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Monetary Policy and Asset Price Bubbles [*]

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Abstract

This paper assesses the linear and non-linear dynamic effects of monetary policy on asset price bubbles. We use a Principal Component Analysis to estimate new bubble indicators for the stock and housing markets in the United States based on structural, econometric and statistical approaches. We find that the effects of monetary policy are asymmetric so the responses to restrictive and expansionary shocks must be differentiated. Restrictive monetary policy is not able to deflate asset price bubbles contrary to the "leaning against the wind" policy recommendations. Expansionary interest rate policies would inflate stock price bubbles whereas expansionary balance-sheet measures would not.

Keywords: Booms and busts, Mispricing, Price deviations, Interest rate policy, Unconventional monetary policy, Quantitative Easing, Federal Reserve.

JEL Classification: E44, G12, E52.

1. Introduction

Asset price bubbles are a threat to financial and macroeconomic stability. However, no consensus has been reached regarding how policymakers should deal with them. The role of monetary policy remains disputed. Two closely related questions arise in the policy debate. Does monetary policy contribute to fuel asset price bubbles and is it able to deflate bubbles? This debate has recently resurfaced with the implementation of unconventional monetary policy since 2008 and their potential adverse effects on financial stability.[1] There are two broad opposite views in the literature on this issue. Borio and Lowe (2002), Cecchetti et al. (2003) and Woodford (2012) are in favour of a "leaning against the wind" approach which considers that expansionary monetary policy contributes to the emergence of asset price bubbles and restrictive policies can reduce them.[2] An alternative view, the "modified Jackson Hole consensus", would not recommend using monetary policy to deal with bubbles and financial stability issues and rely on macroprudential tools (see e.g. Gerlach, 2010, Svensson, 2012 and Collard et al., 2017).[3] Bernanke and Gertler (1999, 2001) suggest that a "cleaning afterward" approach would be more optimal and Svensson (2016) demonstrates that the net benefit of the "leaning against the wind" strategy is negative as it entails a higher unemployment rate during both crisis and non-crisis periods.

This policy debate echoes the lack of consensus in the theoretical literature on how to represent the formation and dynamics of bubbles.[4] Within existing bubble models, the role of monetary policy is not clearly established: its effect on asset price bubbles depends on the nature of bubbles. First, in a rational bubble model à la Blanchard and Watson (1982), the asset price is decomposed into a fundamental value, equal to the sum of expected cash-flows, and a bubble component, which is a rational stochastic deviation from the fundamental value growing with the discount factor. Within this framework, Galí (2014) shows that bubbles are linked to monetary policy because the discount factor is related to the real interest rate. With nominal rigidities, central banks influence the real interest rate and higher rates would increase the size of the bubble. This effect would dominate the negative impact of a restrictive monetary policy on the fundamental value since rational bubbles have explosive conditional expectations (Diba and Grossman, 1988). Second, in models accounting for financial market imperfections, Allen and Gale (2000, 2004) suggest that expansionary monetary policy would feed bubbles through the credit dynamics. This transmission channel is also emphasized by Gruen et al. (2005) and Christiano et al. (2010) who suggest central banks to adopt a "leaning against the wind" approach by limiting sharp credit expansion. Allen, Barlevy and Gale (2017) modify Galí (2014)'s framework to include a crowding-out of resources and credit-driven bubbles and show that higher interest rates can dampen bubbles.[5] The objective of this paper is to shed

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[1] Borio and Zabai (2016) and Juselius et al. (2016) fear that the benefits of unconventional monetary policies would decline with time while the risks to financial stability would increase. This view echoes the argument by Taylor (2009) that interest rates "too low for too long" in the United States (US) between 2001 and 2004 have triggered the housing market boom and subprime crisis.

[2] These authors also claim that price stability is not a sufficient condition to promote financial stability. Blot et al. (2015) find that there is no stable link between price and financial stability indicating neither that price stability promotes financial stability nor that price stability is correlated with financial instability.

[3] See Smets (2014) for a recent survey on the attitude of central banks towards financial stability.

[4] See Brunnermeier and Oehmke (2013), Scherbina (2013) or Martin and Ventura (2017) for surveys.

[5] A third strand of models does not give much role to monetary policy as private agents' behaviour is the key determinant of bubbles. They arise after some positive news, generally technology innovations, triggering a rise in the fundamental value, which is amplified either by coordination failures of rational arbitrageurs or by investors' overconfidence. For instance, Abreu and Brunnermeier (2003) and Ofek and Richardson (2003) develop models emphasizing informational frictions or heterogeneous beliefs. Kindleberger (2005) and Schiller (2015) document

light on this debate and to disentangle empirically the competing theoretical predictions about the effect of monetary policy on asset price bubbles. Contrary to the vast literature dealing with the impact of monetary policy on asset prices (see e.g. Rigobon and Sack, 2004 or Bernanke and Kuttner, 2005) or on asset price volatility (see e.g. Bernanke and Gertler, 1999), we focus specifically on the effect of monetary policy on the bubble component of asset prices.

