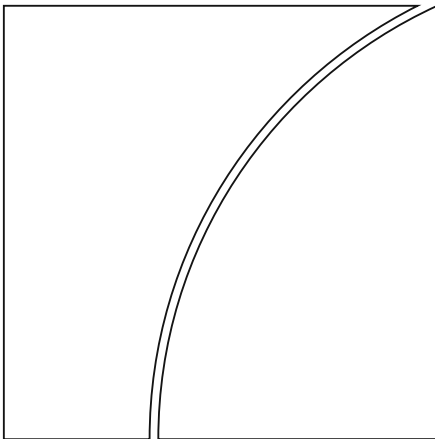# BIS Working Papers
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Modelling the Time-Variation in Euro Area Lending Spreads

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Modelling the Time-Variation in Euro Area Lending Spreads

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Abstract

Using a Markov-switching VAR with endogenous transition probabilities, we analyse what has triggered the interest rate pass-through impairment for Italy, Ireland, Spain and Portugal. We find that global risk factors have contributed to higher lending rates in Italy and Spain, problems in the banking sector help to explain the impairment in Spain, and fiscal problems and contagion effects have contributed in Italy and Ireland. We also find that the ECB's unconventional monetary policy announcements have had temporary positive effects in Italy. Due to the zero lower bound these findings are amplified if EONIA is used as a measure of the policy rate. We did not detect changes in the monetary policy transmission for Portugal.

JEL classification: E43; E52; C32

Keywords: Lending rates; interest rate pass-through

1 Introduction

The events triggered by the global financial crisis of 2008 – 2009 have proved to be some of the most significant economic phenomena observed in recent decades. The costs of the downturn have far exceeded those of any previous post-WWII recession. Moreover, not only did financial developments trigger the downturn, but as events unfolded, the financial sector found itself at the epicentre of the crisis. The collapse of major financial institutions, reduced asset values, the interruption of credit flows, the loss of confidence in bond and credit markets, and the fear of default by euro area countries, were all extraordinary economic occurrences. In addition, aggressive monetary interventions during the crisis charted new ground both in scale and in scope.

For economists, the consequence of these events has been a revival of the macroeconomic-finance nexus, as well as a growing interest in nonlinear modelling approaches. The analytical models that have become standard in the field over the last generation seem to have been unsuited to explaining what was occurring during this unusually significant episode, and now unable to incorporate most of the widely accepted accounts of it. If the economy is subject to important nonlinearities, certain results that derive from linear models do not carry over, with major implications for the monetary policy transmission channel.

Interest rate pass-through is of central importance for monetary policy. With the adoption of a common currency, the euro area was faced with the challenge that a single policy had to account for the heterogeneity among its members. As such, the transmission of monetary policy of the European Central Bank (ECB) has been an important topic for researchers. Before the financial crisis, many studies found that while interest rates appear to be sticky in the short run, there exists complete long-term pass-through, and the adoption of a single monetary policy has improved the transmission and the velocity of the short-run pass-through [Bindseil and Seitz (2001), Angeloni et al. (2003), Sander and Kleimeier (2004), De Bondt (2005), Affinito and Farabullini (2006), Gambacorta (2008)].

However, the recent crises - the global financial turmoil and the euro area sovereign debt crisis - have put the banking systems under severe stress. If interest rates were far higher than in Germany and the associated credit squeeze persisted, this would threaten one of the fundamental aims of the euro area: to create a single market with an integrated economy. Indeed, there was clear evidence of a fragmentation of financial markets, and of a significant diver-

gence between lending and policy rates in the euro area in 2011-13.[1] This, in turn, has had heterogeneous effects on the monetary policy transmission across the different member states [Cihak et al. (2009), Gambacorta and Marques (2011), Ciccarelli et al. (2013), Al-eyd and Berkmen (2013), Illes and Lombardi (2013), Paries et al. (2014), Hristov et al. (2014a)].

While the breakdown in the pass-through has been documented thoroughly, numerous questions remain unanswered. What has driven the change in the interest rate pass-through among euro area member countries during the crisis? What were the trigger variables? Are there country-specific fundamentals that affect lending spreads or is it a matter of flight-to-quality and flight-to-safety concerns? We consider the hypothesis that nonlinear dynamics have driven lending spreads during the crisis. Initial shocks to economic fundamentals may have been exacerbated by endogenous mechanisms. How does the pricing of risk take place and can we identify endogenous factors triggering amplification? The answers will assist in the monitoring and pricing of risk, as well as in the prevention of financial fragmentation. This paper joins a growing literature that has centred on identifying nonlinearities using formal statistical methods.[2]

We investigate the heterogeneous effects of monetary policy across several euro area countries through the lens of a quasi nonlinear Vector Autoregressive model – a VAR subject to regime shifts with endogenous transition between the states. We incorporate the switching mechanism through time-varying transition probabilities that help us identify potential triggers.[3] In this setup, model uncertainty takes the form of different modes that follow a Markov process. It can be thought of as a model encompassing a number of possible representations of the world.

A few studies have investigated the joint variation of macro fundamentals and credit spreads by incorporating the possibility of regime shifts [David (2008)], and a handful have documented the change in interest rate pass-through. Cihak et al. (2009) use a standard bi-variate VAR in the spirit of De Bondt (2005) and a general equilibrium framework to show a slowdown in the pass-through. They also analyse unconventional monetary policy measures and demonstrate that to a certain extent they had helped alleviate the problem. Ciccarelli et al.

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[1] The fragmentation of financial markets in the euro area has decreased over the past year.
[2] See Silvestrini and Zaghini (2015) for an up-to-date survey of the theoretical and empirical contributions exploring the linkages between financial factors and the real economy in nonlinear frameworks.
[3] Evidence that macroeconomic time series follow a Markov process has led macroeconomists to develop monetary policy frameworks with regime shifts. For example, Svensson and Williams (2009) have developed a general form of model uncertainty that remains tractable, using so-called Markov-jump-linear-quadratic models. There is a growing body of Markov-switching DSGE and VAR models [Sims et al. (2008), Davig and Doh (2009), Farmer et al. (2009), Farmer et al. (2011), Bianchi (2012) among others].

(2013) quantify the heterogeneous effects of monetary policy on GDP across the member states by means of a recursive VAR and document the time-variation in interest rate pass-through. Furthermore, they show that the effect on GDP of monetary policy shocks is amplified through the credit channel, and that the bank-lending channel has been non-existent due to unconventional monetary policy measures of the ECB.

Hristov et al. (2014b) examine the effectiveness of the Outright Monetary Transmission Program (OMT) of the ECB by means of a time-varying parameter VAR (TVP −VAR) based on Primiceri (2005), and Paries et al. (2014) capture the breakdown in interest rate pass-through by a single equation framework. The model is extended to account for bond yields, which partly explain the lending spreads. Hristov et al. (2014a) document the incompleteness of the pass-through after the crisis using a panel VAR and a DSGE model. Aristei and Gallo (2014) also use the simple bi-variate framework of De Bondt (2005) in the context of a Markov-switching VAR (MS-VAR) and a Markov-switching Vector Error Correction Model (MS-VECM) with exogenous probabilities, and establish lower efficiency and time-variation in the transmission of monetary policy. Our study differs significantly from Aristei and Gallo (2014), since our framework has the important addition of endogenous transition probabilities to address the question at hand.

There are a couple of novel studies that argue that if one takes several considerations into account, the high lending rates might be explained even in the face of near zero policy rates. Ciccarelli et al. (2013) and Illes et al. (2015) suggest that after the crisis the interbank rate might not be a good proxy for bank funding costs and thus should not be taken as a major determinant for the lending rates, because access to funds was impaired after the meltdown. Illes et al. (2015) create a benchmark for bank funding costs for each country both in the short and the long term, which accounts for the levels of the lending rates. They construct a weighted average cost of liabilities (WACL), which consists of several components including covered bonds, five-year credit default swaps, deposit liabilities and open market operations. Borstel et al. (2015) utilise a factor-augmented VAR (FAVAR) model to incorporate sovereign and bank funding risk, conventional and unconventional monetary policy, and argue that it is not the interest rate pass-through that has changed, but rather its composition. Our results hold even if we account for the zero lower bound and the impairment in the bank funding channel.

Through endogenising the transition probabilities between different regimes, we find that global risk factors have contributed to higher lending rates in Italy and Spain, problems in the

banking sector help to explain the impairment in Spain, and fiscal problems and contagion effects have contributed in Italy and Ireland. We also find that the ECB's unconventional monetary policy announcements have had temporary positive effects in Italy. Due to the zero lower bound these findings are amplified if EONIA is used as a measure of the policy rate. For Portugal we cannot identify any significant change in the interest rate pass-through, in contrast to the other countries.

The paper is organized as follows. The next section introduces the key data in our study, namely lending rates and sovereign bond yields. Section 3 presents the econometric methodology of the paper. Section 4 lays out the main results, and section 5 discusses potential problems and extensions to the main specification. Finally, section 6 concludes.

2 Lending Spreads and Sovereign Bond Spreads

We consider the heterogeneous time-variation across the euro area of long-term lending rates to non-financial firms. In particular, we use interest rates on loans over €1 million, with maturities over one year, to non-financial firms for new businesses (other than revolving loans and overdrafts, convenience and extended credit card debt) [ECB (2003)].

We examine four countries in this study: Italy, Spain, Ireland, and Portugal, and consider the spread of the long-term lending rate to Germany, $r_t^h = R_t^h - R_t^{DE}$, where R_t^h is the long-term lending rate in country h at time t. All countries are identified by the respective two-letter ISO code. As a link between the short-term policy rate and the long-term lending rate we include ten-year government bond yield spreads relative to Germany as an endogenous variable, $g_t^h = G_t^h - G_t^{DE}$, which reflect country-specific market sentiment.[4]

Monthly data for the evolution of lending rate spreads and government bond yield spreads over time (with evolution over time denoted by different colours) is shown in Figure 1 for each of the four euro area countries studied in this paper, Italy, Spain, Ireland and Portugal.

We can see that at the start of the sample period, in 2004, government bond yield spreads tended to be close to zero for all four countries. Lending rate spreads for Italy, Spain and Ireland were even negative in many cases in 2004. At the height of the euro area sovereign debt crisis, government bond yield spreads rose to much higher values than the lending spreads, before falling back considerably towards the end of the sample period.

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[4] Since the Markov switching estimator performs better for long time series, Greece has not been considered, since only shorter time series are available for Greece than for Italy, Spain, Ireland and Portugal.

Government bond spread versus lending rate spread relative to Germany

In per cent; over time   Figure 1
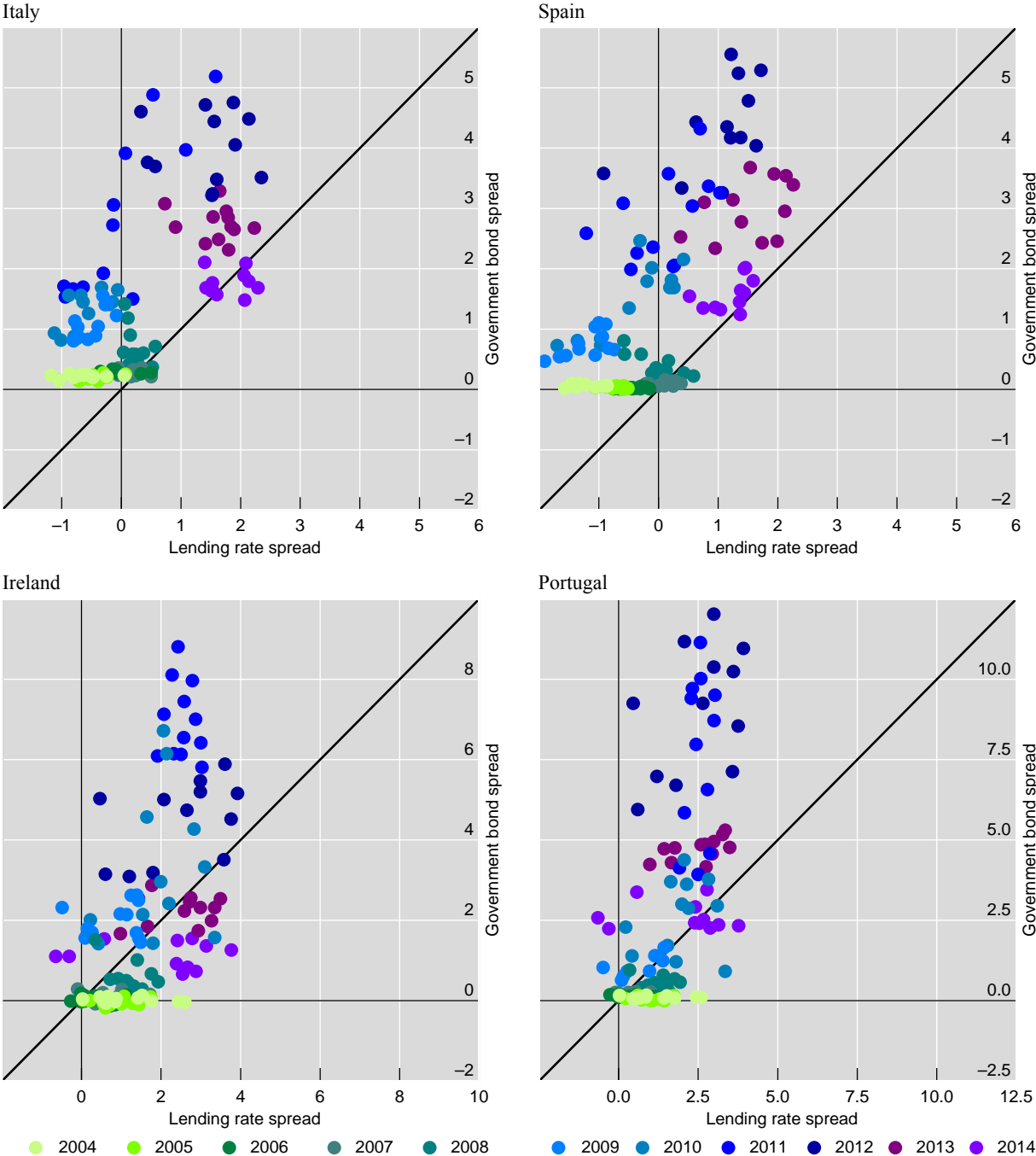
Source: ECB, FRED database, authors' calculations

