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March 2020

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This version: 20 March 2020

Abstract. The critical role of house prices for macroeconomic and financial stability is widely acknowledged since the global financial crisis. While house prices showed spectacular increases and even a bubble-like behaviour in the pre-crisis years, their fall thereafter was accompanied by deep recessions in many countries. Loose monetary conditions, such as the easy availability of credit, are often blamed to be fuelling such booms. In this paper, the link between credit and house prices is investigated for the euro area in a nonlinear model framework. This choice is motivated by the idea that the linkages between these two variables can be governed by a regime-switching behaviour. Threshold VAR (TVAR) models are estimated, which comprise real house price and credit developments, business and monetary conditions. Optimal breakpoints are determined via a grid search. The relationship between the variables is not stable. If output growth and interest rate changes serve as thresholds, two regimes can be distinguished. Conversely, if house prices and credit control the regime change, three regimes are more appropriate. Nonlinear impulse responses suggest that credit developments respond to house prices, while the reverse causality is less significant. Thus, the modest recovery of credit at the current edge can only be partially attributed to the recent acceleration of house prices in the euro area.

JEL: C34, E51, E52

Keywords: Threshold models, house prices and credit, regime switching

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1 Introduction

The critical role of house prices for macroeconomic and financial stability has been widely acknowledged since the global financial crisis (Bagliano and Morana, 2012). Even before the outbreak of the crisis, researchers emphasized the risk of booming house prices for the business cycle. While house prices showed spectacular increases and sometimes even a bubble-like behaviour in the pre-crisis years, their fall thereafter caused deep recessions in many countries. Loose monetary standards, such as easy credit conditions, low interest rates, high loan-to-value ratios and permissive lending approvals, have been often blamed to have triggered house price overvaluations prior to the crisis, most notably in countries with more liberal mortgage markets. Several studies have stressed the role of credit cycles as a major driver for the excessive developments in the real estate market, such as Kiyotaki and Moore (1997), Borio and Lowe (2004) and Wachter (2015), even more as almost 40 percent of total credit to the private sector is devoted to housing purchases. In most euro area countries, the financial conditions are heavily shaped by the availability of bank loans. However, the ties between house prices and credit might not be stable and can vary substantially over time. Against this background, this paper investigates the linkages between house price and credit developments in the context of a non-linear framework (whereby nonlinearities are captured by TVAR models), which represents a novelty compared to the existing literature.

The interlinkages between house prices and credit can materialise through various channels (see Hofmann, 2003; Gerlach and Peng, 2005; Fitzpatrick and Mc Quinn, 2007; Brissimis and Vlassopoulos, 2009; Oikarinen, 2009a and 2009b; Gimeno and Martinez-Carrascal, 2010; Duca, Muellbauer and Murphy, 2011; Avouyi-Dovi, Labonne and Lecat, 2014; Lindner, 2014; Piazzesi and Schneider, 2016), but the direction of causality is often difficult to underpin. First of all, credit developments can be transmitted into house prices via interest and mortgage rates: lower interest rates can lead to a rise in the demand for loans for house purchases and, subsequently, in a rise of residential property prices.[2] At the same time, the reverse causality also

[2] Favara and Imbs (2015) show that an exogenous expansion in mortgage credit has significant effects on house prices. Focusing on the US branching deregulations between 1994 and 2005, the authors argue that the banks operating in deregulating states experienced significantly higher deposit growth, and lower deposit costs. They also charged significantly lower rates, presumably because some of the cost savings were passed through to borrowers. Credit terms improved, more borrowing could be observed, and the demand for housing increased. In areas where housing supply is inelastic, the response of house prices was more pronounced, while it was muted in areas where housing supply is elastic. In addition, Aastveit et al. (2020) find that the decline in housing supply elasticities in the US has also played a role, with a stronger response of house prices to monetary policy in the most recent house price recovery.

holds: in an environment of low interest rates, housing becomes relatively more attractive as an investment opportunity. In turn, credit supply is also affected, as booming house prices improve the balance sheet of banks and following the increase in the banks' capitalisation, a larger amount of loans can be granted, which also tends to fuel housing booms.

Furthermore, the wealth and collateral channels may play a role: provided that the increase in house prices is not only short-lived, higher house prices would raise households' wealth, thus leading to a higher path of consumption over the lifetime (Carroll, Otsuka and Slacalek, 2011) and additional demand for credit, as real estates are often used as collateral for bank lending: homeowners can finance the additional expenditures through additional borrowing, as the higher collateral value increases their borrowing capacities. At the same time, higher property valuations make banks' assets less risky, as the increased value of the collateral pledged reduces the likelihood of defaults on existing loans while also their holdings of real estate assets increase in value, which may influence the risk taking capacity of banks and increase their willingness to expand their lending.[3]

Finally, it has also been argued that, if housing represents a high share in households' portfolio, increases in house prices can signal an increase in expected returns and, therefore, lead to a corresponding shift of resources. For example, Corradin, Fillat and Vergara-Alert (2014) find that accelerating house prices increased the likelihood of housing purchases and investment in the US. Overall, due to these complex interactions and self-reinforcing mechanisms, there is no unique direction of causality between house prices and credit and, therefore, the relationship between house prices and credit should be investigated by simultaneous models.[4]

Several authors have investigated the fundamental drivers of the house price dynamics in order to explain the significance of their imbalances, but no clear-cut conclusions have emerged so far. In their seminal study, Tsatsaronis and Zhu (2004) conclude that real house prices can be explained by consumer price inflation, the yield curve and bank credit, but national differences in mortgage markets also matter. In addition, other variables, such as disposable income, demographic shifts and tax incentives for home ow1nership may exert an impact on house prices in the long run. House prices also appear to be more sensitive to short-term in-

3 Part of the additional available credit may be used to purchase property, pushing up property prices even further, so that a self-reinforcing process may evolve.

4 The evidence from the literature regarding the nexus of housing prices and credit is quite extensive; however, the results tend to disagree about the direction of causality, whereby the discrepancies can be ascribed to different reasons, such as institutional differences among the countries, methodological approaches, sample size and data sets used in the studies.

terest rates in countries where floating mortgage rates are widely used. More aggressive lending practices seem to be associated to stronger feedbacks from house prices to bank credit. Goodhart and Hofmann (2008) report multidirectional linkages between house prices, monetary variables and the macroeconomy in advanced countries. The impact of shocks to money and credit aggregates are higher in periods of booming house prices. However, this effect turns out to be hardly significant as reflected in the large confidence bands of the impulse responses.