Not all asset price variations are bubbles, however the empirical identification of the fundamental value and the bubble component is challenging. Asset price bubbles arise in many theoretical frameworks, and empirical tests are ill-designed to identify bubbles as they fail to disentangle between bubbles and misspecifications of the underlying theoretical model (Gurkaynak, 2008). However, the term "*bubble*" remains extensively used in the literature to characterize periods when asset prices rise sharply and seem disconnected from fundamentals. In practice, the terms "deviations", "price distortions", "booms and busts", "mispricing" or "over and undervaluations" could be used interchangeably. For simplicity, we use the term "bubble" to refer to these deviations. Whatever the term used, these deviations are a concern for many reasons. First, they may generate a misallocation of capital. Second, increases in asset prices driven by the bubble component may entail risk for financial stability jeopardizing the functioning of the financial system. Third, bubble bursts are associated with financial crises and with deeper and longer recessions.[6] Fourth, the transmission of monetary policy may be impaired if the dynamic of bubbles goes against the response of fundamentals to monetary policy. Thus, it is crucial to disentangle asset price movements driven by fundamentals from movements resulting from the bubble component.

Three empirical approaches - structural, econometric and statistical - may be considered to identify bubbles, but none of them has reached consensus. First, according to a structural model, the bubble is a deviation of the asset price from expected discounted cash flows. Second, the fitted value of an econometric specification accounting for a data-rich information set would provide an in-sample proxy for the fundamental value. Third, the literature has also relied on a statistical definition of the excessive level of asset prices, measured with a statistical filter (see e.g. Bordo and Wheelock, 2007, Goodhart and Hofmann, 2008, or Jordà et al., 2015). All these approaches present advantages and weaknesses to estimate the fundamental and non-fundamental components of asset prices.

The main empirical challenge to investigate our research question is to provide a measure of the bubble component of a given asset. Our objective is not to assess which bubble model is best, nor to date the beginning of bubbles so as to detect them in real-time[7], but to estimate *ex post* historical deviations of asset prices from their fundamental value, from their best in-sample fitted value, and from their trend. We develop a new bubble indicator using a Principal Component Analysis (PCA) to extract the common denominator of structural, econometric and statistical approaches used in the literature. This provides us with an agnostic and conservative representation of bubbles. By construction, the first principal component boils down to a model averaging of the structural, econometric and statistical approaches and maximizes their common variance, whereas the idiosyncratic dynamics of each approach will be dropped. So the first principal component can be considered as a robust measure of the

those episodes in financial history when increases in asset prices have been observed after technological booms, which were believed to give rise to a "new era".

[6] It must be noted that financial crises are not only triggered by asset price bubble bursts. Financial leverage and credit booms also matter for financial stability (see Adrian and Shin, 2008). This issue is left for further research.

[7] Phillips et al. (2011), Phillips and Yu (2011), Homm and Breitung (2012) and Phillips et al. (2015) propose recursive unit root tests to detect bubbles in real-time. These works focus on the explosive behaviour of bubbles so to measure their start in real-time, rather than the deviation of prices to their fundamental which is our object of interest.

bubble component. To that end, we first estimate fundamental or trend values and compute deviations of asset prices from these fundamental and trend values. We do so for stock and housing prices in the United States (US). Second, estimating a PCA of these deviations, we compute bubble indicators for each asset price. We acknowledge that looking at aggregate variables, both on the asset and geographical dimensions, might hide heterogeneous dynamics between subcomponents of the asset class (banking vs. intermediate goods sectoral indices, for instance) or between US states (North Dakota vs. Florida, for instance).[8] However, this aggregation bias would lead us to underestimate bubble episodes and so the potential effect of monetary policy on bubbles. Significant estimates would be all the more robust that they should be seen as a lower bound of this effect.

The question of whether monetary policy may trigger asset price booms and busts has been extensively dealt with in the literature. Detken and Smets (2004), Ahrend et al. (2008) and Khan (2010) observe that stock and housing prices tend to increase excessively when short-term interest rates are below the level suggested by a Taylor rule. Taylor (2009) asserts that monetary policy in the early 2000s has fueled the housing boom in the US. This view has been challenged by Dokko et al. (2011) and Kuttner (2012) who suggest that the housing market dynamics would not have been strongly modified if interest rates had followed the Taylor rule. Del Negro and Otrok (2007) also conclude that monetary policy weakly contributed to the housing price dynamics in the US. Besides, Bordo and Wheelock (2007) provide evidence of a weak correlation between interest rates and excessive stock price increases. However, these papers do not rely on a structural identification of the bubble component of asset prices but focus on episodes of asset price dynamics that they consider as excessive. The closest papers to ours are Basile and Joyce (2001), Galí and Gambetti (2015) and Beckers and Bernoth (2016), but they only rely on the rational bubble model. The first assesses the contribution of monetary policy to the variance of bubbles. The second suggests that monetary policy tightening in the US may increase asset prices depending on the size of the bubble component. A policy rate hike reduces the fundamental value, but increases the bubble component since the bubble grows with the interest rate. For a small bubble component, the standard negative effect on the fundamental value dominates, whereas monetary policy tightening feeds the bubble and increases asset price when the bubble component is large. The third paper challenges the finding of Galí and Gambetti (2015) by including a time-varying expected equity risk premium as a key determinant of stock prices.

