-cyclical purposes. Thus the relation should have a negative sign since a negative interest rate shift can be associated with a positive shift in real industrial output rate of growth.

We first consider a simple correlation matrix between $\widehat{\Delta lr_t}$ and $\Delta lind_t$, the monthly rate of growth of industrial output. The matrix is computed over the full sample as our estimation accounts for regime changes that could determine changes of weights in the target function between inflation and capacity output.

Our *a priori* is that if a national banking system filters the monetary policy, adjusting it to the national real cycle and adapting in this way the common monetary policy to idiosyncratic domestic conditions, a negative correlation between $\widehat{\Delta lr_t}$ and $\Delta lind_t$ is to be expected. On the contrary, if the national banking system hinders the transfer of monetary impulses from the ECB to the real sector for various exogenous reasons, such as high risk profile of firms due to excessive leverage and/or tougher national supervision (which may discourage bank lending), the correlation should be positive or null.

[Insert table 4 here]

From table 4 we detect a significant real effect of the pass-through only in the case of Germany and France (where, however the correlation is

[10] Equation (3) can be interpreted as a reaction function of the national banking system to the common monetary policy.

negative and significant only in the case $\Delta lind_{t+6}$). The table reports both the results of Spearman rank-order, which accounts for large outliers, and standard Pearson correlations. These findings are corroborated by the variance decomposition in figure 6, obtained from the estimates of eight $\Delta lind_t$, $\widehat{\Delta lr}_t$ bivariate VAR systems, one for each country.[11] The variance of the $\Delta lind_t$ time series, which is explained by the estimated changes of the long-term interest rate on loans to non-financial firms $\widehat{\Delta lr}_t$, is different from zero only for Germany.[12]

[Insert figure 6 here]

5. Conclusion

This paper investigates the pass–through from the money market to the bank long-term interest rate on loans to non-financial corporations for eight EMU countries over the 2003-2013 time period. We find significant differences between the short and the long-run pass-through specifications of the national banking systems.

A positive relationship between the interest rate on loans to non-financial corporations and the spread between domestic and German government bond interest rates is found only in the long-run pass-through of countries that are involved in the European sovereign debt crisis. Indeed, shifts in the spread bring about corresponding changes in banks' cost of funding.

[11] The lag order of the $\widehat{\Delta lr_t}$, $\Delta lind_t$ bivariate VAR systems used to compute the variance decompositions is selected with the help of Wald lag exclusion tests. They suggest the following orders: 6 lags for Germany, 3 lags for France, the Netherlands, Spain, and Portugal, 4 lags for Ireland and Italy, and 2 lags for Greece.

[12] At this point we should dispel the doubt that the ECB is acting only in response to the German real cycle. A correlation analysis among the $\Delta lind_t$ time series shows that, in many cases, the contemporaneous correlation between the domestic cycle and the German one is positive and significant (a result which also holds using the Spearman index). This corroborates the idea that a significant part of the malfunctioning of the monetary policy transmission, with respect to the counter-cyclical target, is not due to asymmetries in the real cycle but to asymmetries in the reaction function of national banking systems. For the sake of space the correlation table is not reported here but it is available upon request to the corresponding author.

The non-linear short-run processes too are heterogeneous across countries. Various transition variables generate mostly non synchronized regime shifts. In particular, the spread between the domestic and German government bond interest rates plays a crucial (and potentially destabilizing) role in the short-run transmission mechanism in Italy, Spain, Portugal, and the Netherlands. Thus a preemptive management of the spread can be seen as kind of auxiliary monetary policy tool. Its relevance cannot be underestimated.

Our findings identify, both in the short and in the long-run monetary policy transmission, a source of heterogeneity which is largely independent from both central bank decisions and national bank's behavior. This raises the difficulty of implementing a common monetary policy in a fragmented European banking system where "one size cannot fit all".

Table 1. Johansen cointegration Trace test statistics

Panel A.

List of variables in the cointegration vector: lr_t, eur_r					
	Hypothesized N. of Cointegration Relationships	**Trace Test Statistic**	**Probability of Rejection of the Null Hypothesis**	**N. of lags in VAR**	**Deterministic Trend Assumption**
BD	None at most 1	27.388* 1.124	[0.004] [0.920]	3	Restricted constant
FR	None at most 1	25.003* 1.449	[0.009] [0.871]	2	Restricted constant
NL	None at most 1	24.588* 1.706	[0.010] [0.827]	5	Restricted constant

Panel B.

List of variables in the cointegration vector: lr_t, eur_r, spr_t, dum_t					
	Hypothesized N. of Cointegration Relationships	**Trace Test Statistic**	**Probability of Rejection of the Null Hypothesis**	**N. of lags in VAR**	**Deterministic Trend Assumption**
ES	None at most 1 at most 2	75.115* 20.210 3.727	[0.000] [0.243] [0.788]	4	Restricted constant
GR	None at most 1 at most 2	36.553* 16.286 1.803	[0.034] [0.164] [0.810]	9	Restricted constant
IT	None at most 1 at most 2	35.887* 10.228 1.904	[0.004] [0.623] [0.792]	4	Restricted constant
IR	None at most 1 at most 2	49.186* 16.705 6.812	[0.013] [0.464] [0.397]	4	Restricted constant
PT	None at most 1 at most 2	39.938* 17.324 3.796	[0.013] [0.122] [0.455]	7	Restricted constant

Notes. *: significant at the 5% level; lr_t: long-run lending rate to non-financial corporations; eur_r: 3 month Euribor; spr_t: spread between the 10 years domestic government bond interest rate and the 10 years Bund interest rate; $dum_t = \begin{cases} 0, & if\ t \leq 2008{:}09 \\ 1, & if\ t > 2008{:}09 \end{cases}$. We use the country acronyms of the ECB.