By applying a user cost approach to explain the development of house prices, Glaeser, Gottlieb and Gyourko (2010) conclude that the predicted impact of interest rates is much lower once several model extensions, such as mean-reverting behaviour of long-term interest rates, repayment conditions and credit-constraints of home buyers are included. Lower loan-rejection and higher loan-to-value ratios have only modest effects and, thus, cannot justify the observed hikes in US house prices. By contrast, Duca, Muellbauer and Johnson (2011), by augmenting the user cost indicator with a credit availability index based on loan-to-value ratios for first time home buyers, find that their new index has high explanatory power for house prices to rent ratios. Therefore, they conclude that low credit standards are a trigger for house price booms. Gattini and Hiebert (2010) emphasize a strong role of demand and supply shocks to explain the house price dynamics in the euro area. Real house prices are related to fundamental drivers in the long run, including real housing investment, real disposable income per capita and the real interest rate. By using a composite index incorporating stock and house price developments, Gerdesmeier, Reimers and Roffia (2010) report evidence that credit aggregates, long-term interest rates and the investment-to-GDP ratio can predict asset price booms and busts, whereby asset prices include both stock and house prices. Following Gerdesmeier, Lenarcic and Roffia (2014), user costs, demographics, unemployment, debt-to-income ratio, disposable income and the housing stock are crucial to explain house prices in the long run on the basis of a demand-inverted approach.

While the bulk of the empirical evidence is based on linear models, nonlinearities have been largely ignored. As an exception, Calza and Sousa (2006) conclude that credit shocks have larger effects on GDP in periods of tighter credit regimes, although the impact is not very strong. Himmelberg, Mayer and Sinai (2015) argue that house prices are more sensitive to monetary conditions in periods when nominal interest rates are low. Buyers might be keener to enter the housing market when prices increase, probably due to the fear that a further delay would result in even higher prices. If prices fall, market participants tend to be reluctant to buy residential properties or sell their homes due to loss aversion. If nonlinearities are substantial, linear models are not suited to capture these effects.

Against this background, this paper explores the relationship between credit and house prices by allowing for nonlinearities. The nonlinear model includes the linear specification as a nested case.. In case of stationary variables, threshold VAR (TVAR) models can be used to examine the interlinkages between those variables. Conversely, if the series are nonstationary and cointegrated, a threshold VEC (TVEC) model can be considered to be more appropriate. Both approaches can distinguish between different regimes. As a matter of principle, changes between the regimes occur once certain thresholds are crossed. Moreover, while the process within each regime is essentially linear, the coefficients are regime-specific. It is also important to note that the threshold candidate variable triggering the change in regime can be chosen among all the variables that are included in the model. Furthermore, in a TVEC, the error correction term can be used to steer the switching between the regimes.

Our main analysis focuses on the inter-relationship between house prices and credit, which are both heavily dependent on the framework conditions, i.e. the state of the economy. Conversely, house prices and credit can also influence the economy. As a consequence, the framework contains, in addition to house prices and credit, also two variables that mirror the state of the economy, which are proxied by the prevailing business and monetary conditions. It is worth mentioning that our limitation to four variables is actually enforced by the fact that VAR models demand a rather high number of observations, a problem that becomes even more serious if multiple regimes exist.

To obtain first insights into the sources of potential nonlinearities, univariate models are specified for all individual series. Nonlinearities are not very strong at this level. While the equations for interest rates and credit for housing purchases do not show significant deviations from the linearity assumption, output and house prices exhibit some signs of threshold behaviour. By contrast, the nonlinear evidence based on multivariate models exhibits a much richer picture, implying that the relationships between the variables are subject to regime changes. Since credit and house prices do not exhibit any signs of non-linear cointegration, specifications based on stationary TVAR models seem to be appropriate. If output growth and interest rate changes are chosen as thresholds, two regimes can be distinguished. Conversely, three regimes are more appropriate if house prices and credit control the regime change. The generalized Impulse responses functions (GIRF) show that credit responds to house prices in the short run, while the reverse direction is less important. Hence, the modest recovery of credit at the current edge is partially driven by the recent acceleration of house prices, possibly due to the dominance of collateral effects. It should be noted that the results refer to the euro area as an aggregate. At a country level, experiences can substantially differ, due to, for instance, heterogeneous institutional conditions in national mortgage markets. This notwithstanding, the analysis from a euro area perspective is decisive in terms of monetary policy, as the latter is conducted for the entire monetary union.

The rest of the paper is structured as follows. Section 2 presents some descriptive analysis of the developments of house prices and credit in the euro area. The econometric methodology for threshold models is discussed in Section 3. Section 4 presents the results, and Section 5 concludes.

2 Credit growth and house prices in the euro area

In the period preceding the financial crisis, the euro area experienced strong house price growth, especially directly before the crisis. Loans for housing purposes expanded at much higher rates for more than a decade, probably driven by favourable business and monetary conditions. The development came to a halt during the crisis. Credit started to increase at much lower rates, and house prices stagnated at the euro area wide level. While house prices recovered partially since then, credit growth has remained rather subdued, likely also because of the strong acceleration before the crisis. At the same time, the moderate dynamics could also indicate high repayments due to the strong increases in mortgage loans in the period preceding the crisis. The repayments can substantially weigh on the net lending figures at the current stage.

-Figure 1 about here-

Although a unique direction of causality between house prices and credit cannot be established, plausible lead-lag structures can be considered. While the contemporaneous correlation between the q-o-q growth rates is about 0.7, the dynamic correlations do not exceed 0.2. Likewise, the evidence for self-reinforcing spirals is not overwhelmingly strong. Figure 1 displays the developments of real loans for housing purposes and residential property prices in the euro area and the four largest member states, i.e. Germany, France, Italy and Spain, which overall represent roughly 75 percent of euro area GDP.

Booming house prices clearly promoted the economy especially in France, Spain and, to some extent, also in Italy in the period before the crisis, with increasing output gaps and higher inflation pressure. During the financial crisis, the fall in house prices triggered loan defaults and led to deep recessions, particularly in Spain, which had experienced a strong acceleration in house prices prior to the peak. The effects were also long-lasting, as output exceeds its potential value in France and Spain only since a few years. The negative effects of the crisis turned out to be even more persistent in Italy. House prices continued to fall on average until recently, presumably driven by declining prices for existing and less renovated homes. Real house prices showed a fundamentally distinct pattern in Germany due to different institutional characteristics, such as tighter regulation in mortgage markets and higher incentives for households to rent. German house prices decreased prior to the financial turmoil, partially in response to the former overinvestment in the first years after the German unification. In contrast to the experience in other countries, however, house prices accelerated markedly in the aftermath of the crisis. The recent rise in property prices in the euro area is heavily rooted in the German evolution.

Despite the common monetary policy and advances in the integration of financial and capital markets, such as the formation of the European banking union, the credit experience was also different across the member states. In the decade preceding the crisis, credit for housing purposes expanded at relatively high rates in France, Italy and most spectacularly in Spain. Since the crisis, loans accumulated at much lower rates and even decreased in Spain. While the stock of credit still increased in France, house prices stagnated until recently. Similar to the developments in property prices, credit exhibited a distinct pattern in Germany. After the rise around the introduction of the euro, loans stagnated for a long period. They gained pace in recent years, but the acceleration still lags behind house price developments. Overall, the behaviour of credit and house prices shows substantial heterogeneity across euro area members, at least among the largest countries. This notwithstanding, in line with the common monetary policy in the euro area, the subsequent analysis is conducted at the euro area wide level, where national particularities are likely to cancel out.