The contribution of this paper is to assess the dynamic impact of monetary policy shocks on our bubble indicators. It departs from the existing literature in four ways. First, our definition of the bubble component does not rely exclusively on a single model but hinges on different representations of bubbles. Second, monetary shocks are identified by orthogonalising the policy instrument to the central bank information set as well as to private agents' one and to macroeconomic and financial variables following the Romer and Romer (2004) approach. Third, we disentangle the effects of expansionary and restrictive shocks to consider the potential asymmetries of the impact of monetary policy.[9] Fourth, we account for both standard and unconventional monetary policy by using shadow rates estimated by Wu and Xia (2016) and Krippner (2013, 2014) as an indicator of the overall monetary policy stance. We also disentangle the effects specific to the policy rate and central bank's balance sheet policies in a second step. We are thus able to compare the response of the stock and housing bubble indicators to shocks to different instruments. We investigate the dynamic impact of monetary

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[8] See for instance Del Negro and Otrok (2007) who estimate the effect of monetary policy on the common component of housing prices in the US using a dynamic factor model.

[9] Gali and Gambetti (2015) use a time-varying VAR model, but do not differentiate the sign of monetary shocks.

policy over a 2-year horizon, to account for its potentially slow and delayed effects on a given asset price bubble, by estimating local projections *à la* Jordà (2005) and controlling for the other asset price bubble. In particular, we take advantage of the flexibility of local projections to analyse the potential asymmetric effects of restrictive and expansionary monetary shocks.

A key message of this paper is that the effects of monetary policy are not symmetric and the responses to restrictive and expansionary shocks must be differentiated. The main result is that restrictive monetary policy cannot help deflating stock or housing price bubbles whereas expansionary policies do fuel stock market bubbles. This result is also confirmed on euro area (EA) data. Although this result does not support the predictions of rational bubble models (we find no evidence that higher rates increase bubbles), there is no evidence in favour of the "leaning against the wind" strategy as monetary policy is not a relevant instrument for central banks to deflate bubbles. However, the effect of expansionary monetary policy depends on the policy instrument. Whereas interest rates cuts – standard expansionary monetary policy – would inflate stock price bubbles, the risk that quantitative easing would inflate asset price bubbles does not materialize in the data. We even find that expansionary balance-sheet measures tend to lessen the bubble component of stock prices. From a policy perspective, central banks should be aware that there is a risk for expansionary interest rate policy to inflate stock price bubbles.

The remainder of this paper is organized as follows. Section 2 addresses the identification of asset price bubbles and section 3 the empirical strategy. Section 4 presents the effects of monetary policy on asset price bubble components. Section 5 concludes.

2. Identifying asset price bubbles

Asset price bubbles are unobserved and there is no consensus on the most appropriate way to identify them empirically. This reflects theoretical controversies illustrated by Brunnermeier (2008): "*Bubbles are typically associated with dramatic asset price increases followed by a collapse. Bubbles arise if the price exceeds the asset's fundamental value*". Such a definition emphasizes two dimensions of asset price bubbles. Bubbles rely on the notion of a fundamental value and on excessive movements of prices without reference to the fundamental value. Thus, Fama (2014) defines bubbles as "*an irrational strong price increase that implies a predictable strong decline*". We consider 3 specifications to identify bubbles. Under a "structural" approach, the bubble component is captured as the deviation from the fundamental value captured by the estimation of a discounted cash-flow model. Using a larger set of information to estimate the fitted value of the fundamental may provide a better proxy in-sample. This data-rich approach is called "econometric". Finally, we resort to a "statistical" approach to identify excessive deviations from a statistical trend. None of these models has reached consensus, but our main assumption is that together they capture the main properties of asset price bubbles.

2.1. A range of bubble models

The asset price (P_t) can be decomposed into a fundamental value (F_t) and a bubble component (B_t). None of these components is observable and must therefore be proxied. Under full information rational expectations and when agents are risk-neutral, the fundamental value is the sum of expected discounted future cash-flow payments:

$$F_t = \sum_{i=1}^{\infty} \left(\frac{1}{1+\rho}\right)^i E_t(D_{t+i}) \tag{1}$$

For stock and housing markets, cash-flows (D_t) are either dividends or rents, and the discount factor (ρ) is the long term interest rate. When the transversality condition holds, equation (1) is the solution to the standard asset pricing equation and $P_t = F_t$. Under rational expectations, a bubble solution exists ($B_t > 0$) such that equation (2) is:[10]

$$E_t(B_{t+1}) = (1 + \rho)B_t \qquad (2)$$

Assuming risk-neutral agents, a constant discount factor and constant cash-flows, equation (3) determines the fundamental value as follows:

$$F_t = D/\rho \qquad (3)$$

Equation (3) can then be used to identify the bubble component. We depart from the standard model by adding a time-varying proxy for the risk-premium, consistent with Beckers and Bernoth (2016), which would account for a time-varying risk aversion and the risk-taking channel of monetary policy. Henceforth, we first estimate equation (4) relating the asset price to the current cash-flow, the discount factor and the risk-premium with OLS:[11]

$$P_t = \alpha_0 + \alpha_1.D_t + \alpha_2.\rho_t + \alpha_3\phi_t + \epsilon_t^{OLS} \qquad (4)$$

where P_t is the log of the real price of a given asset (stock or housing), D_t is the associated cash flow of this given asset (real dividend or rent), ρ_t is the time-varying discount factor captured by long-term sovereign interest rates, and ϕ_t a proxy for the risk premium measured by the VIX index. ϵ_t^{OLS} is the residual and stands for the deviation of the asset price from fundamentals – purged from the risk-premia.