Table 2. Cointegration equation estimates

$$lr_t - \beta_0 - \beta_1 eur_t - \beta_2 spr_t - \beta_3 dum_t = \varepsilon_t \quad (5)$$

	β_0	β_1	β_2	β_3	$supQ_T^{(t)}$ date
BD	-0.446 (-0.626)	-1.024* (-4.379)			1.589 2008:11
FR	-2.864* (-19.171)	-0.550* (-9.719)			1.966 2007:02
NL	-3.064* (-19.171)	-0.439* (-8.111)			1.588 2008:10
ES	-1.287* (-18.700)	-0.907* (-41.099)	-0.207* (-10.223)	-0.674* (-11.686)	1.927 2012:01
GR	-3.461* (-17.130)	-0.623* (-9.552)	-0.088* (-6.661)		1.435 2012:08
IT	-2.164* (-34.881)	-0.787* (-41.462)	-0.062* (-3.343)		2.071 2009:03
IR	-3.268* (-21.273)	-0.531* (-11.561)	-0.098* (-3.086)	1.05* (6.138)	6.509* 2008:09
PT	-2.557* (-30.785)	-0.711* (-26.439)	-0.093* (-8.351)		0.701 2010:03

Notes. t ratios are in parentheses; *: significant at the 5% level; $supQ_T^{(t)}$: Hansen and Johansen (1999) parameter constancy test; $dum_t = \begin{cases} 0, & if\ t \leq 2008{:}09 \\ 1, & if\ t > 2008{:}09 \end{cases}$.

Table 3. Short-run pass-through LSTAR/ESTAR estimates

	BD	FR	NL	ES	GR	IT	IR	PT
b_{11}	0.014 (0.50)	-0.256* (-2.15)	0.139* (4.53)	0.125* (2.27)	-1.004* (-3.06)	0.117* (2.018)	0.495* (4.28)	0.715* (4.25)
b_{12}	0.084* (3.53)		0.030 (0.92)	0.081* (1.75)	1.312* (4.73)	0.070 (1.36)	-0.016 (-0.12)	-0.291 (-1.23)
b_{13}			0.022 (0.59)	0.039 (1.01)	-0.307 (-1.58)	-0.033 (-0.52)	0.047 (0.34)	-0.138* (-1.93)
b_{21}	0.451* (5.95)	0.126* (2.56)	0.327* (2.50)	0.415* (3.21)	0.308* (7.191)	0.638* (5.84)	-0.390 (-1.06)	0.259* (5.56)
b_{22}	0.650* (3.07)		-0.299 (-1.087)	-0.228* (-2.09)	0.016 (0.31)	0.508* (3.85)	-1.199* (-6.10)	0.065 (1.05)
b_{23}			0.728* (3.86)	0.347* (3.22)	0.158* (3.04)	0.677* (2.84)	1.358* (5.04)	0.150* (2.74)
α_1	-0.006* (-3.63)	-0.673* (-4.39)	-0.024* (-3.03)	-0.173* (-4.23)	-0.442* (-2.99)	-0.179* (-4.20)	-0.095* (-2.03)	-0.011 (-0.11)
α_2	-0.397* (-2.04)	-0.140* (-6.00)	-0.029 (-0.68)	0.225* (2.93)	-0.056* (-2.12)	0.036 (0.26)	-0.134* (-4.11)	-0.084* (-2.73)
γ	76.433* (3.30)	25.735* (2.29)	104.754* (1.84)	43.593* (3.08)	29.587* (2.84)	70.485 (1.61)	44.725* (3.32)	93.332* (3.01)
c	0.129* (12.80)	0.187* (4.76)	0.060* (8.23)	0.098* (8.04)	-0.161* (-6.051)	0.166* (8.97)	0.039* (3.25)	0.156* (16.84)
Transition variable	$eursd_{t-1}$	*Time*	Δspr_{t-1}	Δspr_{t-2}	Δlr_{t-3}	Δspr_{t-1}	Δvxo_{t-1}	Δspr_{t-4}
AR(1)	2.871 [0.090]	1.523 [0.217]	1.230 [0.267]	0.001 [0.956]	0.126 [0.723]	1.352 [0.245]	0.414 [0.520]	0.424 [0.515]
AR(5)	23.747* [0.000]	6.642 [0.249]	3.020 [0.697]	2.003 [0.848]	4.787 [0.442]	12.959* [0.023]	1.595 [0.902]	5.708 0.356
HET(1)	0.417 [0.518]	1.432 [0.231]	3.360 [0.067]	0.000 [0.998]	0.436 [0.509]	5.464* [0.019]	0.076 [0.783]	0.041 [0.839]
HET(5)	0.846 [0.655]	4.887 [0.430]	7.055 [0.216]	9.717 [0.084]	2.226 [0.817]	10.418 [0.064]	0.244 [0.998]	3.303 [0.654]
LLF	313.165	144.006	291.491	242.973	170.235	223.818	160.181	231.234
JB	27.888 [0.000]	8.895 [0.012]	9.686 [0.008]	10.645 [0.005]	174.831 [0.000]	48.054 [0.000]	58.464 [0.000]	52.052 [0.000]
LM-NLT	5.593 [0.000]	1.790 [0.078]	4.121 [0.000]	4.252 [0.000]	1.830 [0.017]	2.024 [0.007]	1.901 [0.012]	2.849 [0.000]
LR(NL)	62.935 [0.000]	17.341 [0.008]	67.845 [0.000]	76.759 [0.000]	42.870 [0.000]	50.5492 [0.000]	39.809 [0.000]	39.549 [0.000]