3 Modelling regime switching behaviour

The empirical analysis of the relationship between credit and house prices is investigated in a nonlinear framework. By its very nature, the latter includes the linear benchmark as a nested case and the more general models are selected once the linearity assumption is rejected. Threshold autoregressive (TAR) models provide a convenient instrument to deal with nonlinearities and they allow for a switching between different regimes once certain conditions are met. In principle, Markov switching (MS) models offer an alternative approach to deal with the nonlinearities.[5] The basic difference between the TAR and MS models relates to the process that triggers the regime-change. In the TAR environment, the switching is modelled explicitly in terms of observed variables, whereas the regimes are unobservable under the MS approach. Under the latter approach, the switching is governed by a (hidden) Markov chain with estimated transition probabilities. However, in this case it is difficult to determine why the process is in a particular regime and in a certain period. Therefore, the switching between the regimes can be better described in the TAR framework. Because of the limited number of observations, only a few regimes can be distinguished, and usually, there are not more than two or three regimes that can be detected (which would imply one or two "switches").

In order to explore the sources of potential nonlinearities and the presence of regime switches, the analysis is first carried out using univariate autoregressions. It has to be kept in mind that, if the number of lags is sufficiently large, these models can be viewed as an approximation of a more complex multivariate framework. Although univariate models are less appealing from an economic point of view, they often outperform multivariate models in terms of forecasting (Brockwell and Davis, 2016).

After analysing the linear autoregressive framework, an extension of it is analysed using the SETAR (self-exciting TAR) approach. A feature of this model is a feedback loop which regulates the process when its output becomes unusual. Specifically, the autoregressive coefficients and deterministic terms can take different values depending on the fact, whether a past value of the process is above or below a certain threshold δ. In the case of a two regime SETAR model, it follows:

5 The TAR and MS specifications are not nested. Hence, there is no statistical test that can be used to compare these models as rivalling hypotheses.

(1) $$y_t = (\alpha_0 + \sum_{i=1}^{p} \alpha_i y_{t-i}) I(y_{t-d} \leq \delta) + (\beta_0 + \sum_{i=1}^{q} \beta_i y_{t-i}) I(y_{t-d} > \delta) + u_t$$

where the indicator function *I* is equal to 1, if the corresponding argument is true and 0 otherwise. Lags of the variable *y* are included to ensure the white noise properties of the error term *u*. A switch between the regimes can occur with a delay of *d* periods. The unknown parameters *d* and δ are estimated by a grid search. The values of the variable are sorted in descending or ascending order and a certain percentage (i.e. the "trimming parameter") of the lowest and highest values is excluded to ensure a minimum number of observations in both regimes. For each lag *d*, the sum of squared residuals is estimated for all tentative thresholds. The optimal model minimizes this criterion.

Testing for the presence of nonlinearities is nonstandard, as the threshold parameter is not identified under the null. Therefore, an information criterion must be used that may offer a useful guideline to find the optimal model. As a rule of thumb, the threshold parameter should not be penalized in this exercise, although the evidence is not unique in this respect (Strikholm and Teräsvirta (2006)). Hansen (1999) suggested an *F*-type test to determine both the presence of a threshold and the appropriate number of regimes.[6] Based on the sum of squared residuals (*SSR*) of two models with *i* and *j* regimes, the statistic

(2) $$F_{ij} = n(SSR_i - SSR_j) / SSR_j$$

is calculated under three options. The first two options examine the null hypothesis of linearity (one regime) against the alternatives of two (1 versus 2) or three (1 versus 3) regimes, i.e. one or two thresholds. The third option is represented by a specification test. Once the presence of a threshold is confirmed, the test explores whether a model with 2 or with 3 regimes (2 versus 3) is more in line with the data. The *p*-values are obtained through bootstrap methods, based on the residuals of the model under the null.

While SETAR models investigate whether nonlinearities can be attributed to the evolution of a specific variable, the presence of nonlinearities in the economic relationship between different

6 Further approaches are based on functions of certain test statistics calculated for particular threshold values. These functions might include the mean or the supremum of the individual statistics (Andrews and Ploberger (1994)).

series can also be tested in context of TVAR models. In terms of notation, TVAR models with two regimes can be represented as follows:

(3) $$Y_t = (A_1 Y_t + B_1(L)Y_{t-1})I(z_{t-d} \leq \delta) + (A_2 Y_t + B_2(L)Y_{t-1})I(z_{t-d} > \delta) + u_t$$

where Y denotes a vector of m endogenous variables, A_1 and A_2 stand for the parameters of the contemporaneous relationships. $B_1(L)$ and $B_2(L)$ are polynomials in the lag operator, and z is the threshold variable, normally included in Y. The matrices A and B are regime specific. The TVAR can also have a recursive structure, i.e. the A matrices are lower triangular, according to the Cholesky decomposition. In principle, the threshold variable z can be exogenous. However, an endogenous z offers a somewhat richer interpretation, as a switch between the regimes can also occur from shocks hitting other variables of the system. If the TVAR is rewritten in a reduced form, the model equations can be estimated separately using OLS within each regime, conditional on the threshold parameter. Similarly to the univariate case, the threshold is determined through a grid search. For a pre-specified number of regimes, the model is estimated for all potential threshold values. The ultimate threshold is the value where the determinant of the residual covariance matrix is minimized. Additional thresholds (more than two regimes) are estimated conditional on the former thresholds (Gonzalo and Pitarakis, 2002).

In the context of TVAR models, likelihood ratio tests are applied, which represent a multivariate extension of the univariate F tests (Hansen, 1999). The LR test statistic:

(4) $$LR_{ij} = n[\ln(\det(\hat{\Sigma}_i) - \ln(\det(\hat{\Sigma}_j)]$$

is based on the determinants of the residual covariance matrices of models with i and j regimes. For a pre-selected threshold variable, the test is run under three variants. The first two options examine the null of linearity against the alternatives of models with one or two thresholds. Once nonlinearity is detected, the third test investigates the null of a TVAR with one threshold against the alternative of a TVAR with two thresholds. As the distribution of the statistic is nonstandard, p-values are extracted from the bootstrap procedure.

In the context of the SETAR and TVAR analyses, the time series must be stationary. Hence, the models are better suited for variables expressed in their growth rates rather than in levels, as the former are often stationary.[7] In case of nonstationary variables, regime switching could be modelled within the threshold VEC (TVEC) approach, provided that a long-run equilibrium exists. The switch between the regimes could then be controlled by the error correction term as the deviations from the long run might be smaller in absolute value than the threshold or not. By contrast, the cointegration vector is restricted to be stable across the regimes, as it represents fundamental economic principles. Since in such a case, both the threshold and the cointegration parameters are involved in the estimation procedure, the grid search becomes highly complex. Hansen and Seo (2002) and Seo (2011) proposed a TVEC model which is restricted to two variables and two regimes.[8] Each regime can then embody different deterministic components and short-run dynamics. If an external variable is selected as potential threshold, the estimation can be simplified, as the interdependency between the error correction term and the threshold parameter is suspended. In that case, multivariate specifications (more than two variables) may be allowed, see Gonzalo and Pitarakis (2006) and Krishnakumar and Neto (2015).