Possible reasons for this include that the average duration of the lending could differ across countries, and that yields in Germany may have been affected by a flight to safety. Moreover, the smaller increase in the lending spreads could also be due to a shift in demand for lending during the financial crisis, with fewer and more highly rated firms borrowing during the crisis, as well as due to a shift in the supply of credit in the form of credit rationing. Furthermore,

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lending to firms relying on borrowing in the corporate bond market rather than from banks are not captured in these lending rates. Although lending rate spreads did not rise as much as government bond spreads during the crisis, lending rate spreads tended to remain elevated towards the end of the sample period.

3 Econometric Methodology

To study the changes in the interest rate pass-through we assume the following data generating process of a structural VAR with time-varying parameters,

$$A_0(s_t)y_t = \mathbf{a}(s_t) + A_1(s_t)\,y_{t-1} + \cdots + A_l(s_t)\,y_{t-l} + \boldsymbol{\varepsilon}_t. \tag{1}$$

$A_0(s_t), \dots, A_l(s_t)$ are coefficient matrices, $\mathbf{a}(s_t)$ is a vector of constants, and the structural innovations $\boldsymbol{\varepsilon}_t$ follow a normal distribution with zero mean and stochastic volatility $\boldsymbol{\varepsilon}_t \sim N(\mathbf{0}, \boldsymbol{\Psi}(s_t))$. Each set of coefficients is associated with the respective state $s_t = \{1, \dots, n^s\}$, where n^s is the number of regimes.[5] The vector y_t contains n endogenous variables and l denotes the lag order, selected according to standard information criteria. We use three lags for Spain, four for Italy, two for Ireland and two for Portugal. To determine the maximum lag length for the tests we follow Schwert (1989). We assume endogeneity of all the variables in the system and estimate the dynamics of purely exogenous shocks.[6] Assuming a two-state stochastic Markov process ($n^s = 2$), the shifts across regimes are governed by transition probabilities given by the probability matrix

$$P = \begin{bmatrix} p(Z) & 1 - p(Z) \\ 1 - q(Z) & q(Z) \end{bmatrix}. \tag{2}$$

Instead of assuming an exogenous switching mechanism, we set both p and q as the outcome of a probit model with regressors collected in the vector of trigger variables $Z_t = [1, z_{1,t}, \dots z_{k,t}]$. Let s_t^* be a latent variable determined by the following regression:

$$s_t^* = \gamma_0 + \gamma_1 z_{1,t-m} + \cdots + \gamma_k z_{k,t-m} + \omega_t. \tag{3}$$

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[5] Jovanovic (1989) has shown that in case of sunspots it is necessary to distinguish the dynamic of the fundamentals process from the sunspot process. A Markov regime-switching model provides a flexible framework that allows to distinguish between the two processes. The regime shifts can then be interpreted as jumps between multiple equilibria.

[6] The Markov-switching framework implies complications for empirical work that attempts to estimate how interest rate spreads respond to changes in monetary policy. Complications arise due to the nonlinearity in the decision rule, implying the interest rate pass-through is a function of the regime. To borrow language from Leeper and Zha (2003), an interest rate pass-through within the band where the decision rule is approximately linear can be referred to as a "modest" monetary policy intervention. A "non-modest" policy intervention causes agents to alter their inference regarding the current regime, resulting in a response that is greatly at odds with the predictions of fixed-regime models.

The error term in equation (3) follows a standard normal distribution $\omega_t \sim N(0,1)$ and we set the lag of the trigger variables m to 1 to address potential endogeneity problems. This method of using the lag of the trigger variables to address potential endogeneity problems, so that they are predetermined, may be problematic if expectations play a central role.

The vector of coefficients $\boldsymbol{\Gamma} = [\boldsymbol{\gamma_0}, \boldsymbol{\gamma_1}, \dots, \boldsymbol{\gamma_k}]'$ is of primary interest for this study, as the variables governing the transition probabilities would prove crucial for describing the nature of the euro area crisis. Moreover, statistically significant effects of contagion variables on transition probabilities would lead to the conclusion that lending spreads are driven not only by fundamentals but also by contagion, e.g. due to confidence effects. Significance of both fundamentals and contagion variables would indicate that various crisis models are not mutually exclusive.

Under the assumption of two regimes, the threshold for the observable counterpart s_t of the latent variable s_t^* is defined as:

$$s_t = \begin{cases} \mathbf{1}, & s_t^* < \mathbf{0}, \\ \mathbf{2}, & s_t^* \geq \mathbf{0}. \end{cases} \tag{4}$$

Therefore, the transition probabilities are determined by the following probit model:

$$p(Z) = P(s_t = 1 | s_{t-1} = 1) = P(\omega_t < -Z_{t-m}\boldsymbol{\Gamma}) = \boldsymbol{\Phi}(-Z_{t-m}\boldsymbol{\Gamma}). \tag{5}$$

$$q(Z) = P(s_t = 2 | s_{t-1} = 2) = P(\omega_t \geq -Z_{t-m}\boldsymbol{\Gamma}) = 1 - \boldsymbol{\Phi}(Z_{t-m}\boldsymbol{\Gamma}) \tag{6}$$

The complete model, given by (1)–(6), is based on Goldfeld and Quandt (1973), Filardo (1994) and Filardo and Gordon (1998), and nests the case of fixed probabilities if the variables in **Z** are not informative for the probit regression.

The assumption of the existence of two states is not innocuous. In general, specifying a Markov regime-switching model requires a test to confirm the presence and the number of multiple regimes. The first step is to test the null hypothesis of one regime against the hypothesis of Markov switching between two regimes. If the null hypothesis can be rejected, then one can proceed to estimate the Markov regime-switching models with two or more regimes.

Conducting proper inference, however, is exceptionally challenging. In particular, testing for the number of regimes requires the use of nonstandard test statistics and critical values that may differ across model specifications. Cho and White (2007) demonstrate that because of the unusually complicated nature of the null space, the appropriate measure for a test of multiple regimes is a quasi-likelihood-ratio (QLR) statistic. They provide an asymptotic null distribution for this test statistic from which critical values can be calculated. Unfortunately, Carter

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and Steigerwald (2012) establish that the estimator computed using the QLR-likelihood is inconsistent if the covariates include lagged dependent variables. Thus, this test cannot be applied to our modelling setup.

Since we cannot pin down the amount of regimes by statistical inference we have to take another approach. We choose two states for the data generating process in equation (1). Our choice is motivated by the main question – has there been a change in interest rate pass-through, and if so, what has been driving it? One approach to answer the question would be to use a gradual change in parameters as in Ciccarelli et al. (2013) and Hristov et al. (2014b). The other extreme is modelling a binary outcome.

Even though we cannot test for the presence of two regimes directly, the advantage lies in the fact that it does not impose that regimes be significantly different from one another as we use the same prior in both states. If the data do not support distinct parameters, we would find overlapping posterior distributions of the coefficients and similar impulse responses.

On the other hand, if there are more than two regimes and there is even higher fragmentation among euro area members, our results will average the multiple states in two distinct sets and may be interpreted as a lower bound, i.e. the "true" impulse responses would be even more pronounced. Moreover, this is consistent with the Markov-switching literature, where the presence of two regimes is often assumed, which dates back to Hamilton (1989).[7] A final consideration is computational efficiency and the curse of dimensionality. Every additional state reduces the sample size proportionally, while increasing the number of parameters to be estimated exponentially, which argues against additional states.

3.1 Bayesian analysis

We cast the model of equation (1) in a reduced form by pre-multiplying the structural form with the impact matrix $A_0(s_t)^{-1}$ and redefining all matrices accordingly

$$y_t = \mathbf{b}(s_t) + \mathbf{B}_1(s_t)\, y_{t-1} + \cdots + \mathbf{B}_l(s_t)\, y_{t-l} + u_t. \qquad (7)$$

The residuals $u_t \sim N(\mathbf{0}, \Sigma(s_t))$ in equation (7) and their connection to the structural shocks are of primary interest in any VAR study. For this link we choose a standard Cholesky decomposition, which is consistent with the pass-through literature. This choice is motivated by the economic theory that policy rates determine lending rates and can do so instantaneously, but not vice versa. The reduced form VAR(l) model may be rewritten in its VAR(1) form by

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[7] For an application of two state models to monetary policy, term structure and bond/CDS spreads, see for example Amisano and Tristani (2009), Lanne et al. (2010), and Blommestein et al. (2012).

imposing $Y = [y_1 \ldots y_T]'$, $X = [X_1, \ldots X_T]'$, $X_t = [y'_{t-1} \ldots y_{t-l}, 1]'$, $U = [u_1 \ldots u_T]'$ and $\beta = [B_1 \ldots B_l, b]'$:

$$Y = X\beta + U. \tag{8}$$

For estimation we employ Bayesian methods and incorporate the priors following Banbura et al. (2010) (see the technical appendix for details).

3.2   Explanatory variables in the regime-switching VAR

The interest-rate pass through consists of two stages. In the first stage, the ECB lends funds to financial institutions in its open market operations at the policy rate which determines the interbank rate. The second stage is the transmission from the interbank rate to the lending rates for non-financial institutions. The ECB sets the policy rate and adjusts it several times a year, which makes the official policy rate a step-wise function - unsuitable for empirical analysis. Typically, the literature assumes that the first-stage transmission is always perfect in the sense that interbank rates such as the overnight EONIA rate or the EURIBOR rate are a good proxy for the policy rate. Nevertheless, recent studies have noted that this may not be an appropriate choice any more.

On the one hand, Hassler and Nautz (2008) document that the link in the first stage has broken down. On the other hand, the zero lower bound plays an important role in studying interest rate pass-through, because the lending rates move freely even when the constraint is binding for the money market rate. Thus the empirical models might "capture" a breakdown in the pass-through solely due to the flatness of the proxy for the policy rate.

As an alternative the literature has suggested using a different proxy for the policy rate, namely the shadow short rate (SSR) [Wu and Xia (2014), Krippner (2014), Pericoli and Taboga (2015) and Borstel et al. (2015)]. The shadow rate is derived using nonlinear term structure models and is allowed to take negative values, which alleviates the problem of the zero lower bound. Unfortunately, this does not come without a cost. Since the series are based on theoretical foundations, they do not represent interest rates at which economic agents can transact [Krippner (2014)]. Hence they do not reflect the banks' funding conditions.

Thus, choosing a proxy for the policy rate is not a straightforward task. To this end, we explore several different specifications. In the main section we employ EONIA as the policy rate i_t, while in the robustness section we estimate the model using alternative measures - the shadow short rate estimates of Wu and Xia (2014), Krippner (2014), Pericoli and Taboga (2015).