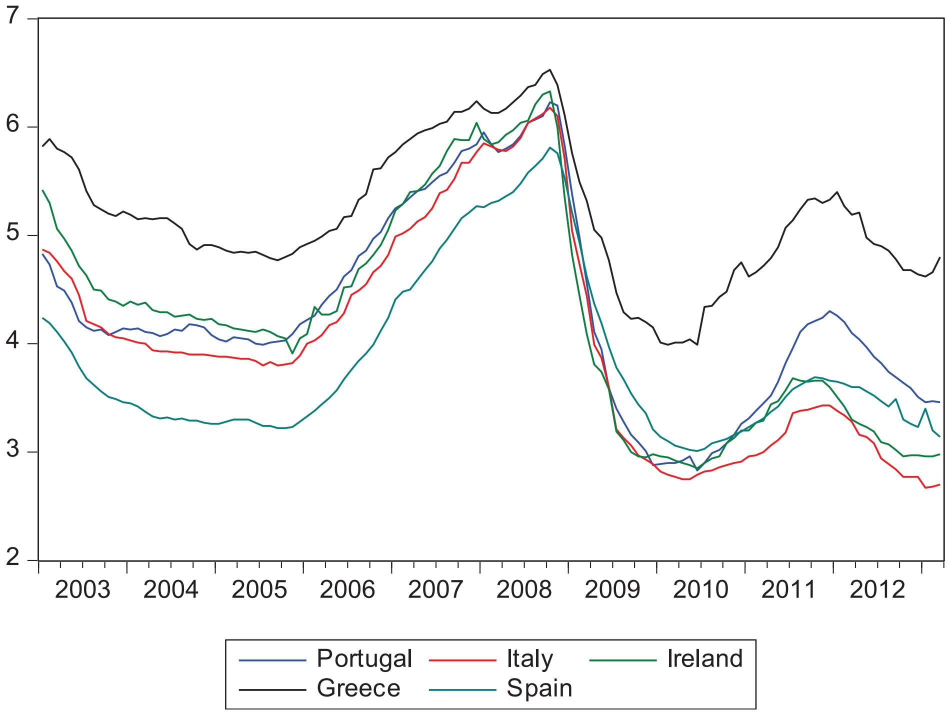
Notes. t-ratios in parentheses and probability values in square brackets; *: significant at the 5% level; LLF: Log Likelihood value; JB: Jarque-Bera normality test; AR(n): Ljung-Box test statistic for n-th order serial correlation; HET(n): Ljung-Box test statistic for n-th order serial correlation of the squared time series; LM-NLT: LM non-linearity test; LR(NL): Likelihood ratio test of the null of non-linearity in the estimates H_0, where H_0: $b_{20}=a_{21}=\ldots=a_{2n}=b_{21}=\ldots=b_{2n}=d_{21}=\ldots=d_{2n}=\alpha_2=\gamma=c=0$.

Table 4. Correlation between the fitted long-term interest rate on loans to non-financial firms $\widehat{lr_t}$ and the rate of change of industrial output $\Delta lind_t$

	Spearman				Pearson			
	$\Delta lind_{t+3}$	$\Delta lind_{t+6}$	$\Delta lind_{t+12}$	$\Delta lind_{t+18}$	$\Delta lind_{t+3}$	$\Delta lind_{t+6}$	$\Delta lind_{t+12}$	$\Delta lind_{t+18}$
BD	-0.090 [0.34]	-0.166** [0.08]	-0.233* [0.02]	-0.206* [0.04]	-0.067 [0.48]	-0.205* [0.03]	-0.215* [0.03]	-0.161** [0.10]
ES	-0.021 [0.82]	-0.086 [0.36]	-0.064 [0.52]	-0.139 [0.17]	-0.051 [0.59]	-0.138 [0.15]	-0.108 [0.27]	-0.127 [0.21]
FR	-0.053 [0.58]	-0.184* [0.05]	-0.061 [0.54]	0.053 [0.60]	-0.079 [0.40]	-0.199* [0.03]	-0.096 [0.33]	-0.063 [0.53]
GR	0.124 [0.19]	-0.065 [0.49]	0.033 [0.74]	0.033 [0.74]	0.080 [0.40]	-0.049 [0.61]	0.052 [0.60]	-0.037 [0.71]
IR	0.020 [0.83]	-0.025 [0.79]	-0.113 [0.25]	-0.060 [0.55]	0.031 [0.74]	0.043 [0.65]	-0.101 [0.30]	-0.021 [0.84]
IT	-0.081 [0.39]	-0.086 [0.37]	-0.121 [0.21]	-0.159 [0.11]	-0.097 [0.30]	-0.207* [0.03]	-0.175** [0.07]	-0.140 [0.16]
NL	0.088 [0.35]	-0.089 [0.35]	0.056 [0.57]	0.045 [0.66]	0.100 [0.29]	-0.112 [0.24]	-0.069 [0.48]	0.053 [0.60]
PT	-0.048 [0.61]	-0.062 [0.52]	-0.064 [0.51]	-0.013 [0.90]	-0.035 [0.71]	-0.048 [0.62]	-0.054 [0.58]	0.008 [0.94]

Notes. Probability values in square brackets; *, **: significant, respectively, at the 5% and 10% levels; $\widehat{lr_t}$: fitted value of the interest rate changes on loans to non-financial corporations provided by the non-linear estimates set out in table 3; $\Delta lind_t$: monthly rate of change of the industrial production index.