Under the "structural" approach, residuals stem from a model where the fundamental value is related to the cash-flows and the discount factor only. An alternative approach – "econometric" – rely on estimating empirical models where asset prices are represented by projections against a wide range of variables. By selecting a large set of macroeconomic and financial variables, equation (5) provides OLS estimates of the best in-sample prediction of a given asset price conditional to a given information set.

$$P_t = \beta_0 + \beta(L).P_t + \beta_1.M_t + \beta_2.F_t + v_t^{OLS} \qquad (5)$$

M_t and F_t are vectors of macroeconomic (rents and dividends, industrial production, GDP, real disposable income, inflation, confidence indicators and oil prices) and financial variables (real long term interest rate, monetary and credit aggregates, other asset prices and the VIX indicator). Lags of the endogenous variable are also included in the estimation. For all explanatory variables, 3 lags are included in the specification.[12] The residuals capture the component of the asset price unrelated to macroeconomic and financial fundamentals.

So far, asset prices are defined as the sum of a fundamental component and some residuals. However, if residuals are normally distributed, they would only capture small and short-lived deviations from fundamentals. Such static bubbles would result from anomalies in financial markets. They are likely to be small (Filardo, 2004) and henceforth are irrelevant for monetary policy and macroeconomic stability. The statistical properties of residuals shed light on the nature of the deviations captured by equations (4) and (5). According to the Cumby-Huizinga test for autocorrelation and for the Portmanteau test for white noise (Table 1), normality and the absence of autocorrelation is clearly rejected for ϵ_t^{OLS} but not for v_t^{OLS}. Consequently, the residuals from equation (4) may capture persistent deviations from the fundamentals – a

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[10] Rational bubbles may also depend on fundamentals as illustrated by Froot and Obstfeld (1991).
[11] We discuss later alternative specifications of the cash-flow model.
[12] Specifications with leads have also been tested but do not change the result and the residual dynamics.

bubble – whereas the residuals from equation (5) would not. For v_t^{OLS}, we consider a measure that takes into account the cumulative and dynamic process associated with a bubble formation. Each single deviation may be small and relatively irrelevant for policy makers but successive positive or negative deviations may signal a persistent deviation. A rolling-window sum of residuals captures dynamic and persistent deviations from fundamentals. We compute the rolling sum of residuals v_t^{OLS} over 36 months so as to capture at each point of the sample the cumulative deviation of a given asset price to its fundamental over a medium-term sample.

In addition to the two models presented heretofore, we also consider a model corresponding to the "statistical" approach where bubbles are defined as significant deviations from a trend. Most of the papers in the literature have relied on a statistical filter to decompose asset prices between trend and cycle. Goodhart and Hofmann (2008) define boom periods as a persistent deviation from the trend of more than 5% and lasting at least 12 months while Detken and Smets (2004) use a 10% threshold. Alessi and Detken (2011) and Bordo and Jeanne (2002) define the boom as a 1.75 and 1.3 standard deviation at least from the trend respectively. For Bordo and Landon-Lane (2013), the boom occurs if a 5 % increase in house prices (10% for stock prices) is followed by a 25% correction within two years. In Jordà et al. (2013, 2015), a bubble is identified when an increase in a given asset price of more than 1 standard deviation from the trend is followed by a correction of 15% at least, over a 3-year period. Our third model, statistical, identifies bubbles with a dummy taking the value 1 (resp. -1) when asset prices are more than 1.5 standard deviation superior (resp. inferior) to the Christiano-Fitzgerald (CF) trend and the value 0 when asset prices are within these bounds, so 87% of the data lies within them.[13] A synthetic description of all models is presented in Table 1.

2.2. Data

We estimate these three models for two asset prices: stocks and housing. Data are available from January 1986 to August 2016 (see Table A in appendix for data description and sources and Table B for descriptive statistics). The stock price index is the S&P500. We use Shiller's benchmark monthly index for house prices. Each asset price is deflated by the CPI.

For the cash-flow model, fundamental value is a function of cash-flows (dividends for stocks and rents for housing prices) and the discount factor. Dividends paid by corporations and rents received by households are provided by the Bureau of Economic Analysis. Data are available at a quarterly frequency. [14] The long-term sovereign interest rates are used as the discount factor. The model is extended to account for a time-varying risk-premium, using the VIX indicator – the Chicago board of trade volatility index –, which is often used as a proxy for uncertainty and market appetite for risk. Yet, this model may not fully account from all available information. We also identify the fundamental component with a model estimated on a large set of information, including macroeconomic and financial indicators such as: real disposable income, inflation, real GDP, industrial production, oil prices, the Conference Board consumer and ISM firm confidence indicators, the VIX, 3-month interbank interest rate, monetary (M2) and credit aggregates (credits granted by commercial banks). Table 1 provides descriptive statistics and the correlation structure for each market.

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[13] The main advantage of the CF filter compared to the Hodrick-Prescott (HP) filter is that the former is one-sided so that the estimation does not affect the last point of the sample.

[14] Quarterly data have been linearly interpolated to monthly frequency.