Hansen and Seo (2002) suggest a supremum LM test for the null of linear cointegration against the alternative of threshold cointegration. Conditional on the cointegration parameter from the linear model, LM tests are conducted over a wide range of threshold values. As the assumption of linear adjustment does not hold, the initial estimation of the cointegration vector by ML techniques might not be appropriate. Hence, the grid search is performed within a confidence interval around the (unique) cointegration parameter. Seo (2006) instead proposes a test for the null of no cointegration against the alternative of threshold cointegration. Under the null, the feedback coefficients are jointly equal to 0. In that case, a long-run does not exist, as no adjustment towards an imaginary equilibrium takes place. The null is rejected, if at least one feedback parameter is different from 0. For both tests, the number of thresholds has to be specified in advance. Furthermore, since the distributions are nonstandard, the *p*-values are calculated by bootstrap methods.

The impact of shocks on certain variables is usually investigated by impulse response functions. In a linear model, the impulse responses are proportional with respect to the size of the shocks

[7] More precisely, TAR models are required to be globally stationary, i.e. local nonstationarities in an inner regime separated by thresholds are allowed (Bec, Ben-Salem and Carrasco, 2004).

[8] In addition, several authors have proposed bivariate models with probably three regimes for the error correction term, see Wang, Chan, and Yau (2016) and Stigler (2019).

and symmetric, as positive and negative shocks lead to the same responses in absolute value. Initial conditions are not important for the results. These assumptions do not hold in a nonlinear model. The responses to shocks can differ, depending on whether the economy starts in a lower or upper regime. If the shocks are small, the process may stay in the same regime, provided that the distance to the threshold is sufficiently large. Switches between the regimes are more likely in case of large shocks or can occur once the process is near the threshold value. To deal with these issues, generalized impulse response functions (GIRF) have been proposed (Koop, Pesaran and Potter, 1996). The response of a variable to a shock depends on the history of the process, the size and sign of the shock at time *t* and the size and the sign of all shocks up to time *t*+*d*, where *d* is the lag of the threshold. In order to remove the impact of the history and intermediate shocks, the GIRF are obtained by simulation methods. Here, the forecasted paths with and without a shock are compared:

$$GIRF(k, u_t, \Omega_{t-1}) = E(Y_{t+k} \mid u_t, \Omega_{t-1}) - E(Y_{t+k} \mid \Omega_{t-1}) \tag{5}$$

In this equation, Ω denotes the information set available before the period of the shock (*u*). *E* is the conditional expectations operator and *k* is the forecasting horizon. The starting conditions determine the state of the process in period *t*. It is obvious that different initial conditions, signs and sizes of the shocks can lead to different impacts. In order to remove these particularities, the GIRF refer to the average of the individual responses. Due to their construction, the GIRF are obtained from the reduced-form residuals. As the latter are typically correlated across the equations, the GIRF lack a structural interpretation. It is for this reason that the residuals need to be orthogonalized, for instance by using the Cholesky decomposition. After the decomposition, the impulse responses reveal the reactions of the variables to structural shocks.

4 Empirical results

The interlinkages between house prices and credit should not be explored by bivariate models, as the relationship might depend on the state of the economy. Therefore, control variables need to be added to enhance the structure of the model . However, large VAR models are not appropriate, as the degrees of freedom are usually rather limited. This restriction becomes even more binding, if different regimes are allowed.

Hence, a parsimonious specification is recommended. To capture the main effects, the VAR includes four variables, namely real house prices and real loans for house purchase as key variables of interest as well as two variables related to the state of the economy, namely the economic and monetary policy conditions. The former variable is measured by real GDP, i.e. nominal GDP divided by the GDP deflator (2010=100). The conduct of monetary policy is often proxied by the short term nominal interest rate. However, in the context of the financial crisis, the rate has become less informative, as it moved quickly to the zero lower bound. The ECB consequently switched to a series of unconventional monetary policy measures, with potential effects on the relationship between credit and house prices. Massive asset purchase programmes were implemented, and deposit rates for commercial banks have been set even below the zero lower bound (2014). In this environment, the standard short term rate is not appropriate to capture the various stages of monetary easing. Therefore, it is replaced by the shadow rate recently proposed by Lemke and Vladu (2017). While both rates coincide prior to the crisis they differ thereafter.

Real and monetary conditions can exert an impact on house prices and loans, but the reversed direction also applies. House prices refer to the price index for owner-occupied new and existing dwellings and credit to the stock of loans for house purchase. Compared to total credit, the loan measure is more closely linked to the real estate market. Its development has been subject to a spectacular development over the past decades. While the share of housing credit was 20 percent in the mid-1980s, it increased to almost 40 percent of total credit to the private sector up to the end of the sample, after a temporary fall during the financial crisis. In order to derive time series in real terms, both house prices and credit are divided through the GDP deflator.[9] The series are seasonally adjusted and measured at the quarterly frequency. The sample period ranges from 1985 Q1 to 2018 Q4. The starting point is chosen to exclude a potential turmoil caused by the second oil crisis. All variables are obtained from the ECB database and are expressed in logs, except of the interest rate, which is measured in percent per annum.

9 Several robustness checks have been performed. For example, instead of loans for housing purposes, total credit to the private sector has been selected as the credit measure. Moreover, the HICP instead of the GDP deflator was used to construct real house prices. While there are some differences in detail, the broad conclusions are not altered.

Stationarity tests are carried out to determine the integration properties of the variables. According to the standard ADF test, all variables include a unit root in their level representation, but their first differences are stationary. To get initial insights into the sources of potential nonlinearities, univariate SETAR models are estimated. Since the *p*-values are obtained by simulation, they can vary by one or two percentage points in different model evaluations. As a consequence, there is some uncertainty around the numbers, as the exact distribution of the statistics is unknown. Even then, however, the evidence for nonlinearities is not overwhelmingly strong. Apart from real output and house prices, the null of linearity cannot be rejected against the alternatives of two or three thresholds (Table 1). While there is some evidence for two regimes, three regimes are less appropriate. Notwithstanding, threshold behaviour might occur in the relationships between the variables.

-Tables 1 and 2 about here-

The interlinkages between the variables are explored in a TVAR or TVEC framework, depending on the results of cointegration tests. The latter tests reveal that a long-run equilibrium between real house prices and loans does not exist (Table 2). No signs for nonlinear cointegration are detected, as both hypotheses of no cointegration and linear cointegration cannot be rejected, therefore excluding the use of a TVEC approach[10]. The results do not depend on the number of switches under the alternative, i.e. whether one or two thresholds are assumed. Therefore, the subsequent analysis proceeds by estimating VAR and TVAR models, which are specified in terms of variables in first-difference, as those are stationary.

In the first instance, a linear VAR is, therefore, estimated and taken as a benchmark. The ordering of the variables (real GDP growth, changes in the nominal interest rate, real credit growth for housing purchases and changes in real house prices) is rather standard (see, for example, Goodhart and Hofmann, 2008). The relative position of credit and house prices is arbitrary, given the complex nature of their relationship. The standard setup assumes that credit comes first, i.e. credit can influence house prices in a contemporaneous way, but not vice versa. One

[10] It may be argued that the threshold cointegration tests are biased, as they refer to bivariate settings. However, to the best of our knowledge, tests for thresholds in larger models and potentially multiple cointegrating vectors are not available. To shed more light on the cointegration issue, the linear specification with all four variables is run. Here, the null of no cointegration cannot be rejected even at the 0.1 level.

claim for this setting is the lower volatility of credit compared to house prices. In any case, the reversed ordering has only minor effects on the results. Figure 2 displays the linear impulse responses, where the structural shocks are identified in a recursive manner (Cholesky decomposition). Since the impulse responses are estimated rather imprecisely, one standard error bands are used instead of the conventional significance levels, as suggested by Sims and Zha (1999). Even under this setting, not all responses are significant.