Figure 1. Long-term loan interest rate and 3 month Euribor



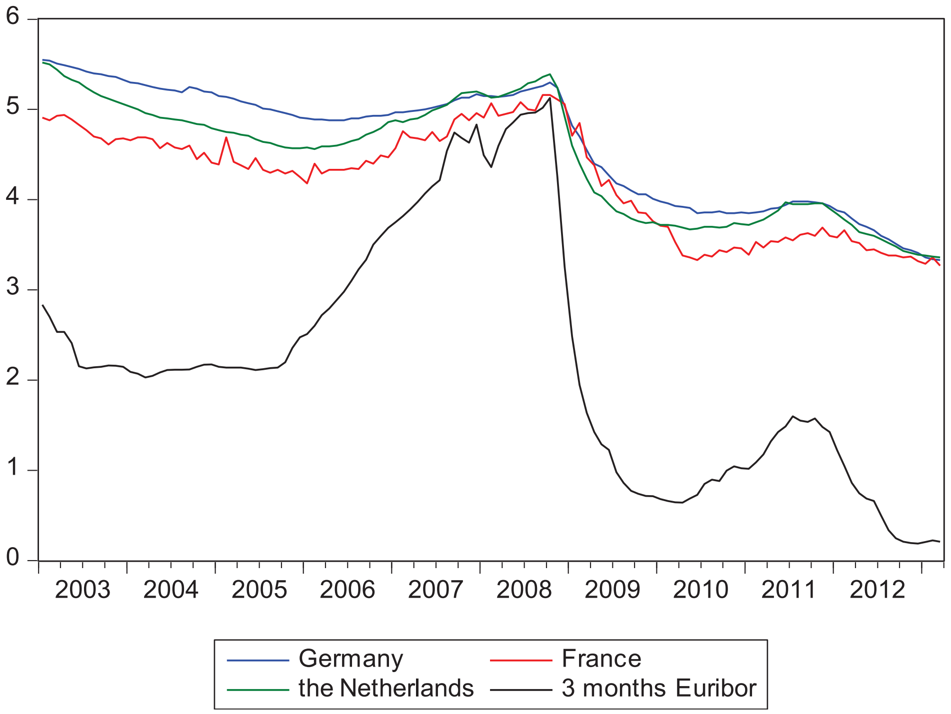
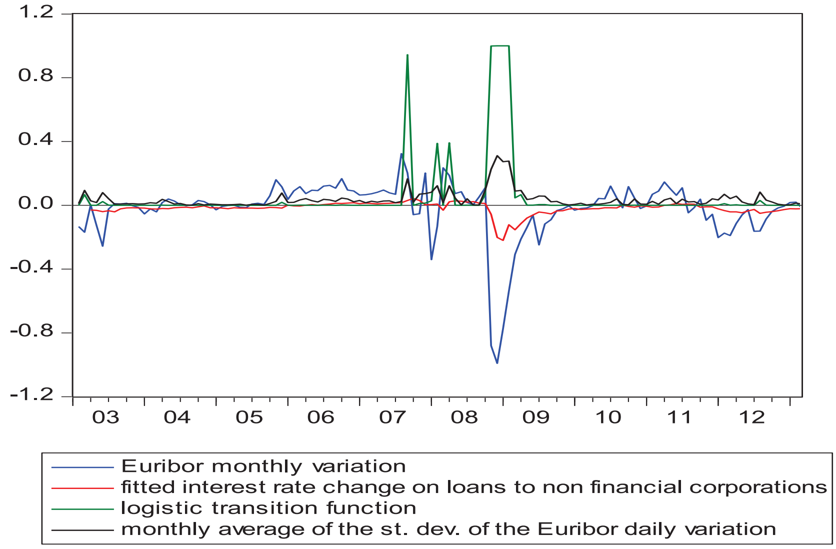


Figure 2. Euribor Changes, Fitted Interest Rate Changes, Transition Variable and Transition Function

Germany



Spain

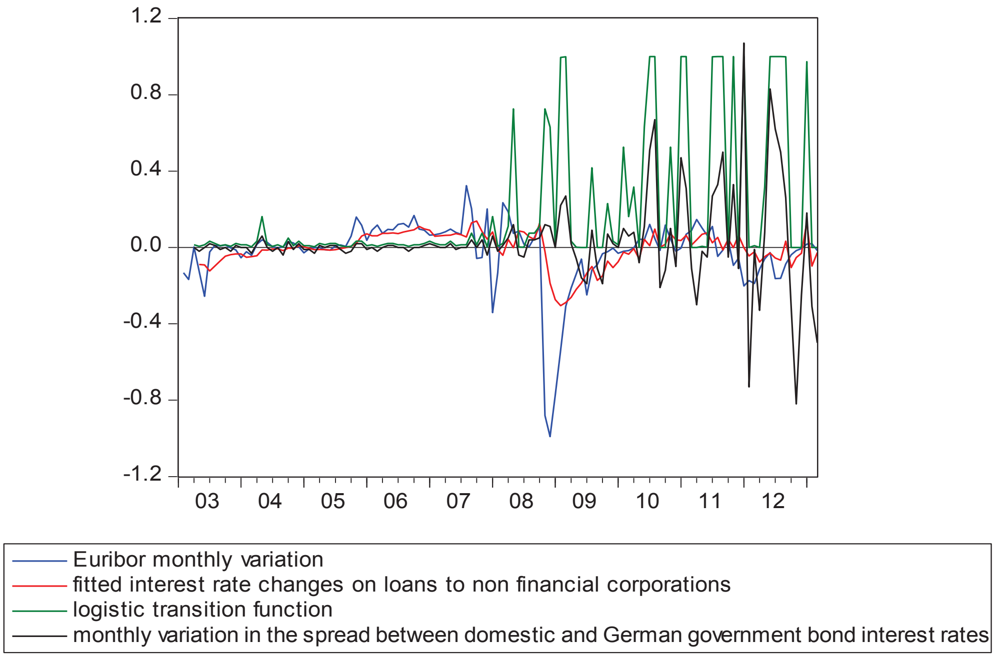
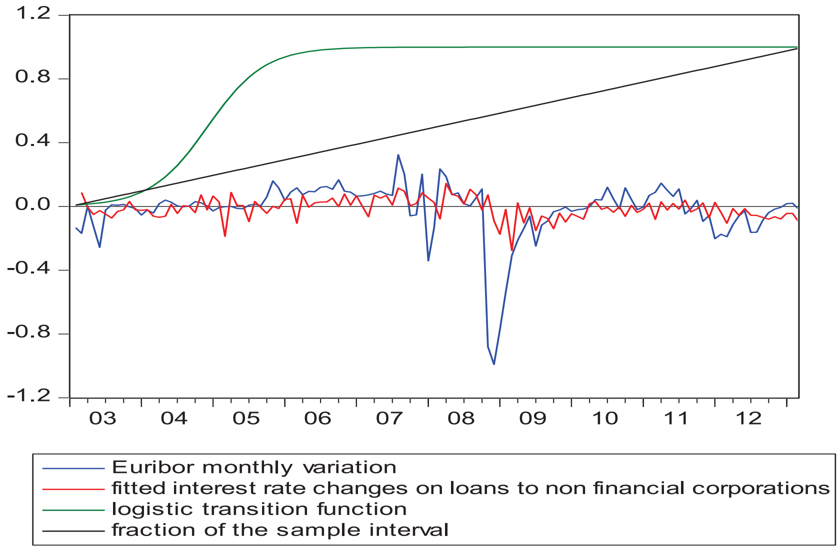