2.3. A Principal Component Analysis of bubbles

As emphasized earlier, there is no consensus on the appropriate method to identify bubbles. That is why we adopt an agnostic approach and consider that bubbles are captured by the common component of the three models described above. In order to summarize the information provided by the residuals ϵ_t^{OLS}, the rolling sum of residuals ν_t^{OLS} and the "statistical" model, we perform a Principal Component Analysis (PCA) to estimate a unique indicator maximizing the common variance of the individual bubble series. In addition to reduce information in one single series, another advantage of the PCA is to remove the evolution of each series that would be specific to that model and provides a robust measure of the bubble component of asset prices. More specifically, PCA seeks a linear combination such that the maximum variance is extracted from the variables. Components reflect both common and unique variance of the variables and may be seen as a variance-focused approach seeking to reproduce the total variable variance with all components and their correlations.[15]

We compute bubble indicators for each market (stock and housing) by estimating the first component of the 3 individual bubble components of each market. Table 1 provides the main characteristics of the estimation of the stock and housing bubble indicators. These two bubble indicators capture 58 and 57% of the variance of their respective 3 bubble components. Besides, the highest loading factor is on the "structural" approach (model 1) for the stock bubble whereas it is on the "econometric" approach (model 2) for the housing bubble.

The individual bubble components (for each approach) together with the bubble indicator (the first principal component) are plotted in Figure 1. The bubble indicator for stock prices coincides with the dummy of the statistical approach where bubbles (crashes) are identified as at least a 1.5 standard-deviation of prices above (below) their trend. The dotcom bubble is also clearly identified by the first principal component and the pure statistical approach. The bubble period would have started in 1999 and would have stopped in 2000. A bubble is also identified in 2007 followed by a crash in stock markets in late 2008. Turning to the housing market, the first principal component identifies a bubble in 2006 and the subsequent crash. This bubble indicator and the outcome of the statistical approach are clearly in line. Over the end of the sample, the bubble indicator is close to zero for both stock and housing markets. Table 1 also shows that the two bubble indicators are weakly correlated (0.13).

2.4. Robustness

The construction of the bubble indicator relies on assumptions related to the estimation methods for the "structural" and the "econometric" approaches, the rolling sum applied to residuals, the length of the rolling-window for the sum of residuals, the filtering method, and the smoothing parameters used for the "statistical" approach. We assess the robustness of the identification of these bubble indicators in section 4.5 by examining whether the impact of monetary shocks on asset price bubbles is affected by changes in the assumptions used to compute the bubble indicators.

To that end, we compute several alternative PCA bubble indicators for which one dimension at a time is modified relative to the baseline assumptions. First, we apply the same transformation on residuals ϵ_t^{OLS} and ν_t^{OLS}. We compute the rolling sum of residuals ϵ_t^{OLS} over

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[15] In practice, computing PCA of a dataset X, an (m × n) matrix, where m is the number of variables and n is the number of observations, entails computing the eigenvectors and eigenvalues of the covariance matrix of X. The eigenvector with the highest eigenvalue, measuring the variance in all variables which is accounted for by that eigenvector, is the first component.

36 months and estimate the first principal component (*PCA_rollingsum*) with the rolling-sum of residuals for both "structural" and "econometric" approaches plus the "statistical" model. Inversely, the PCA may also be computed on standard residuals from equations (4) and (5) plus the "statistical" model (*PCA_norollingsum*). We also estimate *PCA_roll24* and *PCA_roll48* bubble indicators where the rolling sum for residuals v_t^{OLS} of the "econometric" approach are computed over a 24-month and 48-month rolling-window.

Equation (4) and (5) may also be estimated with an error-correction model (ECM) to capture the possibility that prices are a combination of a long-run trend and short-run dynamics. Breitung and Kruse (2015) argue that the dynamics of the asset price and its fundamental should be dealt with jointly and also use an ECM. Therefore, the price of the financial asset is allowed to deviate temporary from its long-term equilibrium and the bubble components for the "structural" and "econometric" approaches are captured by the rolling sum of residuals and used to compute a *PCA_ECM* bubble indicator.

Besides, the estimation of the cash-flow model could be biased by endogeneity as the fundamental value is captured by cash-flows that might be driven by the asset prices. To account for this potential endogeneity issue, we estimate equation (4) with GMM, using industrial production and monthly interpolated GDP as instruments for cash flows, considering that current economic activity would provide information on current cash-flows. Moreover, another limitation of our analysis due to data availability is that we include contemporaneous cash-flows whereas the cash-flow model relies on the forward-looking nature of asset prices. This may introduce a bias in the measurement of bubbles. Consider the case where the central bank implements an expansionary monetary policy. Due to the transmission lags, the fundamentals do not immediately improve. However, rational investors anticipate an improvement of economic perspectives and henceforth a rise in future cash-flows, so the asset price increases. If our model fails to account for this rise in expected cash-flows, the increase in the asset price is mistakenly considered as a bubble. In order to account for expected cash-flows, we estimate equation (4) including the forward values of the respective cash-flows, 12 months and 36 months ahead. We estimate this model with GMM and use private and central bank output forecasts as instruments for future cash-flows. We acknowledge that the realized forward values are not expectations of these cash-flows. Yet, estimating equation (4) with forward values and GMM enables to assess, under some assumptions, the sensitivity of our baseline estimates to the forward-looking behaviour of investors. The corresponding PCA bubble indicators are named *PCA_GMM*, *PCA_GMM12*, and *PCA_GMM36*.