-Figure 2 about here-

The impulse responses are broadly consistent with theoretical reasoning. Unexpected hikes in output, house prices and credit raise nominal interest rates, due to the endogenous reaction of monetary policy. Likewise, a positive interest rate shock tends to reduce output, credit and house price growth. Hence, a tighter monetary policy stance can limit potential bubbles emerging in the real estate and credit markets. Furthermore, real credit and house prices are stimulated through higher demand, implying that the response of these variables to GDP shocks is positive. Shocks to credit and house prices are expected to raise output. In addition, an increase in credit leads to soaring house prices and vice versa.

-Table 3 about here-

Although the impulse responses are broadly in line with the predictions from standard theory, the linear VAR might not be justified.

In fact, the evidence for the presence of nonlinearities is rather strong and more complex than in the univariate case. Therefore, the interlinkages between the variables are not stable and subject to regime switching behaviour. In particular, the threshold is unique, if output growth or interest rate differentials are chosen to trigger the regime change. With respect to house prices and credit growth, a model with two thresholds outperforms the single threshold only at the 0.1 level. Nonetheless, the null of linearity can be rejected against a model of two thresholds at the 0.05 level. Therefore, if house prices or credit are selected as a threshold variable, TVARs with three regimes (low, medium, high) are more appropriate.

The estimated thresholds are shown in Figure 3. The switching points for output growth (0.092) and interest rate changes (0.061) are in the neighbourhood of the horizontal axis. Roughly speaking, the affiliation to a particular regime depends on the sign of the threshold variable. More specifically, if GDP growth falls below 0.092 percent on a q-o-q basis (corresponds to almost 0.4 at annualized rates) the system is located in the lower regime. The higher regime prevails in a faster growing economy. The coefficients of the TVAR and the pattern of impulse responses can change around this point. The higher regime may be interpreted as the normal one, as it includes 80 percent of the observations.

-Figure 3 about here-

Three regimes can be distinguished for real credit growth. Breaking points are 0.591 and 1.065, implying that the lower regime is in place in roughly one third of the cases, most notably since the financial crisis. Note that the evolution at the current edge is still located in the lower regime, despite the recent recovery. The "high regime" holds in little more than 50 percent of the credit expansion. The "medium regime" is more exotic than normal, as it includes only 17 percent of the observations.[11] Similar to credit, real house prices are captured in three regimes, with thresholds of 0.027 and 1.084. About 35 percent of the actual growth rates fall in the lower, 45 percent in the medium and 20 percent in the high regime. Booming house prices can be detected during the economic upswing in the late 1980s and in the period around the introduction of the common currency area. Despite the acceleration in recent years, the current evolution is still in the normal regime.

-Figure 4 about here-

The rejection of the linear VAR does not depend on the choice of the threshold variable. Hence, the relationship between credit and house prices can be investigated under four model specifications. The GIRF shown in Figure 4 focus on the response of house prices and credit to a shock in the other variable, respectively. In contrast to the usual impulse responses, the GIRF

[11] It should be noted that the existence of the mid-range is partially a result of the trimming parameter (0.15). The medium regime is removed, if a higher parameter value is selected.

do not assume that the process stays in a certain regime. Instead, they are robust against the possibility of a regime change.

The responses are rather moderate and short-lived. However, in contrast to the linear approach, the causality runs from house prices to credit, probably due to the dominance of collateral effects. In contrast, the reverse causality is less important. Hence, the modest evolution of credit at the end of the sample can reflect large repayments but might be also linked to the recent acceleration of house prices. The fact that the credit evolution is still in the lower regime does not imply that it acts as a brake for stronger house prices, as the underlying direction of causality is not supported.

5 Conclusions

The critical role of house prices for macroeconomic and financial stability is widely acknowledged since the global financial crisis. While house prices showed spectacular increases and even a bubble-like behaviour in the pre-crisis years, their spectacular fall thereafter was accompanied by deep re-cessions in many countries. Loose monetary conditions, such as the easy availability of credit, are often blamed for the boom prior to a crisis. In this paper, the link between credit and house prices is investigated for the euro area in the context of nonlinear models. This choice is motivated by the idea that the linkages between can be suspected to be governed by a regime-switching behaviour. As a consequence, threshold VAR (TVAR) models are estimated, including real house price and real credit changes, business and monetary conditions. Optimal breakpoints are determined via a grid search. If output growth and interest rate changes serve as thresholds, two regimes can be distinguished. Three regimes are more appropriate, if house prices and credit control the regime change. Non-linear impulse responses reveal that credit respond to house prices. The reverse direction appears to be less important. The credit dynamics at the current edge can indicate substantial repayments due to the strong increases in mortgage loans in the pre-crisis period. Nonetheless, the recent acceleration of house prices in the euro area could also play a role. The fact that credit growth is still in the lower regime does not imply that the acceleration of house prices could have been even stronger.

Finally, monetary policy might have played a role. While the evolution of the business cycle in the euro area is still modest, the prolonged period of low interest rates cut the profit margins of banks, with adverse effects on credit supply. Saving in financial assets have become less attractive, thereby causing additional housing demand.

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Table 1: Tests for SETAR effects

	1 versus 2	1 versus 3	2 versus 3
Real house prices	23.30 (0.017)	39.04 (0.044)	13.39 (0.272)
Real credit growth	9.878 (0.563)	20.73 (0.609)	10.09 (0.640)
Nom. interest rate	5.460 (0.684)	7.244 (0.967)	1.696 (0.997)
Real GDP growth	14.77 (0.041)	23.28 (0.122)	7.670 (0.464)

Notes: Sample period: 1985 Q1-2018 Q4. *F*-type tests for null of linear AR against SETAR alternative with 2 (1 versus 2) and 3 (1 versus 3) regimes, according to equation (2). In addition, test statistics for SETAR models with two versus three regimes (2 versus 3) are reported. Bootstrap distribution, *p*-values (in parentheses) based on 1000 replications, trimming parameter 0.15. Maximum delay 4.

Table 2: Tests for cointegration between real house prices and real credit

	Test statistic
Linear versus threshold cointegration	11.73 (0.986)
No cointegration versus threshold cointegration	34.01 (0.998)

Notes: Sample period: 1985 Q1-2018 Q4. Hansen-Seo (2002) for the null of linear cointegration against threshold cointegration, Seo (2006) for the null of no cointegration against threshold cointegration. TVEC with an unrestricted constant and 2 lags, *p*-values in parentheses. The distributions for the supremum LM tests are based on 1000 replications, trimming parameter 0.15.