Figure 3. Euribor Changes, Fitted Interest Rate Changes, Transition Variable and Transition Function

France



Greece

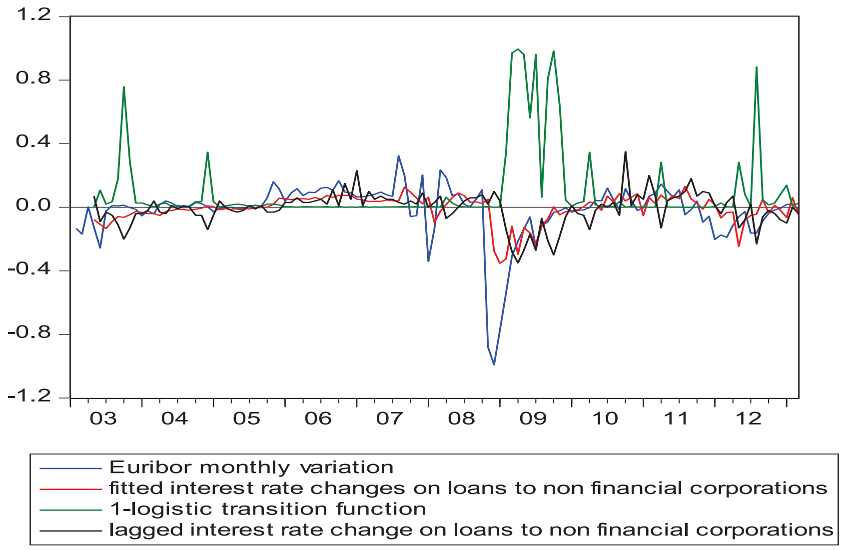
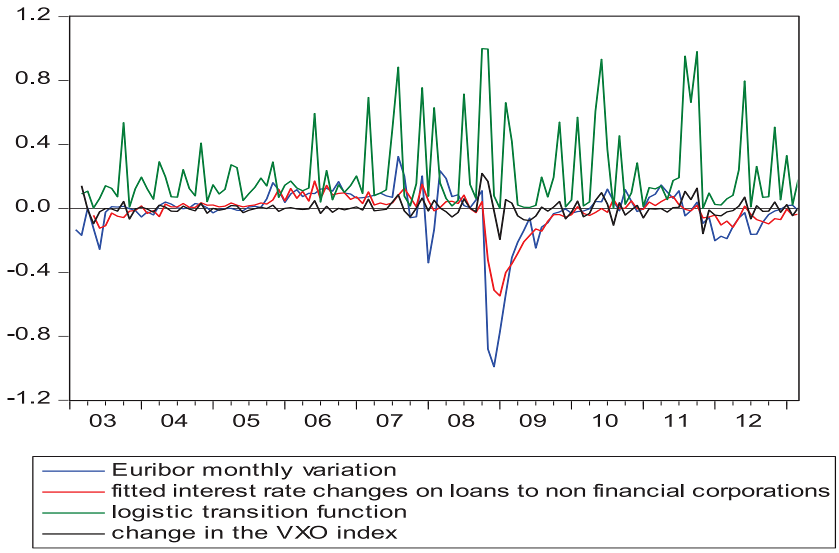


Figure 4. Euribor Changes, Fitted Interest Rate Changes, Transition Variable and Transition Function