VAR variables

Italy

Spain

Portugal

Ireland

— EONIA   — Government bonds   — Lending rate
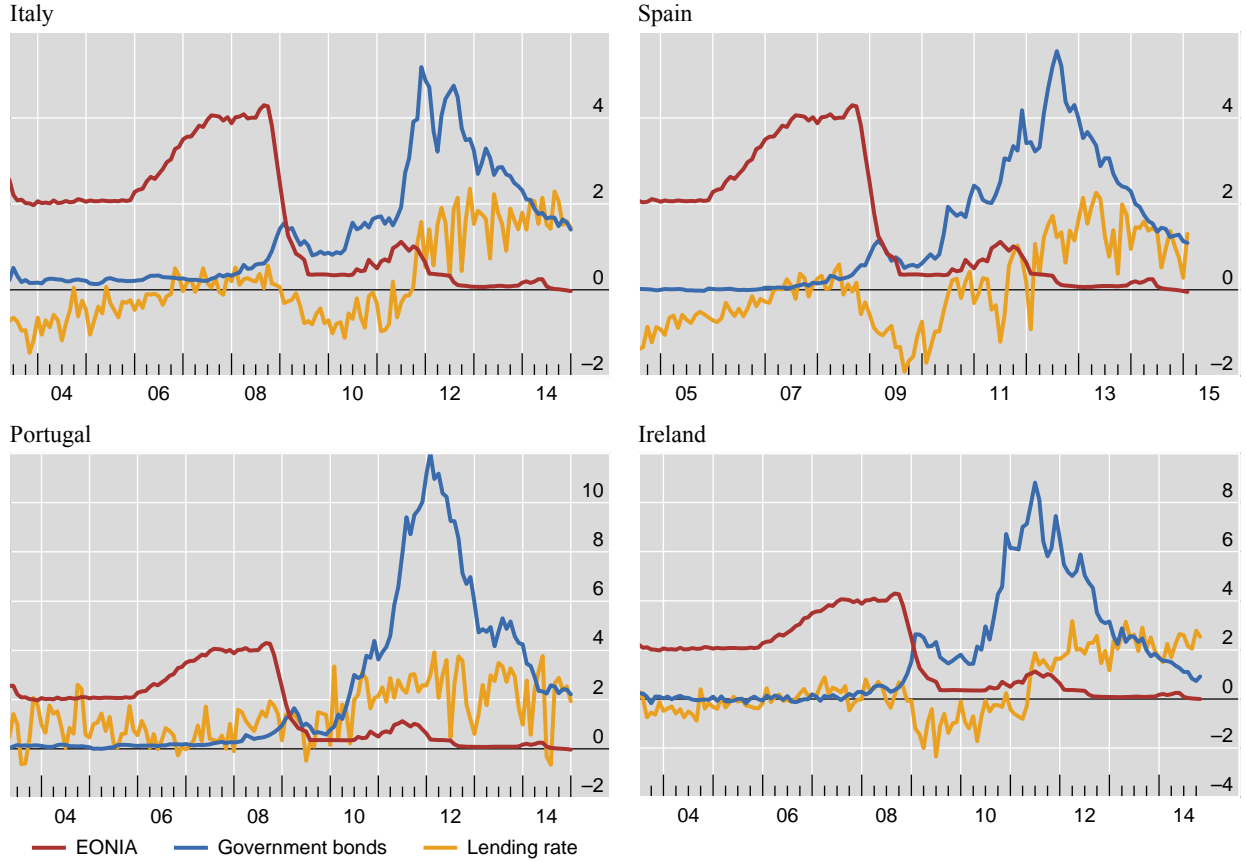
Source: ECB, FRED database, authors' calculations.

As a robustness exercise, we also incorporate a proxy for bank funding costs of Illes et al. (2015), namely banks' weighted average cost of liabilities, and our results reported below remain largely unchanged (see online Appendix I).

To summarize, long-term lending rate spreads are explained by the policy rate i_t, approximated by EONIA or a shadow short rate, and the 10-year government bond spread. The vector of endogenous variables y_t^h for country h at time t is given by

$$y_t^h = [i_t \quad g_t^h \quad r_t^h]. \tag{9}$$

We plot the time series for each country in Figure 2.

3.3   Choosing the trigger variables Z

The choice of trigger variables in the probit model (3) is of crucial importance. Omitting relevant explanatory variables increases the variance of the error term, which potentially biases

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the estimates. Therefore, care should be taken when specifying equation (3) of the regime-switching VAR. We use a multitude of macro and financial variables as indicators, and test each one for the informational content regarding the switching mechanism.

Macroeconomic developments are among the main determinants of interest rate spreads. To capture the impact of macroeconomic fundamentals, three main types of variables will be considered in this study. The full set of variables including data sources may be found in the online appendix.

The first type groups country specific variables such as broad macroeconomic indicators (including industrial production growth, HICP inflation and the debt-to-GDP ratio[8]), financial market information such as bank stock indices and CDS spreads, as well as information regarding a country's borrowing within the ECB's main financing operations (MROs) and long-term refinancing operations (LTROs).

Apart from domestic conditions, interest rate spreads are also influenced by global conditions and contagion, which we capture in a second group of variables. Tighter global liquidity and/or contagion might lead to fund outflows from countries, resulting in larger spreads. There are several price-based or quantity-based measures of global liquidity and contagion in the literature. We take the VSTOXX and the MOVE index to represent market sentiment about global financial conditions. To assess the issue of contagion, we also incorporate lending rate spreads of different countries as potential trigger variables. Another index of interest is the European economic policy uncertainty index from Baker et al. (2015).

A third group of variables captures monetary policy announcement effects. We introduce two dummy variables that aim to capture the effects of policy announcements from the ECB. The first captures the LTRO announcements from July, October, and December 2009. The second variable captures several monetary policy announcements from July, August, and September of 2012. In July, the president of the ECB Mario Draghi communicated the ECB's support for the euro in a panel discussion.[9] In September the ECB announced the Outright Monetary Transactions (OMT) programme. Altavilla et al. (2014) find that these measures alone have reduced sovereign bond yields in Italy and Spain by more than two percentage points.

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[8] We also considered net fiscal balances as a percentage of GDP in addition to the debt-to-GDP ratio, which left the results reported below largely unchanged.
[9] The speech has been often labelled in the media as the "Whatever-it-takes speech". De Grauwe and Ji (2013) have shown empirically that the temporary disconnect of market expectations from fundamentals and the existence of jumps between multiple equilibria was an important element of the euro area crisis.

3.4 Time series properties

Typically, in time series analysis the question of stationarity is meticulously discussed. Testing the VAR variables with an ADF test reveals that most variables appear stationary with the exception of the government bond spreads of Spain and Ireland and the lending rate spread in Ireland. Nevertheless, in MS-VAR models, the stationarity assumption is not needed for the dependent and the independent variables. These models rely upon something quite a bit stronger than stationary residuals, namely that the true residuals (if the regime were known) are independent and normal. Because the regime is not known, that is not really a testable hypothesis. All one can test is whether the standardized residuals are uncorrelated and have constant variance. Note that passing those tests does not make the model valid, just not rejectable. We present tests for normality of the estimated residuals in the online appendix.

Turning to the probit model, the stationarity issue is not as simple. We test all variables for unit roots with the ADF and Pierre-Perron tests. Since many of the variables appear non-stationary, there might be several pitfalls. Park and Phillips (2000) have shown that while the estimator in binary dependent variable models is consistent even with integrated regressors, it has some special asymptotic properties. Riddel (2003) has documented that the explanatory variables might fail to pass the t-tests for coefficient diagnostics even when they are indeed informative.

Therefore we rely on our use of Bayesian techniques and the consistency property to alleviate the problem — we use credible (probability) intervals of the posterior distribution instead of asymptotic intervals. Informally, to assess significance, one can look at the estimated transition probabilities, since the model should reduce to the fixed probability case (flat probabilities) if the variable is insignificant. We base our results on the formal analysis unless otherwise noted. Furthermore, as is standard practice, we estimate first differences and run the models with them as a separate case. This brings the total number of trigger variables per country to twenty five. To address potential multicollinearity issues at the estimation stage, we do not choose any pair of variables with a correlation coefficient greater than 0.5 in absolute value. Correlation matrices may be found in the online appendix.

Another potential issue associated with probit models is the inclusion of dummy variables. This may lead to the problem of quasi-complete separation, which arises when the explanatory variable has too much predictive power over the dependent variable. In the case of binary variables too many coinciding "ones" or "zeros" on both sides of the regression might distort the estimator. Gelman et al. (2008) suggest the use of Bayesian estimation over the standard

maximum likelihood estimator to remedy the problem, or, if the former is undesirable, to drop the variables in question. Therefore, when setting the prior for the trigger variables regression, we omit the dummies at the maximum likelihood stage and choose them separately. For brevity, we report the results only for the regressors that are significantly different from zero for each country, with the exception of the policy announcement variables, which are included in all cases.

Having laid out the econometric model and data, we turn to the estimation stage. Nevertheless, we would emphasize that the empirical model description is illustrative and does not try to incorporate all the technical elements that can be found in the literature on the subjects that are addressed.

3.5   Regime identification

Finally, we turn our attention to the regime identification scheme. How does one identify periods with breakdown in the pass-through? This question requires thorough deliberation. In a single equation framework, where lending rates are explained through policy rates, one may order the states by imposing that the lower regression coefficient of the policy rate is attributed to the first state. In a VAR framework, regime ordering is not as straightforward. Moreover, we model the lending rate spread across two countries instead of the lending rates per se. In this setup, if monetary policy transmission has become heterogeneous, unexpected movements of the policy rate should affect the two counties differently, whereas if both lending rates react in a similar manner, one should not observe any difference in the spread. Hence, we propose three different identification strategies: (i) impulse response (IR) identification; (ii) identification via the Markov-switching constant; (iii) identification via the Markov-switching conditional mean.

The first identification scheme orders the regimes by calculating the impulse responses of the lending rate spread to a shock in the policy rate at each iteration and imposing the "stronger" IR as the second regime. We define "stronger" by calculating the cumulative response for twelve months ahead.[10]

The second strategy allocates the regimes according to the size of the constant in the lending rate spread equation - the higher constant is allocated to the second regime. The economic intuition behind this is that if the homogeneity across countries has changed, this might be

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[10] For robustness we also consider 9 to 18 months ahead, and the findings remain unchanged.

reflected in a level shift of the spread. Whether policy transmission has changed will then be evident by comparing the respective impulse responses.

In the third strategy we calculate the conditional mean of the lending rate spread at each iteration and allocate the higher of the two to the second regime. The rationale is similar to the second regime identification strategy — a level shift of the lending rate is an indication of heterogeneous transmission of monetary policy. The difference to the regime identification strategy above is that using the conditional mean controls for additional information through the other explanatory variables.

Note that neither of these strategies imposes any regimes ex-ante. They separate the data based on parameter mean values, which does not ensure that the posterior distributions do not overlap. Simply put, an ordering by the cumulative IRFs, for example, does not guarantee that the difference between the impulse responses will be statistically significant. Therefore, we do not assume a priori that there has been any change in the pass-through. This can be examined in our model only after we plot the actual impulse responses ex-post.

4   Estimation Results

In this section we present the estimation results for Italy, Spain, Ireland, and Portugal individually, as different risk assessments across countries may give rise to potentially different movements of the interest rate spreads. For example, even when the spreads of all countries respond to the same set of economic news, e.g. about macroeconomic data and/or monetary policy, the spreads in some countries may react more strongly when there are concerns over the pace and sustainability of reforms. At the same time, different countries may be more or less exposed to global factors when cross-border flows differ across countries. For each state we will first look at the transmission of monetary policy to the lending rates via impulse response functions, examine the regime probabilities and inspect the trigger variables. For each country we will focus mostly on a representative set of the significant trigger variables out of the full list given in online Appendix B.

4.1   Italy

The top panel of Figure 3 presents the estimated regime. Following Hamilton (1989), we interpret a value below 0.5 as the economy being in the first regime, and above 0.5 as a realisation of the second regime. In the bottom panel we analyse the state-contingent impulse responses of the lending rate spread.[11] We normalise the EONIA shock across the states. If the

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[11] The full set of impulse responses is available in the online appendix.

monetary policy transmission is homogenous across countries, the lending rates in different countries should not react differently to a monetary policy shock.