Regarding the "statistical" approach, the baseline estimation resorts to a cyclical component estimated with the CF-filter. A HP-filter may also be used as an alternative to disentangle between the trend and the cyclical component of asset prices. Using this filtering method, we build the *PCA_HP* bubble indicator. Besides, the calibration of the CF-filter also matters as it may capture either short-term cycle or medium-term cycle. In the baseline, the procedure filters out stochastic cycles for periods smaller than 18 months and higher than 96 months (8 years). Drehmann et al. (2012) characterize the length of cycles disentangling short-term and medium-term cycles for several indicators, including house and equity prices. They suggest that house and equity price medium-term cycles have a duration of 10 ½ years and 9 ½ years respectively. To account for a longer duration of cycles, we augment this range to 12 from 144 months to perform the PCA (*PCA_CF ≠smoothing*).

We also assess whether the estimation of the first principal components depends on the sample period considered. Therefore, we estimate the eigenvalues and eigenvectors of a PCA over

three subsamples (1986-1996, 1996-2006 and 2006-2016) and then predict the first principal component over the entire sample using the subsample estimates. We henceforth obtain three PCA bubble indicators (*PCA_pre96*, *PCA_9606* and *PCA_post06*). Finally, we estimate the first principal component of all 14 models, including the ones used in the baseline and all alternative models described above (*PCA _all*). Finally, we also test whether our results are driven by the PCA estimation by assessing the impact of monetary shocks on each bubble component of the three approaches separately.

3. The identification of monetary shocks

Analysing the effects of monetary policy requires addressing issues about the identification of exogenous monetary shocks. Several methods have been used in the empirical literature and may lead to some discrepancies in the responses to monetary policy shocks.[16] Our baseline choice is to resort to the Romer and Romer (2004)'s approach augmented following insights from the information friction literature. Concretely, the identification of shocks accounts for the information set of both policymakers and private agents. For robustness purposes, we also use alternative approaches based on high frequency event-study assumptions following Kuttner (2001) or based on the estimation of forward-looking Taylor rules where residuals are considered as the monetary innovation.

3.1. Baseline monetary shocks

The question of the most relevant identification strategy remains an open question. Empirical literature on monetary policy has often resorted to VAR models. However, timing assumptions in recursive identifications – reasonable for real variables and their sluggish reaction to shocks and low sampling frequency – are not relevant when applied to financial variables or fast-moving variables. There is indeed no rationale to suppose that some asset prices move faster than others. Romer and Romer (2004) regress the intended federal funds rate change on the information set of the monetary authority to purge endogenous responses to current and expected future economic developments.

Because of different information sets (Romer and Romer, 2000, or Blinder et al., 2008), the Romer and Romer (2004)'s identification approach may underestimate the extent to which market participants are able to predict future interest rate decisions. As discussed in Blanchard et al. (2013) and Ricco (2015), the presence of information frictions significantly modifies the identification problem. We propose an identification that combines insights from Romer and Romer (2004) and from the information frictions literature. We thus require the estimated monetary shocks to be orthogonal to both central bank's and private agents' information sets and to macro and financial market information.[17] In this context of imperfect information, the new information is only partially absorbed over time and, estimated surprises are likely to be a combination of both current and past structural shocks.

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[16] See Coibion (2012) for a discussion.

[17] The inclusion of both private and central bank forecasts in the regression model enables us to deal with three concerns. First, private agents and policymakers' information sets include a large number of variables. Forecasts have the advantage of encompassing rich information sets. Bernanke et al. (2005) show that a data-rich environment approach modifies the identification of monetary shocks. Forecasts work as a FAVAR model as they summarise a large variety of macroeconomic variables as well as their expected evolutions. Second, forecasts are real-time data. Private agents and policymakers base their decisions on their information set in real-time, not on ex-post revised data. Orphanides (2001, 2003) show that Taylor rule-type reaction functions estimated on revised data produce different outcomes when using real-time data. Third, private agents and policymakers are mechanically incorporating information about the current state of the economy and anticipate future macroeconomic conditions in their forecasts and we need to correct for their forward-looking information set.

Our baseline measures of exogenous monetary shocks are based on the shadow rate measure of Wu and Xia (2016). Because monetary policy has taken many different dimensions over the last years and we ought to consider shocks to unconventional instruments and communication policies (forward guidance, for instance) in addition to shocks to the conventional instrument, we use this shadow rate measure that translates these various dimensions in a single variable expressed in interest rate space to measure the overall stance of monetary policy (labelled *MP*). In a second step, we also estimate shocks specific to the policy rate (labelled *PR*) and to an indicator of the central bank balance sheet size (labelled *Unconv*) to further analyse the response of bubble indicators to conventional and unconventional policies in normal and exceptional times respectively. For the former, we consider the federal funds target rate. For the latter, we consider the monthly change in the size of the Federal Reserve's balance sheet.