Ireland



Italy

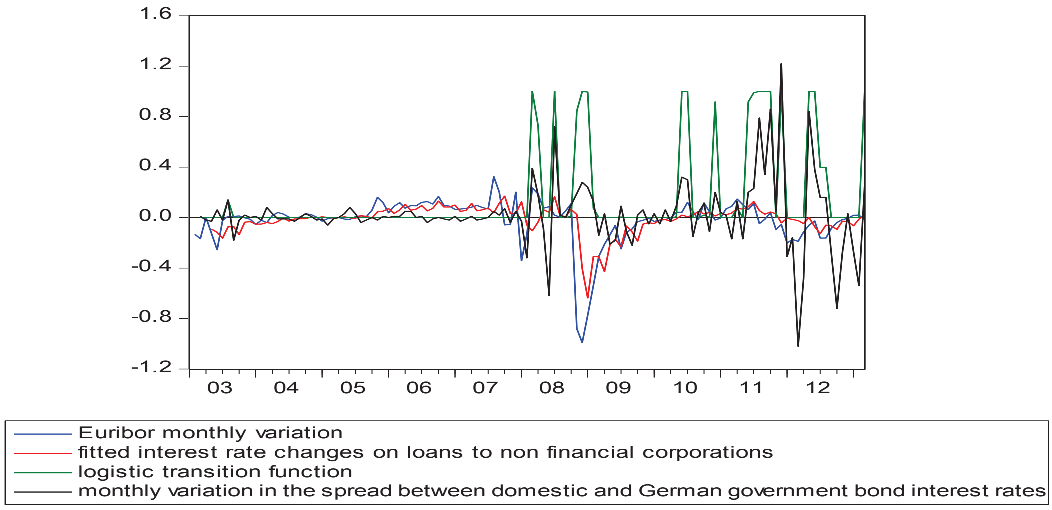
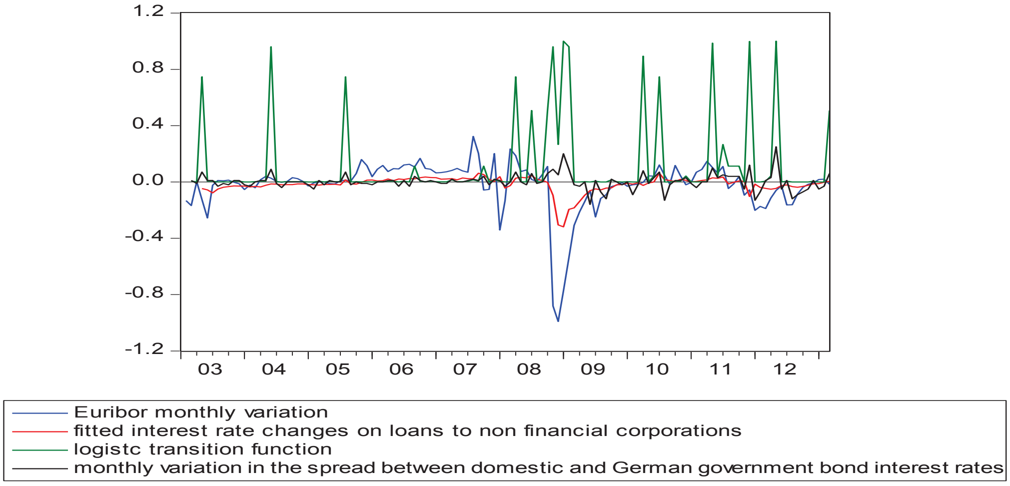


Figure 5. Euribor Changes, Fitted Interest Rate Changes, Transition Variable and Transition Function

The Netherlands



Portugal

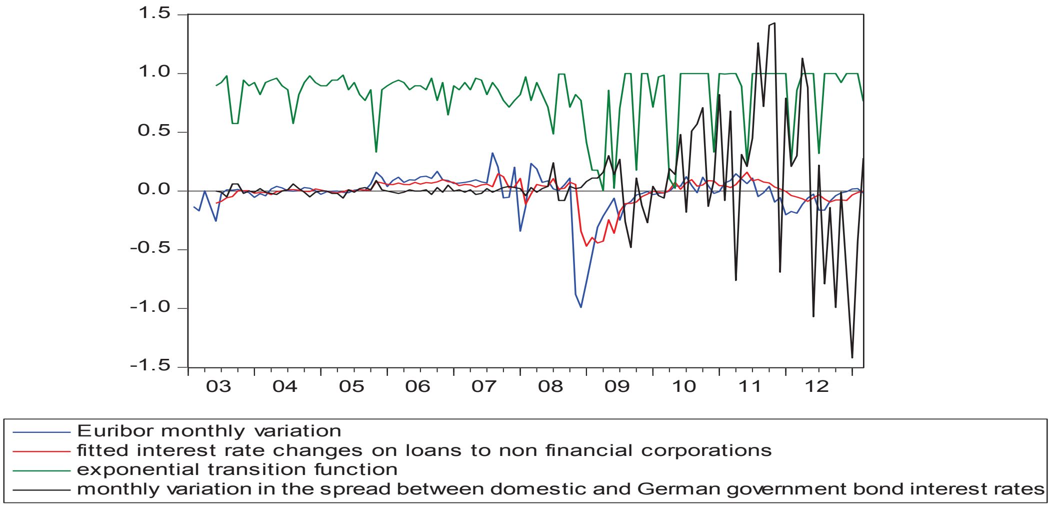
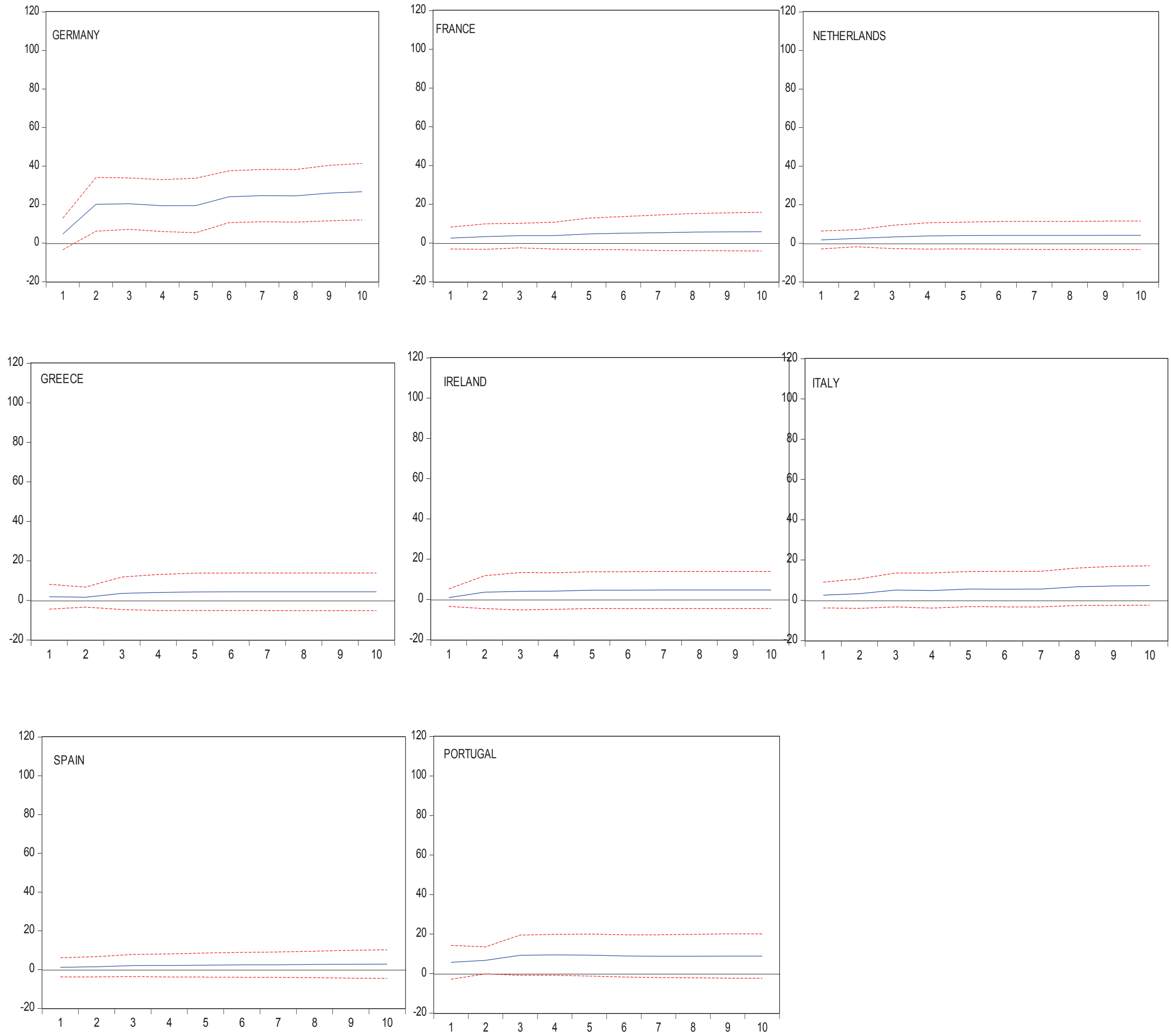