Italy: Estimated probability of the second regime and impulse-response functions   Figure 3

Estimated probability of the second regime[1]

State contingent impulse-response functions of the lending rate spread (r) to:

Shock in EONIA   Shock in g   Shock in r
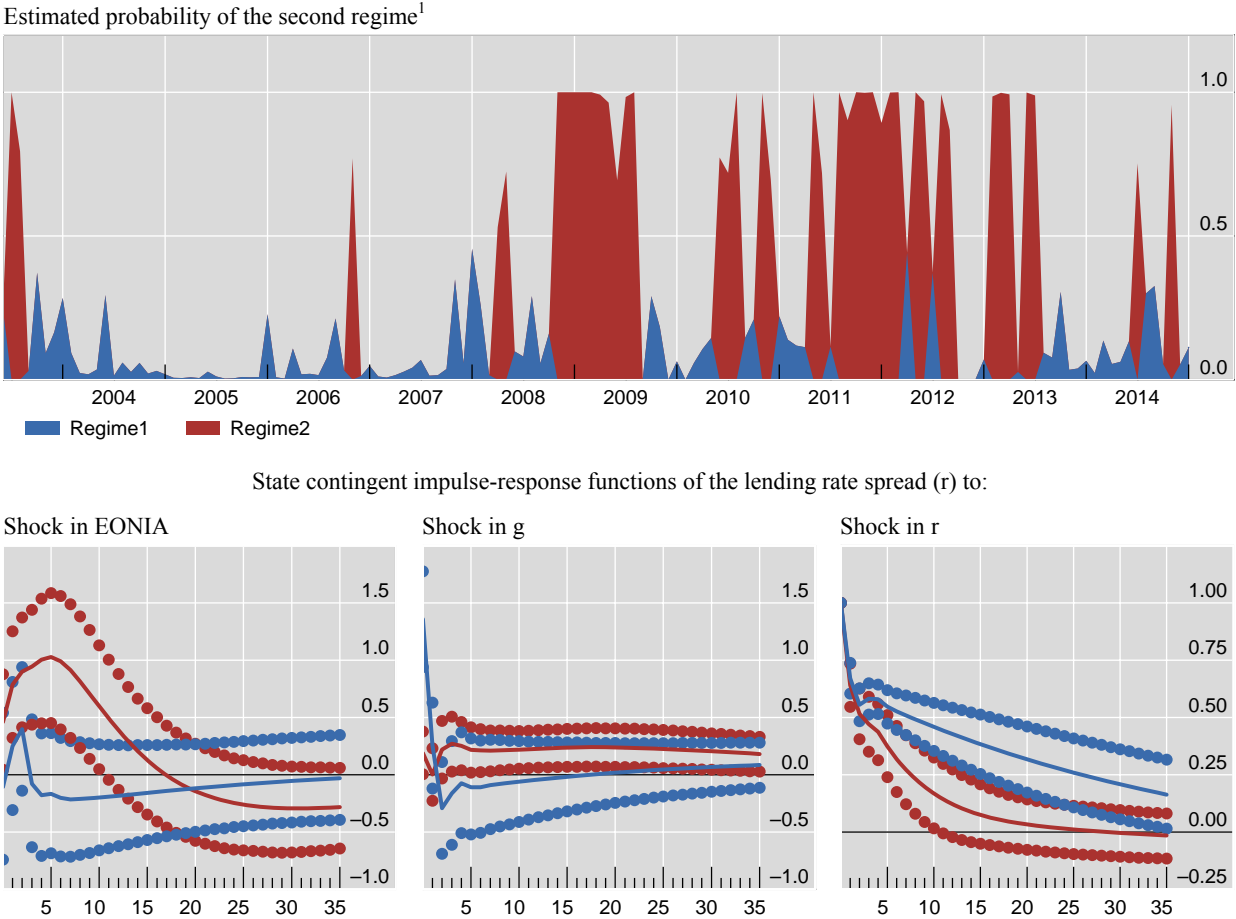
[1] A probability below 0.5 indicates a realisation of the first state and a value above 0.5 a realisation of the second state (top panel).

Source: Authors' calculations.

For Italy, however, this is not the case. A 100 basis point increase in EONIA leads to a significant opening of the lending rate spread in the second regime in contrast to the first, which indicates that the lending rates rise higher than in Germany.[12] Confusing as the estimates might appear, they have a clear economic interpretation – in regime one, market participants behave as if they are in a comfort zone and do not feel compelled or encouraged to pull the lending rates further away from the German rates. However, in the second regime, market participants anticipate a "dark corner" and act to increase the lending rate spreads vis-à-vis

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[12] These results are amplified by the presence of the zero lower bound. With EONIA being flat after the middle of 2012, the response of the spread is characterized by the persistence of the policy rate. This can be observed in the reaction of EONIA to a shock in EONIA for the second regime, and also in the residuals for the policy rate, which are plotted in Figure C.1 in the online appendix. To deal with this problem we also explore using shadow rate estimates for the euro area instead of EONIA, which are discussed in detail in section 4.5.

Germany. The second regime was prevalent during the outbreak of the financial crisis 2008 – 2009, between the months of August 2011 and 2012, and throughout the first half of 2013 — both associated with the euro area sovereign debt crisis. Rising fiscal imbalances and weak demand took a toll on Italy, with the crisis escalating in the autumn of 2011, leading to political turmoil with government bond yields increasing to an all-time high. Following a political change, Italian bond yields stabilized for a short time, but in the beginning of 2013 fears grew again. The economy started to recover slightly in 2013, with a major contributor being an improvement in the current account deficit, which turned positive in 2014.

What contributed to the regime shifts? We examine the trigger variables that are significantly different from zero in the probit equation. A positive coefficient decreases the probability of staying in the first regime and increases the switching probability to the second regime. We plot a representative set of trigger variables in Figure 4. The bottom panel shows the transition probability for the first state.[13] One of the main problems of the Italian economy has been a fiscal burden. Lower tax income and weak demand have put a large strain on government finances. A rising nominal debt-to-GDP ratio and worsening net foreign asset position have been important developments, and are a natural choice for trigger variables. The debt-to-GDP ratio was relatively stable through 2011, but jumped in 2012-14, reaching 160% in 2014. The net foreign asset position fell to minus five percent of GDP in 2011. Both prove to be an important indicator with a positive probability for switching to the regime of impaired monetary policy transmission. Among global financial variables, both the VIX and the Economic Policy Uncertainty index are significant, while the MOVE index does not contain information regarding the regime switching. It is interesting that for Italy the MOVE index is not significant. The correlation between monthly percentage changes in the MOVE and VIX indices over the period February 1990 to May 2014 was 0.46, and over the period January 2004 to May 2014 the correlation was 0.51, suggesting that the two measures of bond market and equity market volatility can reflect different features of uncertainty. Monetary policy in the form of actual borrowing in the MROs and LTROs also did not alleviate the problems, with both variables not influencing the switching probabilities.

The unconventional monetary policy announcements of the ECB have had a temporary positive effect. The dummy variables for the LTRO announcements as well as the "whatever-it-takes" speech and the OMT announcements have strong negative coefficients. Their effects are evident in the transition probabilities — they contribute to the spikes in the middle of

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[13] Note that the transition probabilities are symmetric, as is evident from equation (5).

2009 and 2012. The model suggests that through the strong influence on the transition proba-bilities, the announcements played a major role in the actual regime switches in August 2009 and August 2012.[14]

Italy: Representative trigger variables and transition probabilities   Figure 4

Trigger variables[1]

Transition probability of staying in the first regime $[p(Z)]^2$
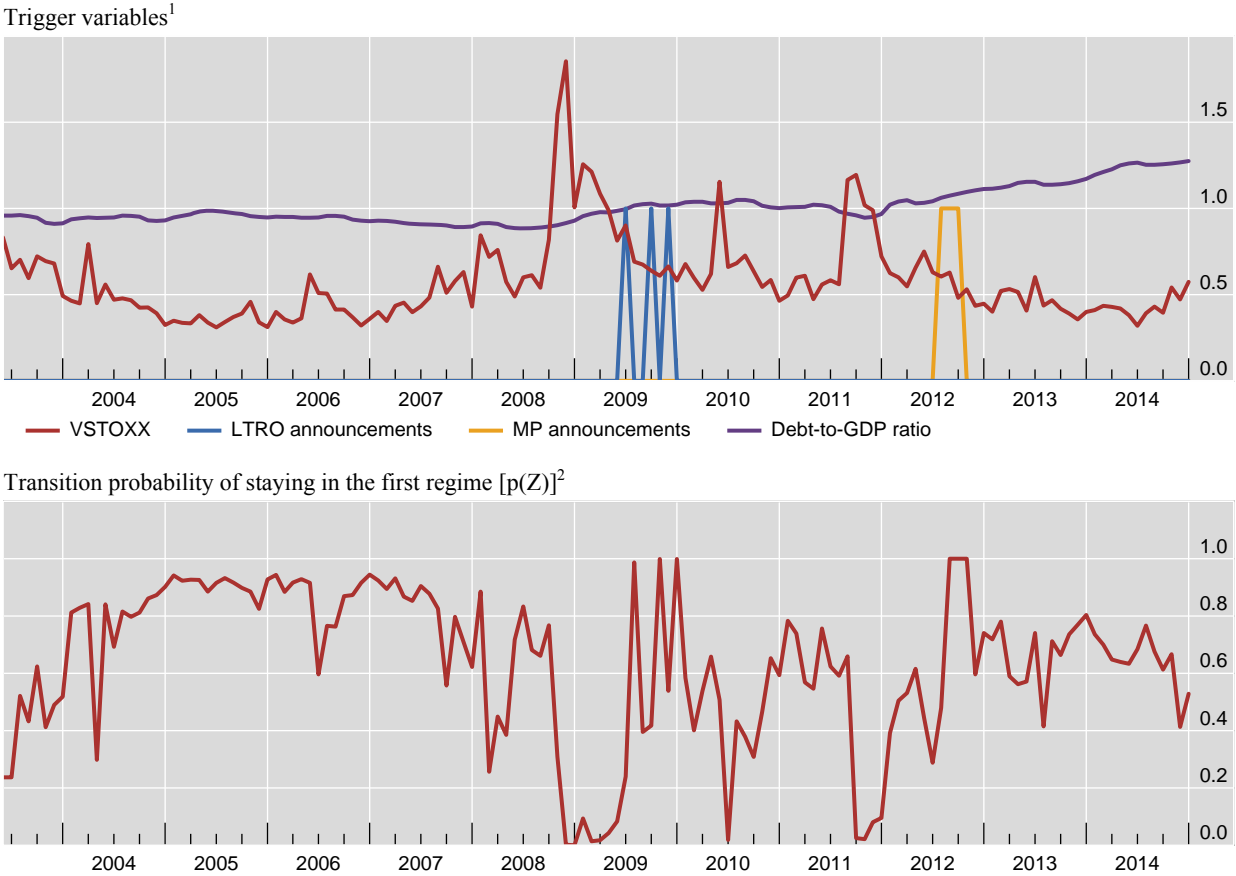
[1] VSTOXX and debt-to-GDP ratio have been rescaled for expositional clarity.   [2] A falling probability indicates a higher chance of switching to the second state and vice versa.

Source: Bloomberg, IMF IFS, authors' calculations.

Another potential matter is the issue of contagion — spillover effects of the sovereign debt crisis across many euro area countries were a major concern for the common monetary policy. Anecdotal evidence suggests that the increasing lending rate spreads originated in certain countries before spreading to further countries. To model this contagion, we estimate the model with lagged lending spreads of Spain, Ireland and Portugal. In Italy market sentiment towards the development of lending rates of other debt-ridden countries influenced domestic lending rates adversely, with Spain being the most important contributor. An interesting point about this finding is that the inverse is not true, as there were no signs of contagion effects

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[14] We present the distributions of the estimated parameters from the probit model in the online appendix.

from Italy to Spain. The reasons are related to the specifics of the Spanish economy (see below).

Although the time series specifications deal with nonlinearity and heterogeneity, they do not allow for cross-sectional dependence. By way of qualification, it must therefore be conceded that we need to be cautious when interpreting these results. There may be spillover effects from one country to another magnifying at times of financial crisis, exposures to common shocks, and the stance of the global financial cycle that could invalidate the cross-sectional independence assumption. These global factors are mostly unobserved. Furthermore, we are aware that the forecasting results of MS models is mixed, and that there are limitations to the use of switching models for forecasting. Although MS models fit well in-sample, they do not necessarily generate superior out-of-sample forecasts [Ferrara et al. (2012)].

4.2   Spain

The model for Spain exhibits both remarkable similarities and notable differences compared to that for Italy. On the surface the impulse response estimates seem equivalent, while the realised states have a higher persistence, with the second regime highly dominant after September 2008 (see Figure 5). This suggests a longer duration of the pass-through breakdown in Spain. The impulse response of the lending rate spread exhibits similar dynamics to the Italian one. However, the drivers of the endogenous transition probabilities uncover stark heterogeneity across both countries.

The policy announcements by the ECB did not strengthen the monetary policy transmission or alleviate the rising spread between the Spanish and German lending rates.[15] The main drivers behind the persistence of the second state seem to have been problems in the Spanish financial sector. After the near collapse of several banks, the Spanish central bank requested funds from the European Financial Stability Facility in June 2012.[16] This is reflected in the model by a negative significant coefficient of the Spanish bank stocks indicator. A rising index implies higher valued banks, and a negative coefficient affects positively the probability of switching from the second to the first regime. Hence, a banking crisis reflected in falling bank stock prices would lengthen the state of impaired pass-through.

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[15] What else can the ECB do to fix the wedge in the relationship between policy and lending rates hampering growth in the euro area periphery? Recently the ECB has unveiled a targeted offer of four-year loans designed to encourage banks to lend more to small- and medium-sized enterprises. To take advantage of the facility, which is available at a cheap fixed rate, banks must sign up to commitments to business lending, similar in design to the Bank of England's Funding for Lending Scheme [ECB (2014)].