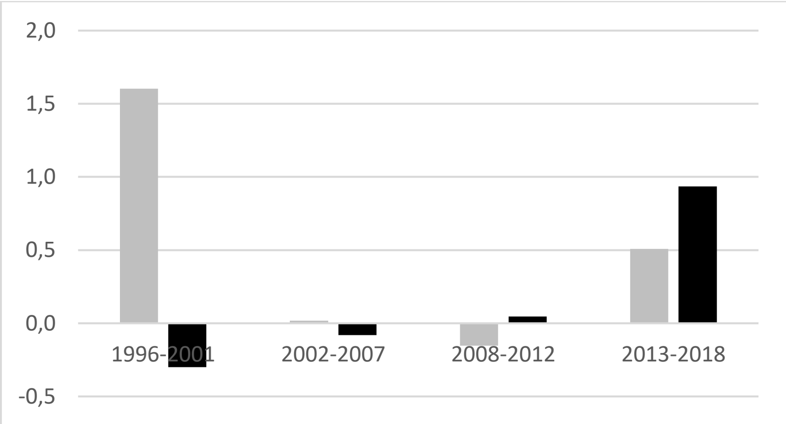
Table 3: LR tests for linearity and threshold behaviour in the stationary VAR model

	1 versus 2	1 versus 3	2 versus 3
Real house prices	55.126 (0.312)	129.89 (0.058)	74.778 (0.120)
Real credit growth	61.878 (0.147)	141.50 (0.021)	79.625 (0.064)
Nom. Interest rate	77.881 (0.012)	140.52 (0.014)	62.639 (0.222)
Real GDP growth	78.134 (0.020)	136.86 (0.028)	58.728 (0.321)

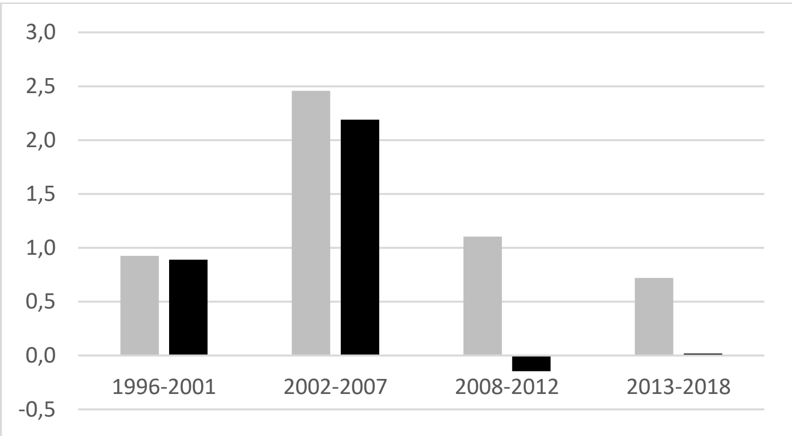
Notes: Sample period 1985 Q1-2018 Q4. LR tests for linear VAR (=1 regime) against alternative of threshold VAR with two regimes (1 versus 2) and alternative of threshold VAR with three regimes (1 versus 3). Last column shows test of threshold VAR with two against a threshold VAR with 3 regimes (2vs3). Bootstrap distribution, *p* values based on 1000 replications. VAR with a constant and 2 lags, trimming parameter 0.15.

Figure 1: Real house prices and loans for housing in the four largest euro area countries and the monetary union