Figure 6. Variance decomposition

Variance of the rate of change of industrial output due to the fitted value of the loan interest rate



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Appendix

Table A.1 Descriptive Statistics

	2003M01-2013M03									
	Mean	Median	Std. Dev.	Skew.	Kur.	JB	AR(1)	AR(5)	HET(1)	HET(5)
Δeur_t	-0.021	0.003	0.180	-2.925	14.792	880.955 [0.000]	62.971 [0.000]	114.16 [0.000]	70.591 [0.000]	90.087 [0.000]
Δlr_t BD	-0.018	-0.010	0.042	-2.138	10.233	359.933 [0.000]	64.488 [0.000]	159.17 [0.000]	55.423 [0.000]	106.44 [0.000]
Δlr_t FR	-0.013	-0.020	0.098	-0.385	5.514	35.146 [0.000]	7.990 [0.005]	20.966 [0.001]	2.191 [0.139]	24.916 [0.000]
Δlr_t NL	-0.018	-0.010	0.059	-2.512	13.049	641.650 [0.000]	81.146 [0.000]	158.83 [0.000]	69.524 [0.000]	97.192 [0.000]
Δlr_t ES	-0.009	0.000	0.093	-1.115	4.589	38.129 [0.000]	74.763 [0.000]	249.16 [0.000]	77.021 [0.000]	183.96 [0.000]
Δlr_t GR	-0.008	0.010	0.103	-0.549	5.142	29.470 [0.000]	38.586 [0.000]	113.93 [0.000]	17.260 [0.000]	36.550 [0.000]
Δlr_t IR	-0.020	-0.010	0.129	-1.880	9.195	267.029 [0.000]	56.883 [0.000]	135.05 [0.000]	53.766 [0.000]	83.418 [0.000]
Δlr_t IT	-0.018	0.000	0.180	-2.428	11.921	524.435 [0.000]	64.777 [0.000]	182.82 [0.000]	24.933 [0.000]	58.770 [0.000]
Δlr_t PT	-0.011	0.015	0.118	-1.840	7.212	159.035 [0.000]	78.099 [0.000]	222.69 [0.000]	67.910 [0.000]	174.56 [0.000]
Δspr_t FR	0.006	0.000	0.092	1.172	13.840	625.249 [0.000]	0.241 [0.624]	7.441 [0.190]	16.348 [0.000]	25.293 [0.000]
Δspr_t NL	0.002	0.000	0.054	0.711	7.676	121.444 [0.000]	1.575 [0.209]	10.768 [0.056]	0.405 [0.525]	13.683 [0.018]
Δspr_t ES	0.0288	0.000	0.231	0.816	9.094	202.302 [0.000]	2.863 [0.091]	11.179 [0.048]	14.913 [0.000]	69.741 [0.000]
Δspr_t GR	0.080	0.010	1.486	-2.117	22.821	2088.29 [0.000]	1.577 [0.209]	19.208 [0.002]	4.114 [0.043]	34.437 [0.000]
Δspr_t IR	0.019	0.010	0.415	-0.881	13.084	532.679 [0.000]	6.142 [0.013]	16.056 [0.007]	8.010 [0.005]	21.801 [0.001]
Δspr_t IT	0.025	0.010	0.266	0.665	9.242	207.051 [0.000]	2.719 [0.099]	10.399 [0.065]	2.046 [0.153]	64.308 [0.000]
Δspr_t PT	0.038	0.010	0.423	0.038	6.625	66.844 [0.000]	8.302 [0.004]	44.101 [0.000]	23.311 [0.000]	81.379 [0.000]
Δvxo_t	-0.183	-0.440	5.102	0.481	7.768	119.299 [0.000]	0.253 [0.615]	6.591 [0.253]	8.363 [0.004]	27.448 [0.000]
$\Delta lind_t$ BD	0.0015	0.004	0.018	-0.910	6.695	86.246 [0.000]	0.430 [0.512]	16.090 [0.007]	9.824 [0.002]	24.824 [0.000]
$\Delta lind_t$FR	-0.001	-0.000	0.015	-0.358	4.158	9.434 [0.009]	5.265 [0.022]	12.915 [0.024]	4.843 [0.028]	8.093 [0.151]
$\Delta lind_t$NL	0.001	0.000	0.029	-0.032	5.271	26.231 [0.000]	15.682 [0.000]	17.232 [0.004]	10.662 [0.001]	12.038 [0.034]
$\Delta lind_t$ES	-0.002	0.000	0.014	-0.469	3.475	5.620 [0.000]	2.833 [0.092]	17.999 [0.003]	10.098 [0.000]	69.287 [0.000]
$\Delta lind_t$GR	-0.003	-0.002	0.028	-0.233	4.998	21.400 [0.000]	40.596 [0.000]	54.593 [0.000]	13.211 [0.000]	24.852 [0.000]
$\Delta lind_t$ IR	0.001	-0.001	0.060	-0.724	4.980	30.604 [0.000]	35.452 [0.000]	37.269 [0.000]	16.670 [0.000]	21.146 [0.001]
$\Delta lind_t$IT	-0.002	0.000	0.015	-0.354	3.447	3.565 [0.168]	0.080 [0.777]	20.740 [0.001]	23.396 [0.000]	60.652 [0.000]
$\Delta lind_t$ PT	-0.001	-0.005	0.032	-0.203	4.215	8.345 [0.015]	30.036 [0.000]	51.108 [0.000]	9.511 [0.002]	12.084 [0.034]