[16] For an overview of the distress in the financial sector, see International Monetary Fund (2013).
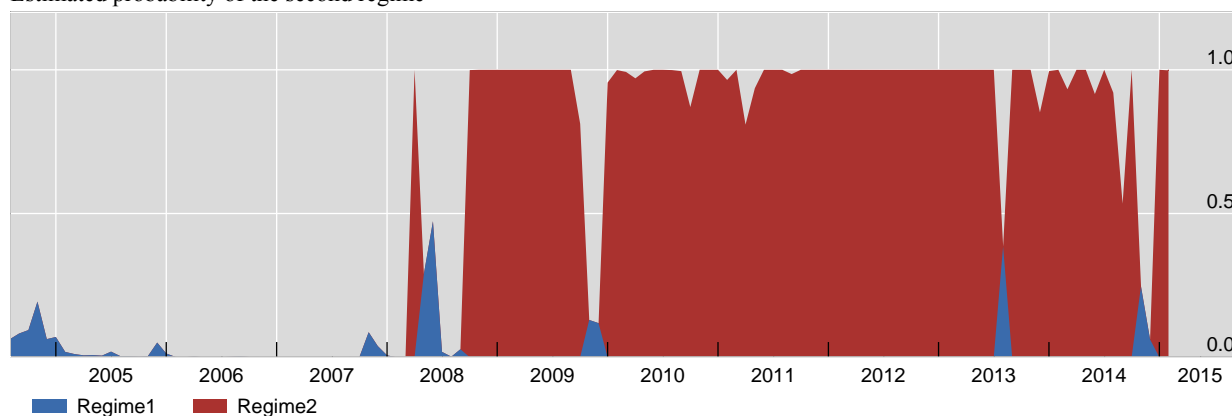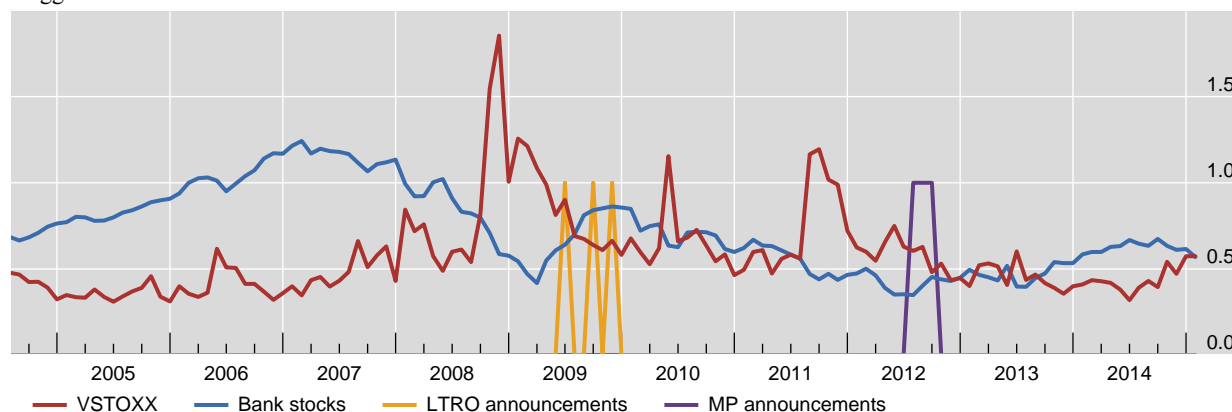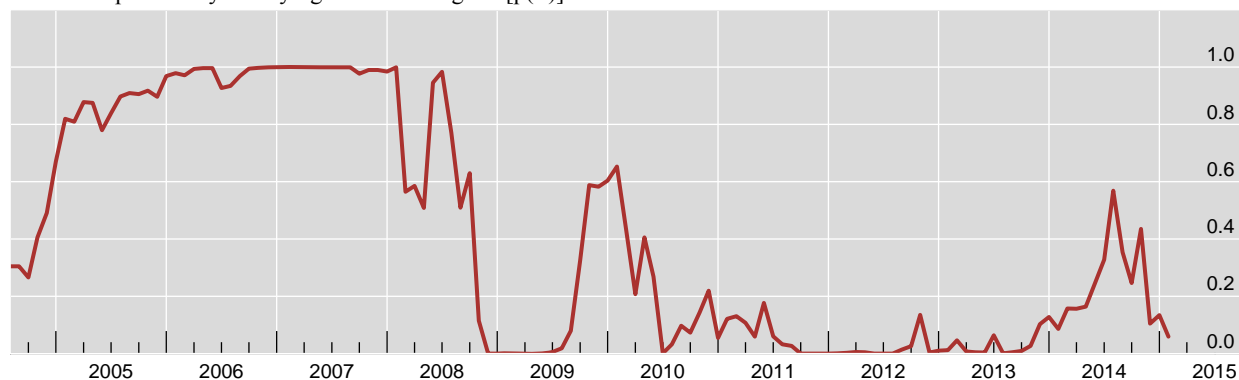
Spain: Estimated probability of the second regime and impulse-response functions   Figure 5

Estimated probability of the second regime[1]

Regime1   Regime2

State contingent impulse-response functions of the lending rate spread (r) to:

Shock in EONIA   Shock in g   Shock in r
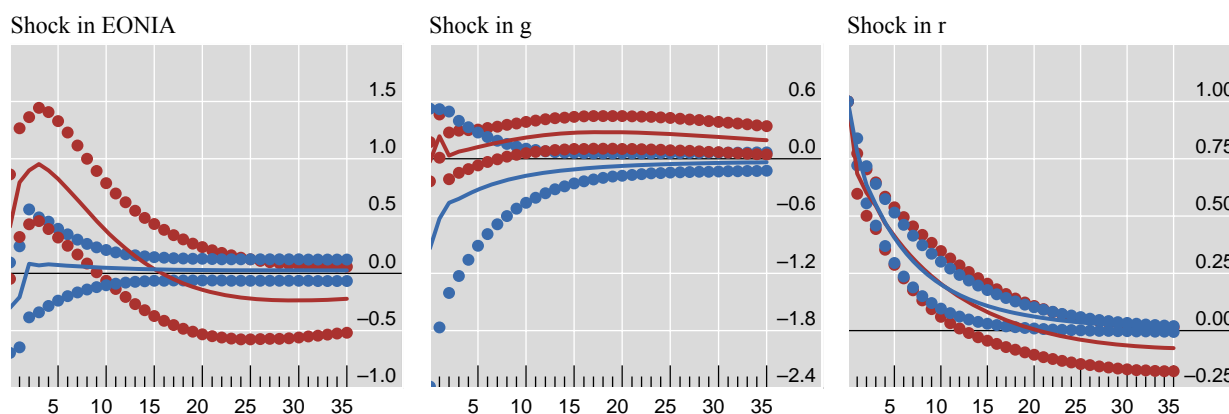
[1] A probability below 0.5 indicates a realisation of the first state and value above 0.5 a realisation of the second state (top panel).

Source: Authors' calculations.

Moreover, the main refinancing operations of the ECB do not appear to have alleviated the problem either. The notion that the problems in Spain were coming from within the country are strengthened by the fact that the lending rate spreads of other countries are not statistically significant, implying no contagion effects.

In terms of estimation, the nominal debt-to-GDP ratio turned out to be problematic, as the model did not exhibit convergence using the Gibbs sampler even with long chains of 100 000 draws. On the other hand, the issue of the zero lower bound does not seem to be of high importance — the EONIA residuals pass all normality tests (see online Appendix C).

Spain: Representative trigger variables and transition probabilities   Figure 6

Trigger variables[1]

— VSTOXX   — Bank stocks   — LTRO announcements   — MP announcements

Transition probability of staying in the first regime [p(Z)][2]

[1] VSTOXX and bank stocks have been rescaled for expositional clarity.   [2] A falling probability indicates a higher chance of switching to the second state and vice versa.

Source: Bloomberg, authors' calculations.

4.3 Ireland

Next we turn our attention to Ireland, where a breakdown in interest rate pass-through similar to that of Italy and Spain is also identified. Figure 7 shows that, on average, the reaction of the lending rate spread is similar, and there is a clear overreaction of the lending spreads to a tightening in the short rate. The realisation and persistence of the second regime are similar to those in Italy during the outbreak of the financial crisis in 2008 and associated with the euro area sovereign debt crisis in 2010, 2011, and 2012, before the monetary policy transmission returns to normal in 2013.

Notably no global variables contain information regarding regime switching. Neither VSTOXX, nor the MOVE index or the European policy uncertainty index have any predictive power regarding the regimes. This points to somewhat different financial conditions in Ireland.

Estimated probability of the second regime[1]
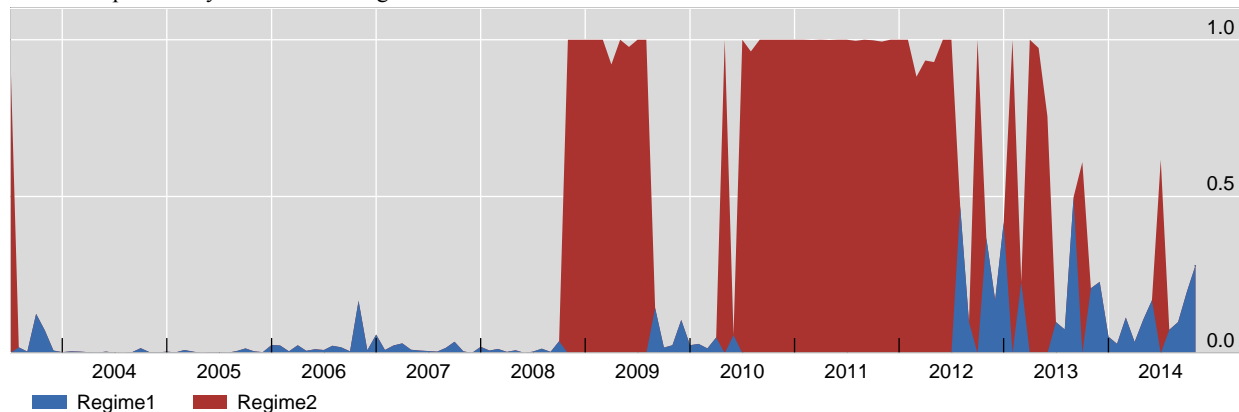
State contingent impulse-response functions of the lending rate spread (r) to:

Shock in EONIA   Shock in g   Shock in r
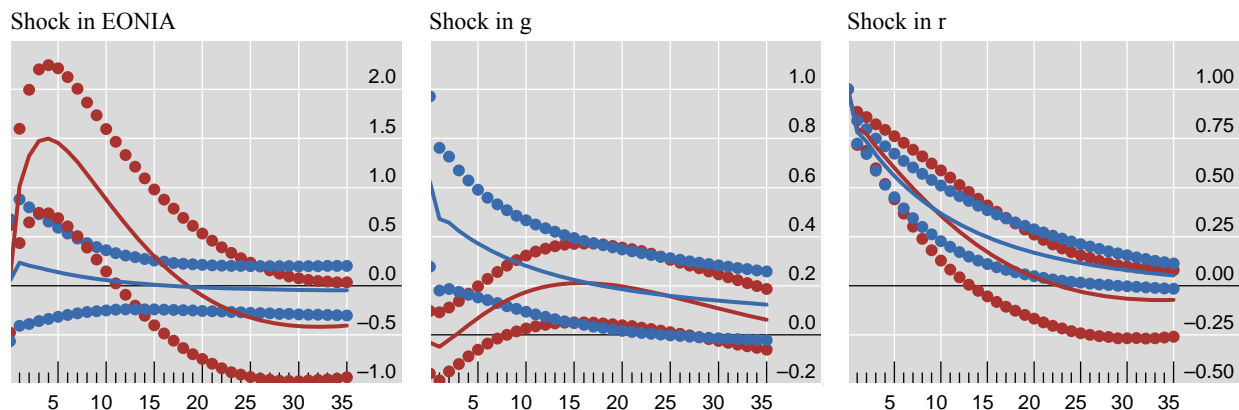
[1] A probability below 0.5 indicates a realisation of the first state and value above 0.5 a realisation of the second state (top panel).
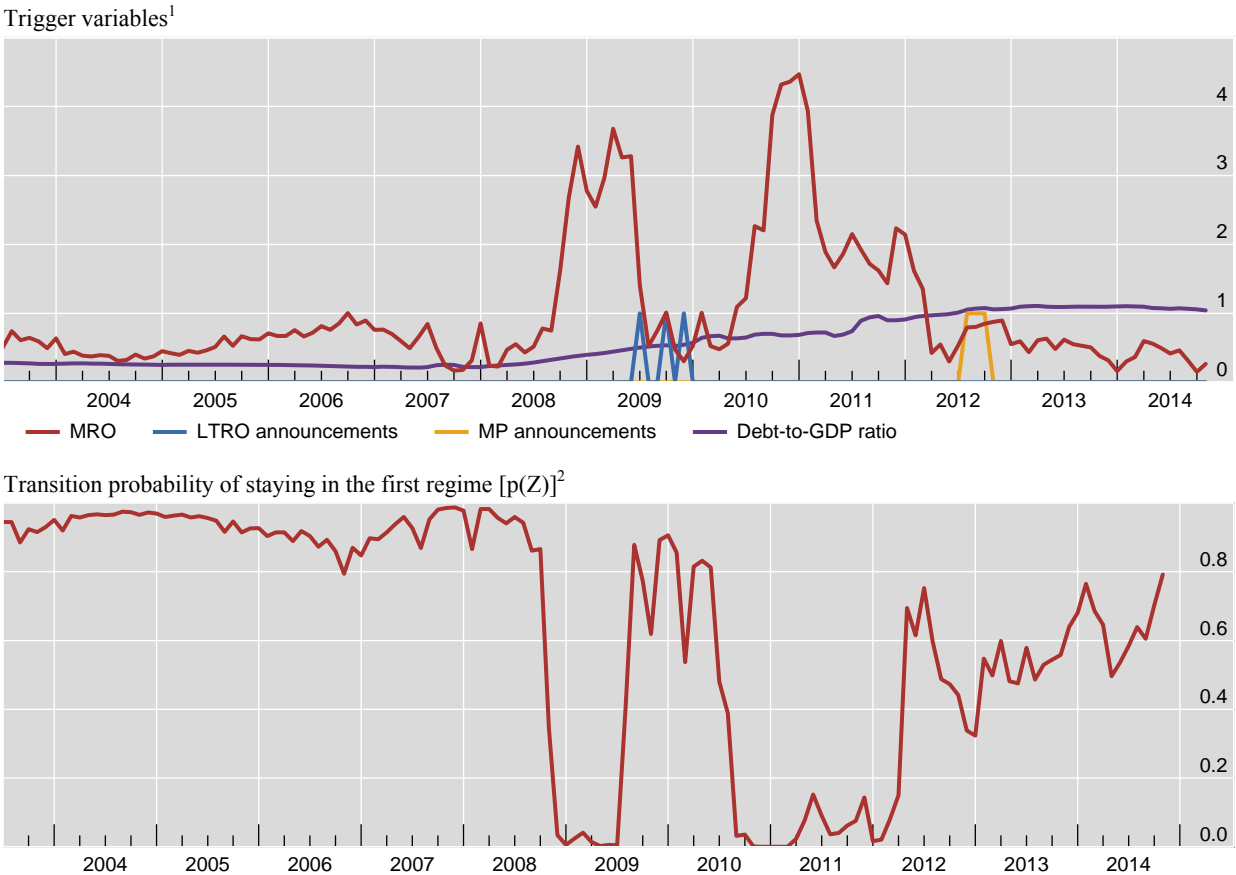
Source: Authors' calculations.

Similar to Italy, the country's debt burden has played an important role in the impairment of the monetary policy transmission – the coefficient is significant and positive. In contrast to Spain, the banking indicators are not informative, from which it can be inferred that the state of the banking sector was not the source of the breakdown in pass-through. This result could reflect that while the banking crisis in Ireland had major effects on the Irish economy in 2008 and 2009, there was a subsequent recovery in the banking sector, and further problems were mostly associated with the government deficit. For example, the bailout programme from 2008 expired in 2010 and was not renewed. Furthermore, in 2011 the Irish finance minister stated that the bank recapitalisation is fully covered by existing funds and loans are not needed,[17] and Ireland exited the bailout programme in 2013 successfully. The ECB's monetary policy announcements did not play a significant role in the return to the first state of normal

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[17] See http://debates.oireachtas.ie/dail/2012/06/06/00097.asp (accessed on 12 October 2015).

transmission, as they are insignificant. Hence, they did not contribute to the return to the first state in the middle of 2012.

The other significant trigger variable is the volume of main refinancing operations of the ECB, which enters with a positive coefficient, implying that an increase in these operations indicates a rising probability of transitioning to the second regime of heterogeneous pass-through. This might seem counterintuitive, but the MROs can be seen as an indicator for the state of the economy; in a "bad state" the ECB provides more liquidity assistance, and as the economy recovers the volumes decrease. If one uses CDS spreads between Ireland and Germany instead of the MROs, the results are similar, since the variables exhibit high positive correlation (see Table E.3 in online Appendix E).

Ireland: Representative trigger variables and transition probabilities   Figure 8

Trigger variables[1]

— MRO   — LTRO announcements   — MP announcements   — Debt-to-GDP ratio

Transition probability of staying in the first regime [p(Z)][2]

[1] MRO and debt-to-GDP ratio have been rescaled for expositional clarity.   [2] A falling probability indicates a higher chance of switching to the second state and vice versa.

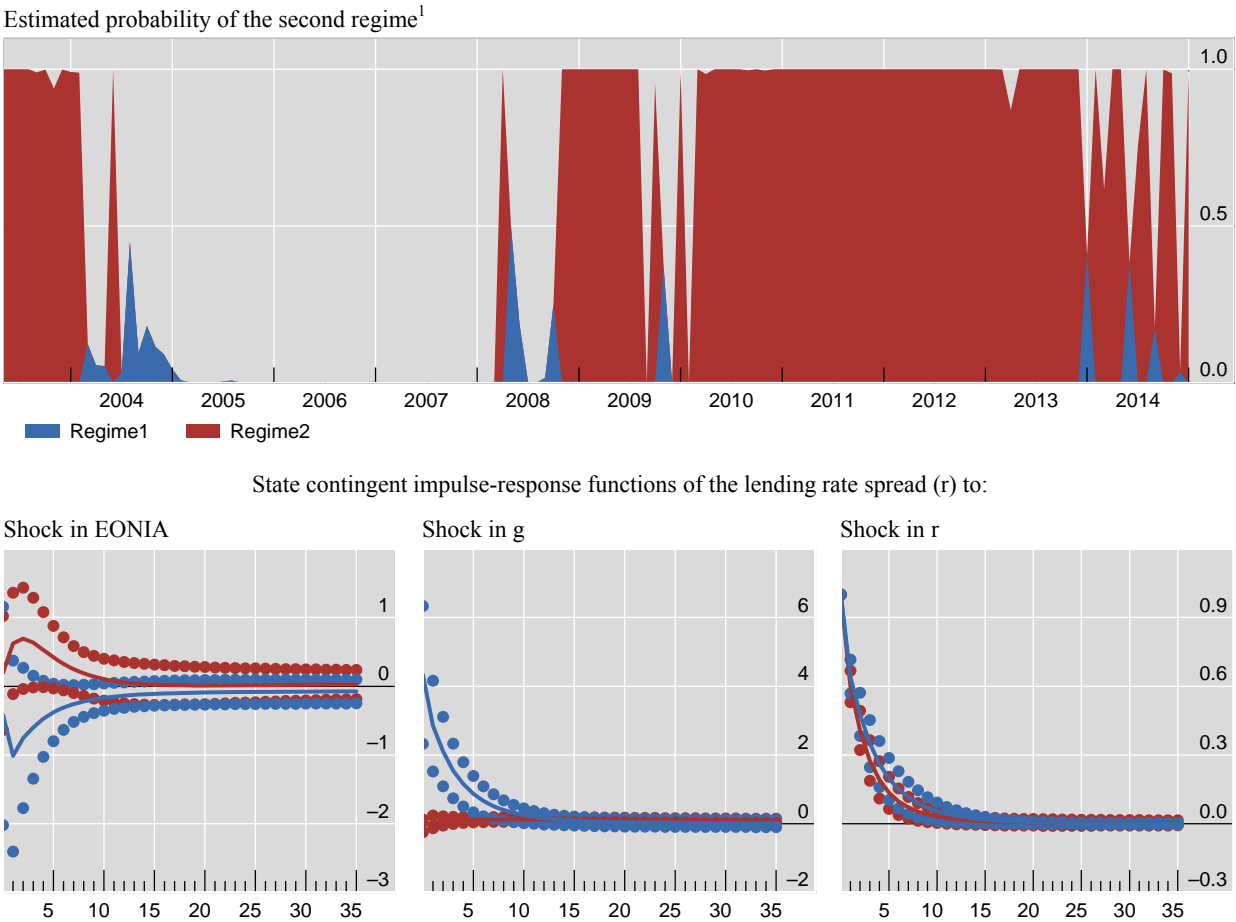Source: ECB, IMF IFS, authors' calculations.

4.4   Portugal

The final country to be examined is Portugal. We do not find a change in the interest rate pass-through. There was no setup in which the Gibbs sampler managed to identify different

reactions of the lending rate spread to an unexpected shock in EONIA. Most impulse respons-es also show no significant interaction across the variables, as is evident in Figure 9. Since the model does not identify two distinct regimes, the estimated states are arbitrary, as they do not carry dissimilar information. The realised states in the top panel of Figure 9 were not robust to the prior specification, in contrast to all the other countries. In all cases the estimated impulse responses of the lending rate spread turned out to be insignificant, irrespective of the regime estimation.

Portugal: Estimated probability of the second regime and impulse-response functions   Figure 9

Estimated probability of the second regime[1]

State contingent impulse-response functions of the lending rate spread (r) to:

Shock in EONIA   Shock in g   Shock in r

[1] A probability below 0.5 indicates a realisation of the first state and value above 0.5 a realisation of the second state (top panel).

Source: Authors' calculations.

Notably the lending rates of Portugal exhibit high volatility (see Figure 2). Surprisingly, they fail to pass a seasonality test, which is not expected from long-term interest rates, and clashes with standard economic intuition. This seasonality might be related to the seasonality to which the construction sector is exposed. Coupled with a flat policy rate, the residuals of the model fail to pass all normality tests (see online Appendix C). Due to these problems, and

specifically the lack of regime identification, we refrain from examining any potential trigger variables.

So far we have presented the baseline results, where EONIA is used as a proxy for the policy rate. This presents one potential challenge — the existence of the zero lower bound. The literature provides other proxies for the policy rate, and we will explore them in detail in the next section.

4.5 Dealing with the zero lower bound

Recently, a number of researchers have used shadow rate models to characterize the term structure of interest rates or quantify the stance of monetary policy [Wu and Xia (2014), Krippner (2014), Pericoli and Taboga (2015)].[18] The shadow short rate metric is a measure for the stance of monetary policy in a zero lower bound environment. The fictitious shadow short rate is a tool to summarize the joint impact of conventional and unconventional monetary policy in a parsimonious manner. Since different shadow rates have been proposed, we present three different estimates in Figure 10. For more details on these shadow short rate estimates, see online appendix H.

Comparison between alternative shadow short rates (SSR)[1]

In per cent   Figure 10
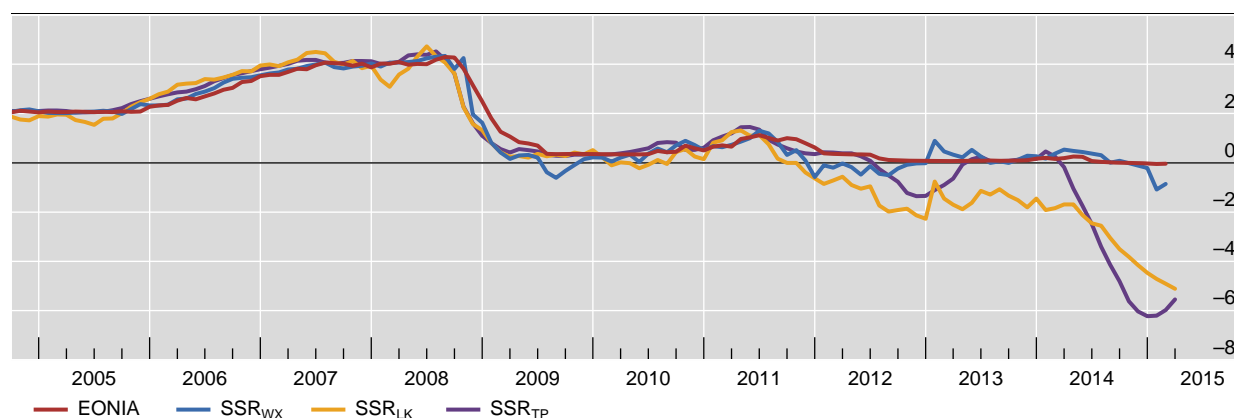
[1] SSR$_{WX}$ was taken from Wu and Xia (2014), SSR$_{LK}$ from Krippner (2014) and SSR$_{TP}$ from Pericoli and Taboga (2015). The latter was interpolated from quarterly data with a quadratic method (match average).