The baseline shock to the overall monetary stance and its two alternatives are estimated with the following equations from which we extract the residuals:

$$\Delta i_t = \beta_0 + \beta_1 \, i_{t-1} + \beta_2 \, \Omega_t + \beta_3 \, \Psi_t + \beta_4 \, X_{t-1} + \beta_5 \, Z_t + \epsilon'_t \qquad (6)$$

$$\epsilon'_t = \beta_6 + \beta_7 \epsilon'_{t-1} + \epsilon_{rr,t} \qquad (7)$$

where i_t is the monetary policy instrument. We assume that the monetary shock must be orthogonal to the contemporaneous policymakers' information set Ω_t, to the private agents' one Ψ_t, to lagged financial market variables embedded in X_{t-1}, and to a vector Z_t of contemporaneous macroeconomic variables. This shock to the overall monetary policy stance, $\epsilon_{rr,t}$, is labelled MP-Shocks-RR ($\epsilon^i_{rr,t}$ for the shock to the policy rate, labelled *PR*, and $\epsilon^{Unc}_{rr,t}$ for the shock to unconventional policies, labelled *Unconv*). A consequence of the timing of the right-hand-side vectors in equation (6) is that monetary shocks affect contemporaneously financial market variables, but do not affect contemporaneously central bank's and private agents' information sets or macroeconomic variables. We believe that the opposite assumptions that monetary shocks are only based on past data or that they do not influence financial markets in real-time are fragile.[18] Because of potential information frictions, ϵ'_t is made orthogonal to ϵ'_{t-1} (equation 7) so that estimated residuals do not contain information from past structural shocks.

The policymakers' information set Ω_t comprises the level and change in FOMC inflation and output projections for current and next calendar years, Ψ_t includes the level and change in US SPF inflation forecasts for 1, 2 and 5 years ahead (resp. next quarter and next year), X_t contains the VIX and the oil price growth rate, and Z_t comprises current and lagged values of the inflation rate, industrial production and the monthly-interpolated real GDP growth rate. We use central bank macroeconomic forecasts (FOMC projections) and private ones (the US Survey of Professional Forecasters, SPF).[19]

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[18] One could argue that there may also be information frictions in financial markets and that financial variables in *t-1* do not incorporate information news from *t-2*, *t-3*, etc. We control for this by estimating equation (6) with two additional lags. The correlation between this alternative series and the baseline is 0.99. These estimates are available from the authors upon request.

[19] The FOMC publishes forecasts for inflation and real GDP growth twice a year in the Monetary Policy Report to the Congress since 1979. Since October 2007, their publication is quarterly. We consider forecasts of the Personal Consumption Expenditures (PCE) measure of inflation and real GDP. We consider the midpoint of the "full range" of all individual FOMC members' forecasts. FOMC projections have been constant-interpolated to monthly frequency. We assess the robustness of our identification with Greenbook projections. The alternative series of monetary shocks using Greenbook projections instead of FOMC projections has a 0.91 correlation with the baseline series. The SPF is collected from around 40 panellists and published by the Federal Reserve Bank of Philadelphia. SPF CPI forecasts are provided as year-over-year percent changes. We consider the median of individual responses.

3.2. Alternative measures of monetary shocks

A first alternative is to follow Kuttner (2001)'s high frequency methodology to identify monetary shocks using changes in the price of futures contracts. Kuttner (2001) identifies monetary surprises by accounting for the forward-looking nature of financial data. For a monetary policy event on day d of the month m, the monetary shock can be derived from the variation in the rate implied by current-month futures contracts on that day. The price of the future being computed as the average monthly rate, the change in the futures rate must be augmented by a factor related to the number of days in the month affected by the change:

$$\epsilon_{kutt,t} = \frac{D}{D-d}(f^0_{m,d} - f^0_{m,d-1}) \tag{8}$$

$\epsilon_{kutt,t}$ is the unexpected interest rate variation which constitutes a monetary shock, $f^0_{m,d}$ is the current-month futures rate and D is the number of days in the month and d the day of the decision. One issue with the Kuttner measure is that it focuses on futures contracts about interest rate only. However, monetary policy has taken many different dimensions over the last years and Wu and Xia (2016) have proposed shadow rate measures that capture the different dimensions of monetary policy in a single variable expressed in interest rate space. However, their measure has a monthly frequency. Krippner (2013, 2014) has estimated shadow short rate (SSR) series at the daily frequency and it therefore enables to apply the Kuttner's high frequency event-study identification of monetary surprises to the daily variation in SSR_t on the policy announcement day:

$$\epsilon_{kripp,t} = SSR_t - SSR_{t-1} \tag{9}$$

Because shadow rate measures are not calendar-based instruments like fed funds futures, there is no need for an adjustment for the remaining number of days. These shocks (labelled MP-Shocks-HF) rely on the financial market participants' interpretation of the overall monetary news disclosed that day, and include private reactions to central bank conventional or unconventional decisions, and to central bank communication released at the same time.

A second alternative for identifying monetary surprises is to estimate a forward looking Taylor rule equation augmented with oil prices and a financial stress index (included in the vector Y_t). This equation is estimated over the full sample. The monetary policy shock (labelled MP-Shocks-TR) is then the residuals of the following equation:

$$i_t = \beta_8 + \beta_9 i_{t-1} + \beta_{10}\pi_{t+6} + \beta_{11}og_{t+6} + \beta_{12}Y_{t+6} + \epsilon_{tr,t} \tag{10}$$

Figure 2 plots the shocks to the overall monetary policy stance using the baseline approach following Romer and Romer (2004) and the two alternative shock series described in this subsection. Table C in the Appendix provides descriptive statistics for these different monetary shocks and their correlation.