Notes. lr_t: long-run lending rate to non-financial corporations; eur_r: 3 month Euribor; spr_t: spread between the 10 years domestic government bond interest rate and the 10 years Bund interest rate; $\Delta lind_t$: monthly rate of change of the industrial production index; JB: Jarque-Bera normality test; AR(n): Ljung-Box test statistic for n-th order serial correlation; HET(n): Ljung-Box test statistic for n-th order serial correlation of the squared time series.

Table A.2 Unit Root tests

	Test Statistic	Number of breaks		Critical Values (5%)	Type of Test
		Intercept	Trend		
eur_t	-2.882	1	1*	-4.51	LM – LS
lr_t BD	-2.022	0	0	-3.44	ADF
lr_t ES	-3.662	1	1*	-4.51	LM – LS
lr_t FR	-4.794	2	2*	-5.66	LM – LS
lr_t GR	-3.406	2*	0	-3.84	LM – LS
lr_t IR	-4.087	2	2*	-5.67	LM – LS
lr_t IT	-3.173	1*	0	-3.56	LM – LS
lr_t NL	-3.309	1*	0	-3.56	LM – LS
lr_t PT	-2.440	0	0	-2.88	ADF
spr_t ES	-1.960	2(1*)	0	-4.45	LM – LS
spr_t FR	-4.023	1	1*	-4.50	LM – LS
spr_t GR	-3.241	2*	0	-3.84	LM – LS
spr_t IR	-4.143	1	1*	-4.45	LM – LS
spr_t IT	-4.497	1	1*	-4.50	LM – LS
spr_t NL	-4.899	2(1*)	2*	-5.73	LM – LS
spr_t PT	-3.138	2*	0	-3.84	LM – LS
$eursd_t$	-3.572	1	0	-2.88	ADF
vxo_t	-3.072	1	0	-2.88	ADF
$\Delta lind_t$ BD	-8.099	0	0	-1.94	ADF
$\Delta lind_t$ ES	-7.421	0	0	-1.94	ADF
$\Delta lind_t$ FR	-6.764	0	0	-1.94	ADF
$\Delta lind_t$ GR	-12.410	0	0	-1.94	ADF
$\Delta lind_t$ IR	-8.702	0	0	-1.94	ADF
$\Delta lind_t$ IT	-3.806	0	0	-1.94	ADF
$\Delta lind_t$ NL	-10.687	0	0	-1.94	ADF
$\Delta lind_t$ PT	-8.109	0	0	-1.94	ADF

Notes. lr_t: long-run lending rate to non-financial corporations; eur_r: 3 month Euribor; spr_t: spread between the 10 years domestic government bond interest rate and the 10 years Bund interest rate; $eursd_t$: monthly average of the daily standard deviation of the 3 month Euribor, computed as the square root of the squared daily variations of the Euribor; vxo_t: monthly VXO index; $\Delta lind_t$: monthly rate of change of the industrial production index. *significant at 5% ; ADF: Augmented Dickey Fuller test statistic; LM – LS : Lee and Strazicich (2003, 2004) Minimum Lagrange multiplier test statistic.