Sources: Wu and Xia (2014), Krippner (2014), Pericoli and Taboga (2015), authors' calculations.
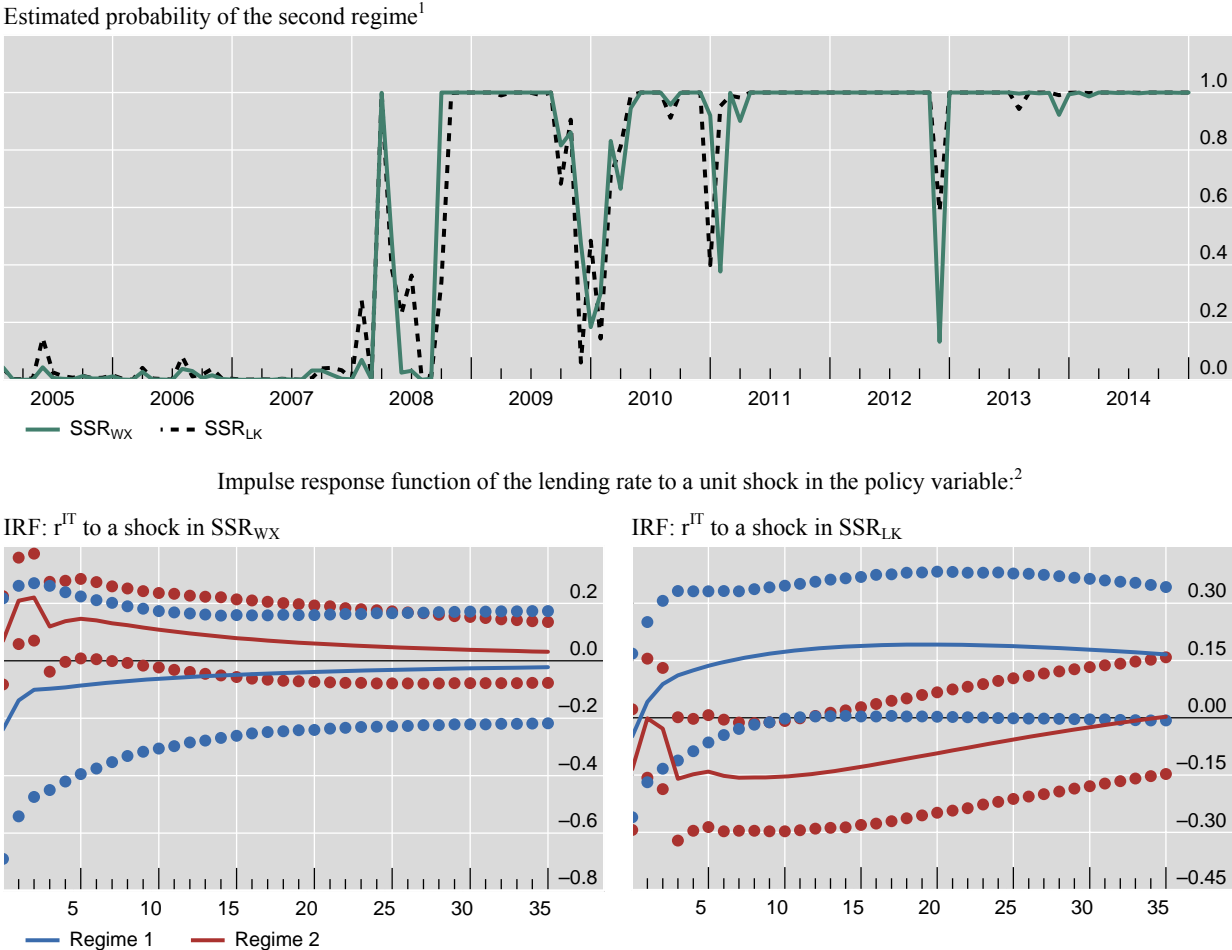
In the following figures, we plot the estimated realised regimes for the shadow rates, and the impulse responses of the lending rate spread to a shock to the policy rate, for each country. Note that we do not estimate the models with the rate of Pericoli and Taboga (2015) due to the drawback that it is only available at a quarterly frequency.

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[18] See also Lombardi and Zhu (2014) for another approach to estimate shadow short rates for the United States.

The main findings are that under Wu and Xia's estimates our results remain qualitatively unchanged for all countries. Quantitatively the responses of the lending rate spread to a 100 basis point increase in the policy rate are smaller, as evident from the figures below – about a quarter of the estimated responses with the EONIA rate in the second state, while in the first regime there is no significant effect, suggesting that monetary policy shocks affected all countries equally. We still do not find distinct states with Portuguese data. One can conclude that using EONIA as a proxy for the policy rate amplifies the results, but that the results are not driven by the zero lower bound.

Italy: Alternative model estimates for each shadow rate as the policy variable   Figure 11

Estimated probability of the second regime[1]

— SSR$_{WX}$ · - · SSR$_{LK}$

Impulse response function of the lending rate to a unit shock in the policy variable:[2]

IRF: r^{IT} to a shock in SSR$_{WX}$ IRF: r^{IT} to a shock in SSR$_{LK}$

— Regime 1   — Regime 2

[1] SSR$_{WX}$ was taken from Wu and Xia (2014) and SSR$_{LK}$ from Krippner (2014).   [2] For the models to be comparable, the estimation sample has been constrained to the shortest data series and the lag length has been kept constant across the SSR models. The model setup is identical to the EONIA scenario.
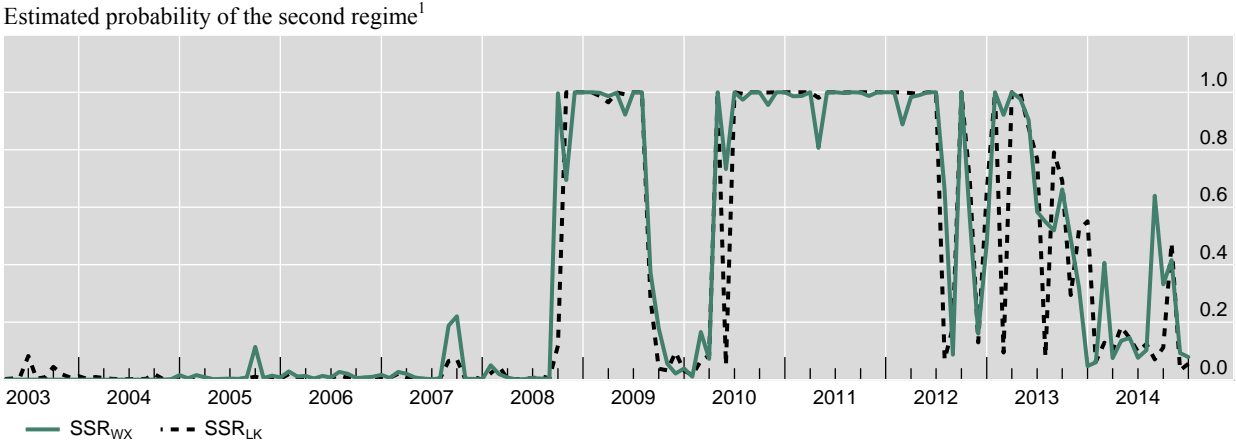
Source: Authors' calculations.

By contrast, if one uses the shadow short rate of Krippner (2014), the estimates paint a different picture for Italy and Spain, while they produce similar responses to Wu and Xia's short rate and EONIA for Portugal and Ireland. For Italy (Figure 11, right panel) the model does not

identify different responses of the lending rate spread initially. However, after three months the spread does react by becoming negative following a tightening of the short rate, which is at odds with the other two rates. This feature is also evident in Spain (Figure 12, right panel), where it is more pronounced – after the fourth month the second state is associated with a negative spread, while the first state is associated with a positive spread following a policy rate shock.

Spain: Alternative model estimates for each shadow rate as the policy variable | Figure 12
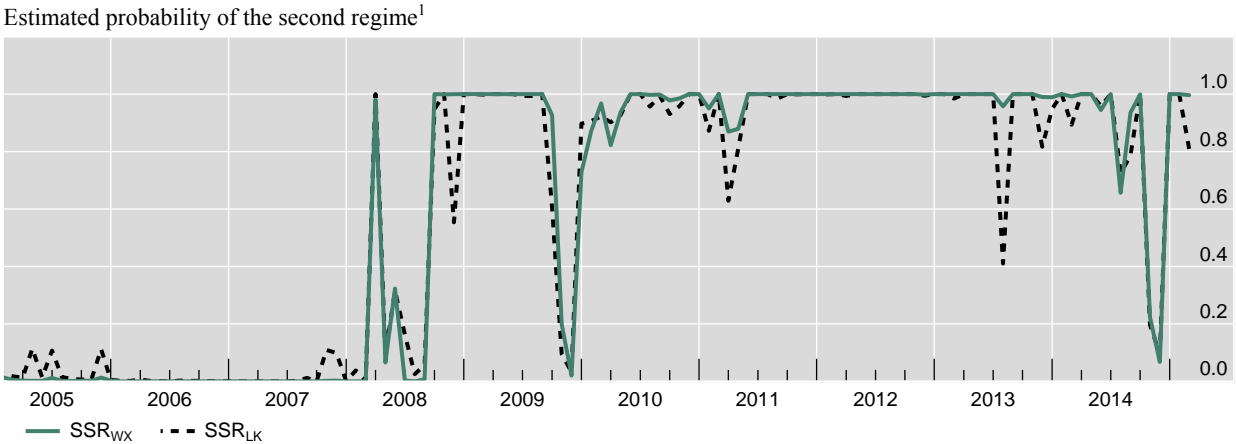
Estimated probability of the second regime[1]

SSR_WX   SSR_LK

Impulse response function of the lending rate to a unit shock in the policy variable:[2]

IRF: r^{ES} to a shock in SSR_WX

IRF: r^{ES} to a shock in SSR_LK

Regime 1   Regime 2
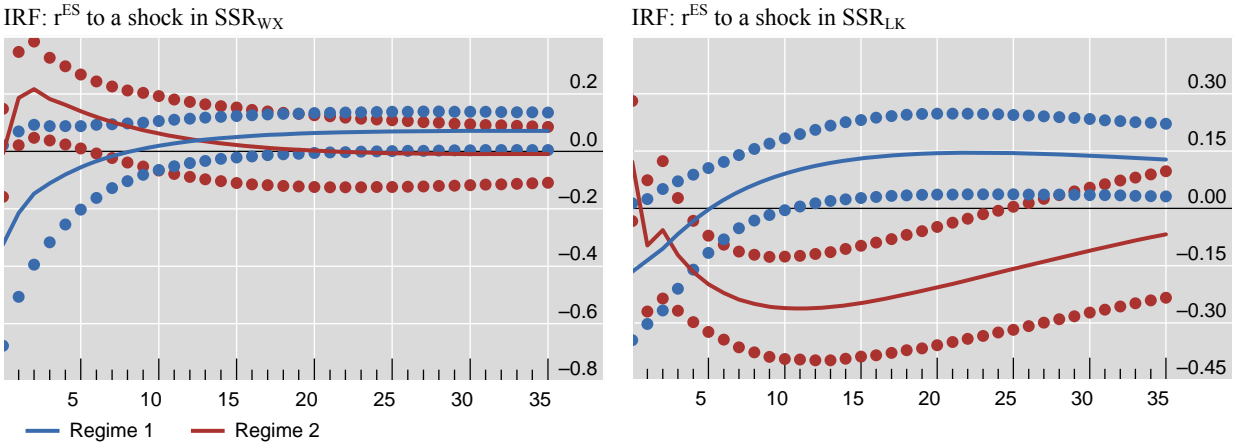
[1] SSR_WX was taken from Wu and Xia (2014) and SSR_LK from Krippner (2014). [2] For the models to be comparable, the estimation sample has been constrained to the shortest data series and the lag length has been kept constant across the SSR models. The model setup is identical to the EONIA scenario.

Source: Authors' calculations.