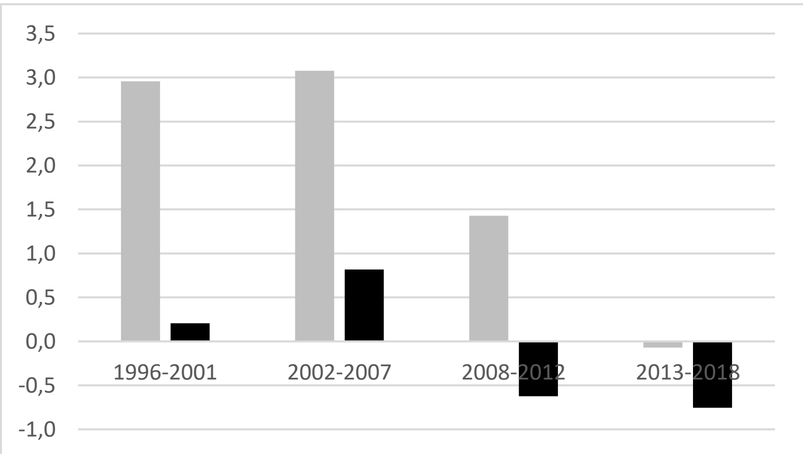
Germany



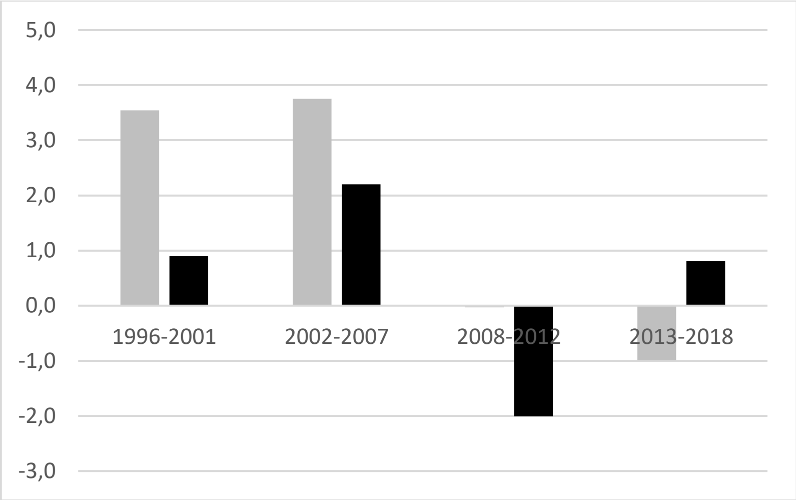
France



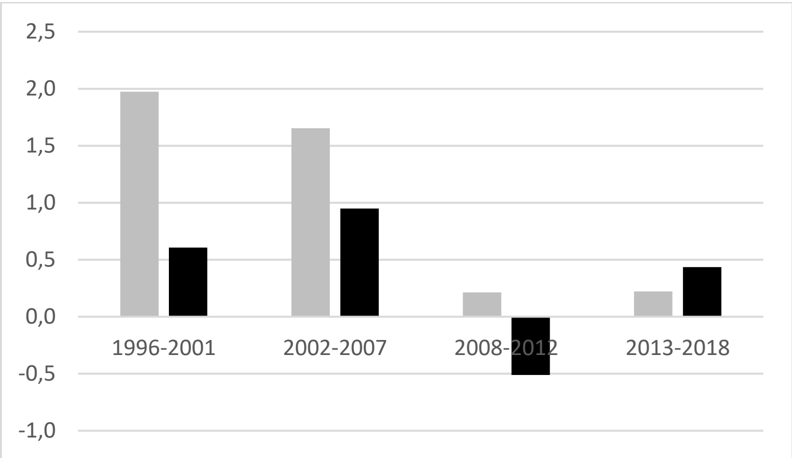
Italy



Spain

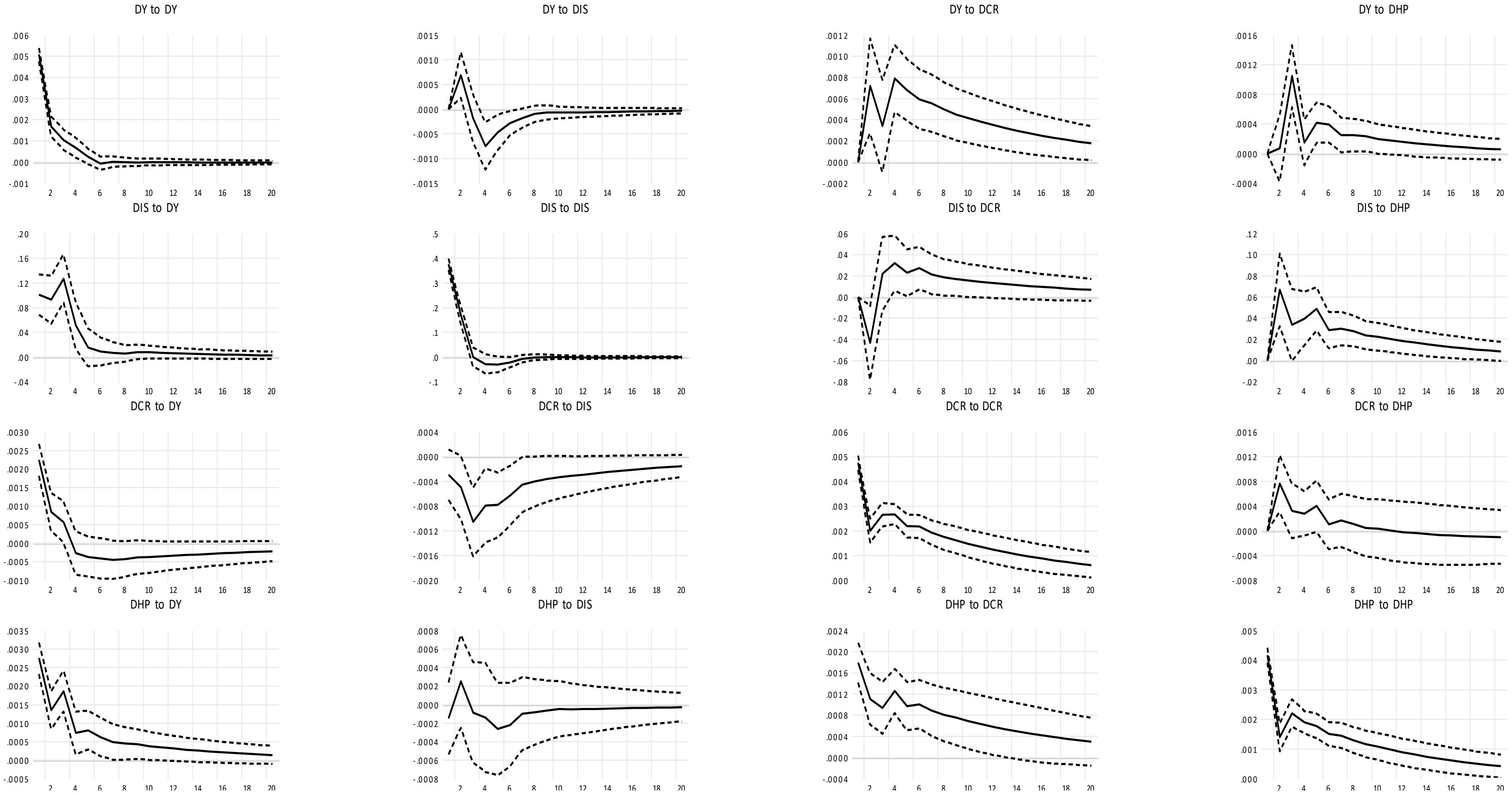


Euro area



Sources: ECB, ECB Statistical warehouse. Notes: Loans for housing purposes (grey) and house prices in selected euro area member states and the entire monetary union. Growth rates in the respective period, q-o-q, geometric average. The nominal variables are divided by the GDP deflator to obtain series in real terms.

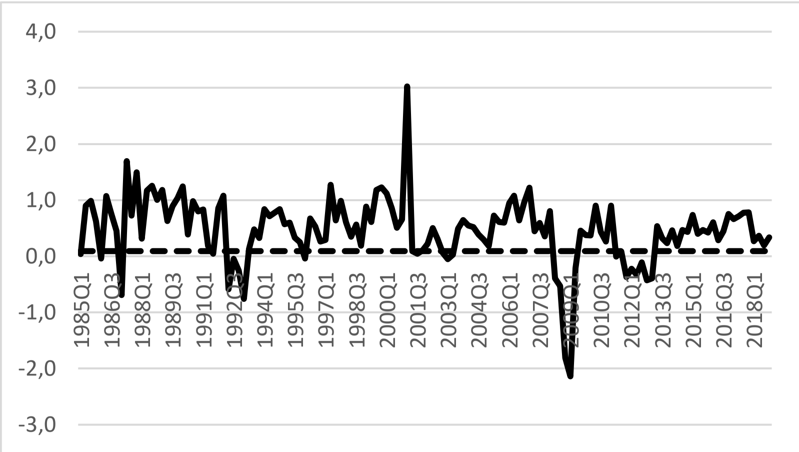
Figure 2: Impulse responses from the linear VAR benchmark



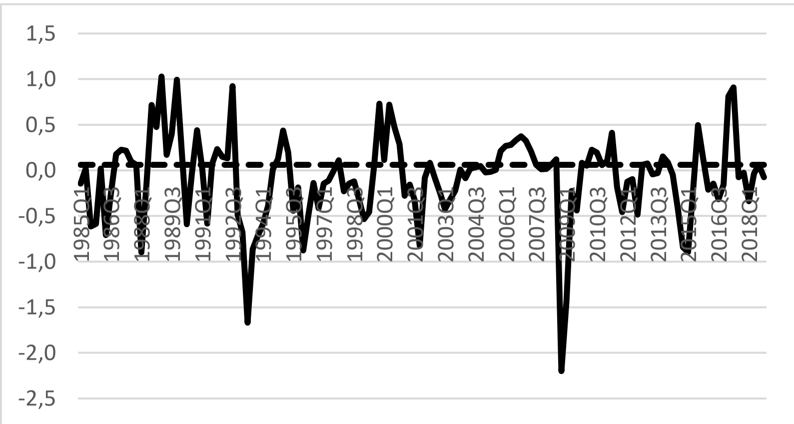
Notes: Sample period 1985 Q1-2018 Q4.VAR with a constant and 2 lags (SBC). Rows denote the responses of the respective variables to shocks of the variables in columns. Cholesky decomposition, ordering real GDP growth (DY), change in nominal short-run interest rate (DIS), real loan growth (DCR) and change in real house prices (DHP). Dashed lines represent one standard error bands.

Figure 3: Thresholds in stationary TVAR models

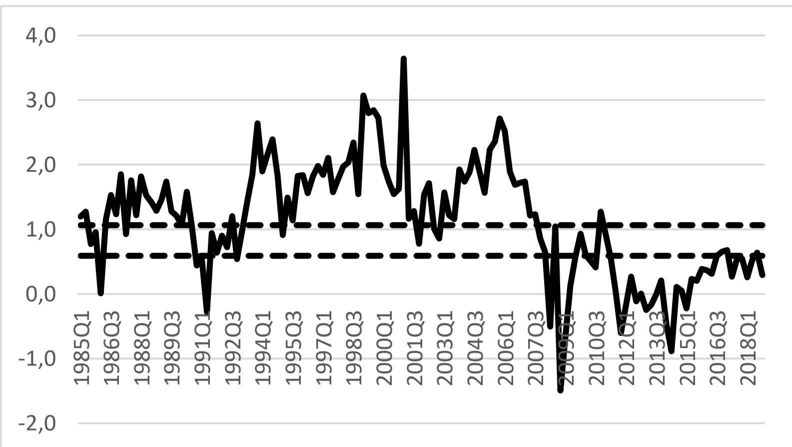
Real GDP growth



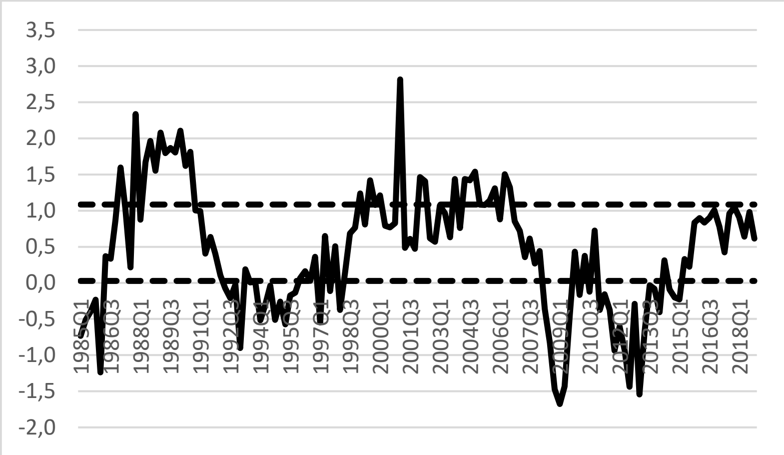
Interest rate changes



Growth in real credit for housing purposes



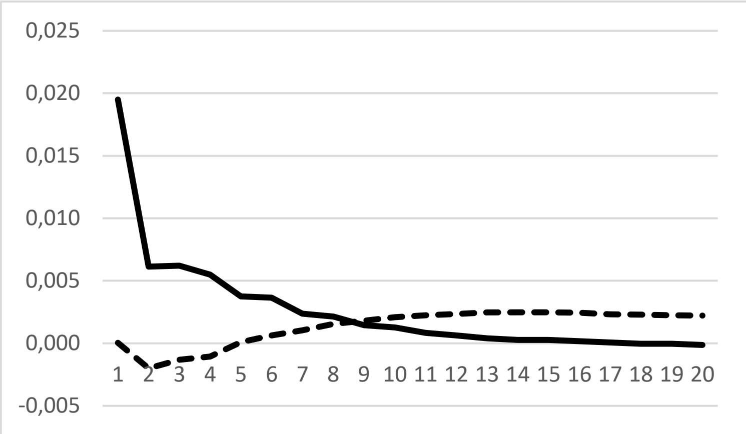
Growth in real house prices



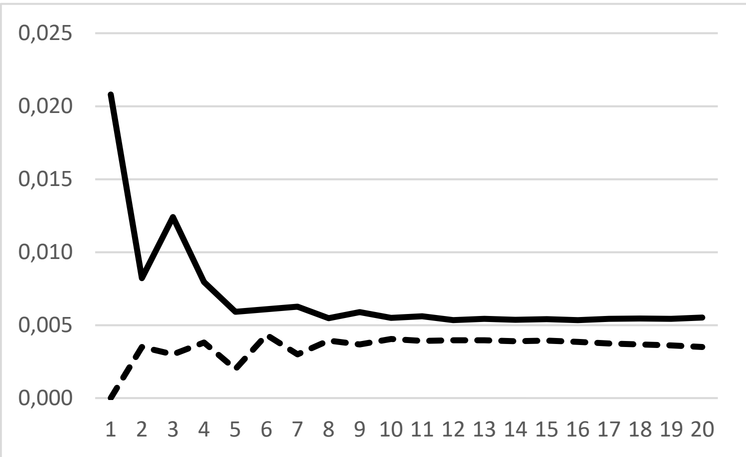
Notes: Sample period 1985 Q1-2018 Q4. Thresholds refer to q-o-q growth rates (changes) expressed in percent. To calculate annualized rates, they should be multiplied by 4. TVAR with a constant and 2 lags, trimming parameter 0.15.

Figure 4: Nonlinear responses of real credit growth and house prices changes

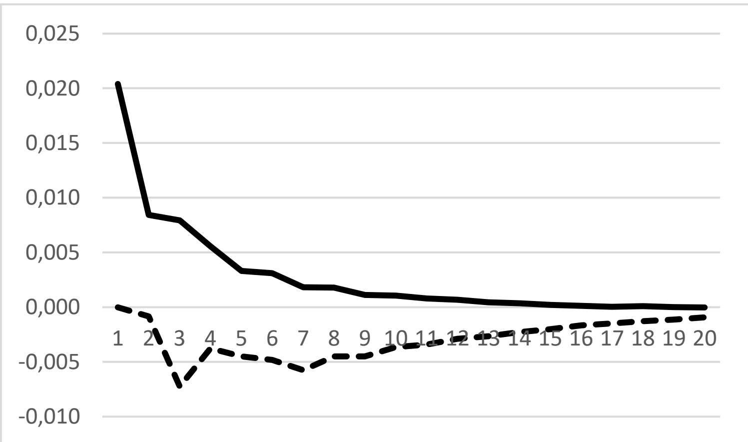
Threshold variable: Real GDP growth



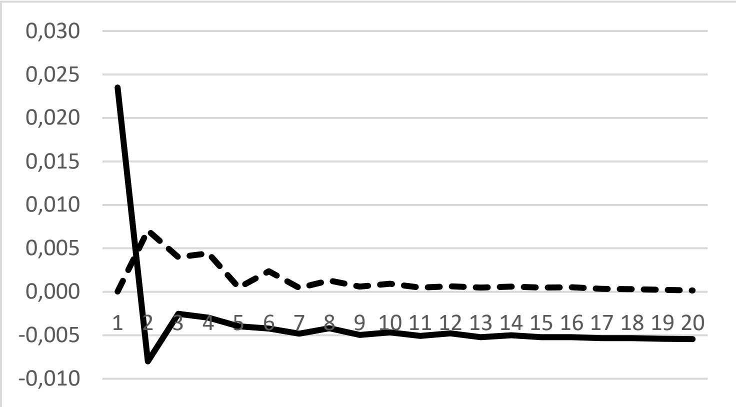
Threshold variable: Interest rate changes



Threshold variable: Growth of credit for housing purposes



Threshold variable: Growth of real house prices



Notes: Sample period 1985 Q1-2018 Q4. TVAR with a constant and 2 lags (SB criterion). Generalized impulse responses are based on 100 histories, 200 replications per history. Responses of real loans growth (solid line) and real house price changes (dashed line) to shocks in the other variable.