Comparing the results for Italy and Spain, it is fair to say that the estimation results with SSRs are less robust. One reason for this could be the sharp decrease of the shadow rate of Krippner (2014), which falls below zero in 2011 and never turns positive, with values reaching minus five percent. This coincides with the whole period of the lending rate spread being positive, hence a downward movement of the policy rate is associated with an upward movement of the
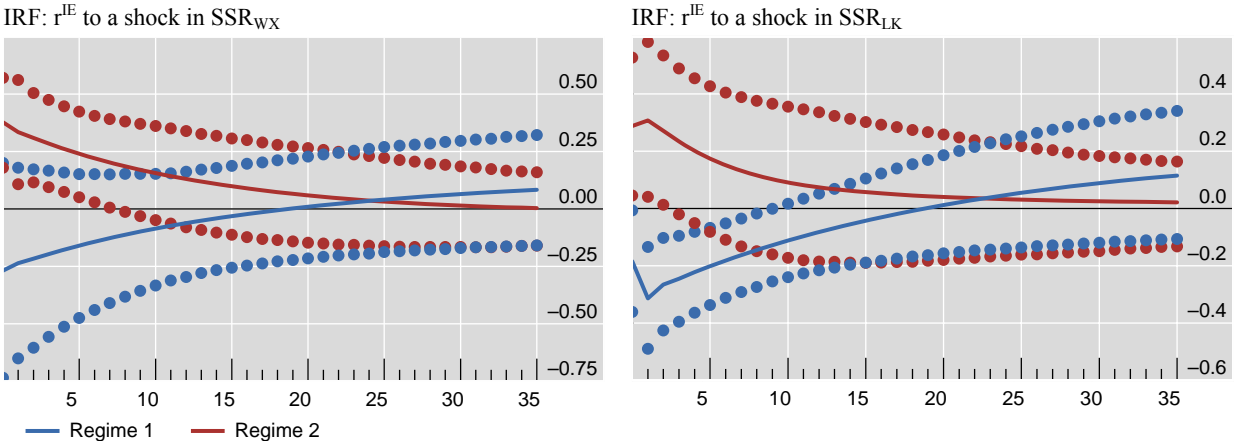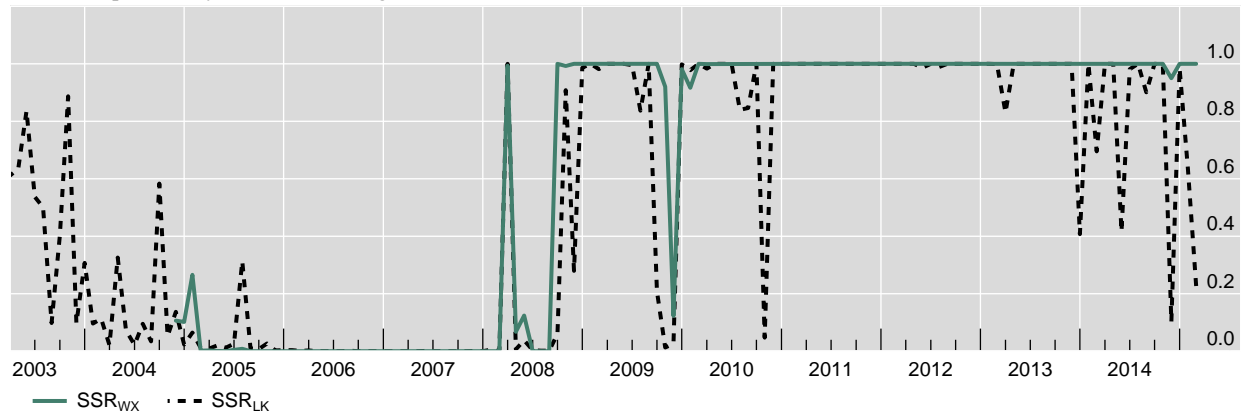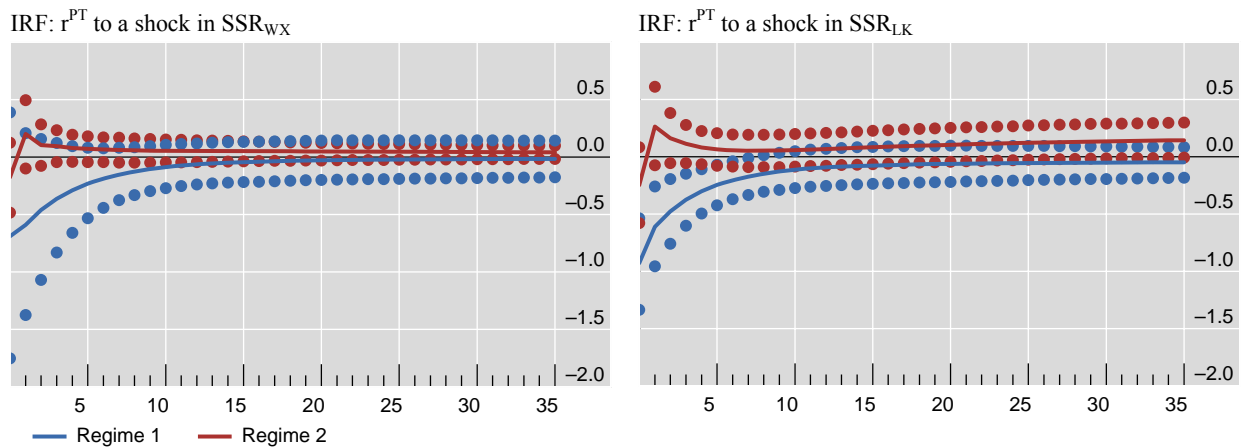
lending rate, which is exactly what we find – following a positive shock, the spread closes and following a negative shock the spread opens.

Ireland: Alternative model estimates for each shadow rate as the policy variable Figure 13

Estimated probability of the second regime[1]

Impulse response function of the lending rate to a unit shock in the policy variable:[2]

IRF: r^{IE} to a shock in SSR_{WX}

IRF: r^{IE} to a shock in SSR_{LK}

[1] SSR_{WX} was taken from Wu and Xia (2014) and SSR_{LK} from Krippner (2014). [2] For the models to be comparable, the estimation sample has been constrained to the shortest data series and the lag length has been kept constant across the SSR models. The model setup is identical to the EONIA scenario.

Source: Authors' calculations.

The results for Ireland and Portugal appear to be robust. They differ mostly in the size of the estimated confidence intervals for the responses, but nevertheless carry the same economic interpretation. The monetary policy transmission has recovered for Ireland following the middle of 2013, and we do not identify any breakdown in pass-through during the financial crisis for Portugal. It is notable that in all models the estimated realization of the first and second regime is highly similar, which is encouraging.

Portugal: Alternative model estimates for each shadow rate as the policy variable   Figure 14

Estimated probability of the second regime[1]

SSR$_{WX}$ ---- SSR$_{LK}$

Impulse response function of the lending rate to a unit shock in the policy variable:[2]

IRF: r^{PT} to a shock in SSR$_{WX}$ IRF: r^{PT} to a shock in SSR$_{LK}$

Regime 1   Regime 2

[1] SSR$_{WX}$ was taken from Wu and Xia (2014) and SSR$_{LK}$ from Krippner (2014). [2] For the models to be comparable, the estimation sample has been constrained to the shortest data series and the lag length has been kept constant across the SSR models. The model setup is identical to the EONIA scenario.

Source: Authors' calculations.

5   Concluding Remarks

The effect of monetary policy on lending rates is central in the policy debate on the design of optimal euro area policy. This topic has received renewed interest among economists and policymakers in the aftermath of the global financial crisis and the euro area sovereign debt crisis. Monetary policy and lending rates are endogenous variables, determined, possibly, by various economic shocks. Can any causal link between the two be established? How do monetary policy and lending rates interact? This paper strives to better understand the mechanisms by revisiting the question whether Italy, Spain, Ireland, and Portugal have experienced heterogeneity in the transmission of the common monetary policy, and investigate what the triggers of these changes were. We approach the question individually for each country through the lens of a Markov-switching VAR with endogenous transition probabilities.

By endogenising the transition probabilities between different regimes, we find that global risk factors have contributed to higher lending rates in Italy and Spain, problems in the banking sector help to explain the impairment in Spain, and fiscal problems and contagion effects have contributed in Italy and Ireland. We also find that the ECB's unconventional monetary policy announcements have had temporary positive effects in Italy. For Portugal we cannot identify any significant change in the interest rate pass-through, in contrast to the other countries.

Following the most recent debates in the literature, we address the potential pitfall of using EONIA as a proxy for the policy rate of the ECB by looking at alternative shadow short rate estimates as proxies for the policy rate. Shadow rate estimates derived from dynamic factor models provide a remedy for the issue of the zero lower bound, yet come at the cost that different models yield different shadow rate estimates. We conclude that the flatness of EONIA as the main policy rate amplifies the results, but does not alter the key findings.

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Technical appendix

For estimation we employ Bayesian methods and incorporate the priors following Banbura et al. (2010). This is achieved by augmenting the vectors of endogenous and exogenous variables through the following matrices:

$$Y_d = \begin{bmatrix} \Lambda \cdot \widehat{\Sigma}/\lambda \\ 0_{n(l-1)\times n} \\ \dots \dots \dots \dots \\ \Sigma \\ \dots \dots \dots \dots \\ 0_{1\times n} \\ \Lambda \cdot M/\tau \end{bmatrix}, \qquad X_d = \begin{bmatrix} J_l \otimes \Lambda \cdot \widehat{\Sigma}/\lambda & 0_{nl\times 1} \\ 0_{n\times nl} & 0_{n\times 1} \\ \dots \dots \dots \dots & \dots \dots \\ 0_{1\times nl} & \epsilon \\ 1 \otimes \Lambda \cdot M/\tau & 0_{n\times 1} \end{bmatrix}. \tag{T.1}$$

Here, $\widehat{\Sigma} = diag(\sigma_1, \dots, \sigma_n)$ is the estimated covariance matrix of the residuals from equation (7), which we weigh by a matrix $\Lambda = diag(\delta_1, \dots, \delta_n)$. The weights δ control how informative more recent lags are compared to older periods. Since our system is generally short, these parameters are not of crucial interest. $M = diag(\mu_1, \dots, \mu_n)$ are the average levels of the endogenous variables in y_t and $J_l = diag(1, \dots, l)$. The parameter λ is the overall tightness of the prior, which ranges from $[0, \infty]$, with 0 being a pure random walk and infinity the OLS estimates. ϵ denotes the prior on the constant. Furthermore, we incorporate Bayesian shrinkage by means of the hyperparameter τ. Finally, the operator " $\cdot$ " denotes elementwise multiplication.

For the choice of these parameters we follow Banbura et al. (2010) and set $\lambda = 0.1; \epsilon = \frac{1}{10000}; \tau = 10\lambda$, and μ_i to the mean of the y_i vector. Combining (8) with (T.1) leads to the following specification:

$$Y^* = X^* \widetilde{\beta} + U^*, \tag{T.2}$$

where $Y^* = [Y', Y_d^{*'}]', X^* = [X', X_d^{*'}]', U^* = [U', U_d']'$. Taking the OLS estimate of $\widetilde{\beta}$ in equation (T.2) as $\widetilde{\beta} = (X^{*'}X^*)^{-1}X^{*'}Y^*$ we impose an inverse Wishart prior on its variance

$$\widetilde{\Sigma} \sim iW\left(\widehat{\Sigma}, T^* + 2 + (1 + n*l)\right), \tag{T.3}$$

where T^* denotes the number of rows in Y^*. Therefore, the posterior distribution of interest becomes

$$vec(\widetilde{\beta})|\widetilde{\Sigma}, Y^* \sim N\left(vec\left(\widehat{\widetilde{\beta}}\right), \widetilde{\Sigma} \otimes (X^{*'}X^*)^{-1}\right). \tag{T.4}$$

Inference on this form of the MS-VAR is straightforward once the vector of realised states $S_T = [s_1, \dots, s_T]$ is known, as the model collapses to $n^s = 2$ linear Bayesian VARs. The vector of regimes may be obtained through the Hamilton filter. Letting $\mathbb{P}_T = [p_1(Z), \dots, p_T(Z)]$ and $\mathbb{Q}_T = [q_1(Z), \dots, q_T(Z)]'$, estimation is carried out via the Gibbs sampler in the following order of events. Given initial conditions for the parameters of interest $\{\widetilde{\boldsymbol{\beta}}_0, \widetilde{\boldsymbol{\Sigma}}_0, \boldsymbol{\Gamma}_0, \mathbb{P}_{T,0}, \mathbb{Q}_{T,0}\}$ and denoting an arbitrary iteration number by j we:

1. Draw $S_{T,j}$ using the Hamilton filter conditional on $\widetilde{\boldsymbol{\beta}}_{j-1}, \widetilde{\boldsymbol{\Sigma}}_{j-1}, \mathbb{P}_{T,j-1}, \mathbb{Q}_{T,-1j}$.

2. Draw $\widetilde{\boldsymbol{\beta}}_j$ conditional on $\widetilde{\boldsymbol{\Sigma}}_{j-1}$ and $S_{T,j}$, equation (T.4).

3. Estimate $\widehat{\boldsymbol{\Sigma}}$ and draw $\widetilde{\boldsymbol{\Sigma}}_j$ conditional on $\widetilde{\boldsymbol{\beta}}_j$, equation (T.3).

4. Estimate the probit model using $S_{T,j}$ and obtain $\boldsymbol{\Gamma}_j, \mathbb{P}_{T,j}, \mathbb{Q}_{T,j}$, equation (3).

5. Set $j = j + 1$.

We employ 50000 iterations and discard the first 35000 as a burn-in phase. In the online appendix we present the trace and recursive means plots to assess convergence in the spirit of An and Schorfheide (2007).

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