When investigating the effect of monetary policy, estimated monetary shocks are expected to fulfil certain conditions. First, we assess the normality and autocorrelation of the estimated shock series. Table C in the Appendix provides the outcomes of these standard tests. These results call for putting less emphasis on Taylor rule type shocks as these shocks exhibit auto-correlation. Second, for our estimated series of monetary shocks to be relevant, they need to be unpredictable from movements in data. We assess the predictability of the estimated shock series with Granger-causality type tests using 9 macroeconomic and financial variables. Table C in the Appendix shows the adjusted R^2 and F-stats of an OLS estimation that aims to test the null hypothesis that estimated series of exogenous shocks are not predictable. It shows that the Romer-Romer-type shock series are relevant to be used in our second-stage estimations.

4. The effect of monetary policy on asset price bubbles

As emphasized previously, there is no consensus in the theoretical literature about the effect of monetary policy on asset price bubbles. In the "leaning against the wind" approach, the link between monetary policy and bubbles is negative so that a restrictive monetary policy shock would reduce the size of bubbles. However, Galí (2014) shows that a restrictive monetary policy would increase the size of bubbles. The empirical strategy aims to disentangle between these two possible responses of asset price bubbles to monetary shocks. To that end, we assess the impact of monetary shocks on asset price bubbles with the local projection method proposed by Jordà (2005). This method is flexible and may easily account for asymmetries in the transmission of monetary policy, and allow us to disentangle the impact of restrictive and expansionary monetary shocks.[20] We also assess the impact of monetary shocks depending on whether the instrument is the policy rate or balance sheet measures.

4.1. Empirical strategy

Assessing the impact of shocks on a given economic variable may be realized either through VAR models or with Jordà's Local Projection method. Whereas the first method enables to take into account the intertwined dynamics of a set of variables, it may be imposing excessive restrictions on the endogenous dynamics and may be prone to bias if the model is misspecified. The second method offers more flexibility in the estimation. In linear stationary settings, the out-of-sample forecasting performance of VARs and local projections is quite similar (see Marcellino et al., 2006, and Kilian and Kim, 2011). However, because a linear low-order autoregressive representation of the data generating process of macroeconomic time series may be deceptive, the robustness of local projections to model misspecification and non-linearity makes them an appealing procedure to recover dynamic responses to exogenous shocks. Considering that these exogenous structural shocks are identified *ex ante*, Jordà (2005) suggests estimating a set of *h* regressions representing the impulse response of the dependent variable at the horizon *h* to a given shock at time *t*:

$$y_{t+h} = \alpha_h + \beta_h \epsilon_t + \sum_{i=1}^{2} \phi_{h,i} \cdot y_{t-i} + \eta_{t+h} \qquad (11)$$

where y_{t+h} is the dependent variable – the bubble indicator - at the horizon h, ϵ_t represents the monetary shock, either to the overall policy stance or to conventional and unconventional measures specifically, and y_{t-i} are lags of the dependent variable (that we set to 2 based on the non-significance of additional lags). We set *h* to 24 periods to measure the effect of monetary shocks on bubble indicators over 2 years. Finally, because the estimated monetary shocks are generated regressors that might cause biased standard errors, we compute heteroskedasticity and autocorrelation robust Newey-West standard errors assuming that the autocorrelation dies out after three lags.[21] This correction also enables to control for potential heteroskedasticity and auto-correlation of the residuals.

Furthermore, we investigate if expansionary and restrictive monetary shocks have different effects and thus if a linear framework may bias the outcome. Such an asymmetric response is

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[20] Cover (1992) illustrates the differentiated impact of expansionary and restrictive monetary policy shocks on output. His results suggest that expansionary shocks would have limited effects while restrictive shocks would be more powerful. These results have been challenged by Wise (1999) and remain discussed (see Angrist et al., 2017, or Tenreyro and Thwaites, 2016).

[21] We have also computed standard errors robust to misspecification using the Huber-White-sandwich estimator and they provide smaller confidence intervals around the point estimate. This generated regressor issue is common to all empirical studies estimating exogenous shocks in a first step as in Romer and Romer (2004), but is more acute when the generated regressors are not normally distributed.

highlighted by the literature analysing the impact of monetary policy on output, suggesting that restrictive monetary shocks would be more powerful than expansionary shocks. Equation (11) may be modified to account for non-linearities:

$$y_{t+h} = \alpha_h + \beta_{I,h}(\epsilon_t.\Lambda_t) + \beta_{\epsilon,h}\epsilon_t + \beta_{\Lambda,h}\Lambda_t + \sum_{i=1}^{2} \phi_{h,i}.y_{t-i} + \eta_{t+h} \qquad (12)$$

where Λ_t is a dummy variable for expansionary monetary policy shocks. This specification aims to single out the potential asymmetric effects of restrictive ($\beta_{\epsilon,h}$) and expansionary ($\beta_{\epsilon,h} + \beta_{I,h}$) monetary shocks on bubble indicators.

4.2. Linear and non-linear evidence

The linear responses over 24 months of stock and housing bubbles to the baseline monetary shocks (i.e. MP-Shocks-RR) are plotted in the upper-panel of Figure 3. Point estimates are shown with 1 and 2 standard-errors confidence intervals. We find that the stock market bubble reacts positively to the restrictive monetary shock starting from the 14th month whereas there is no evidence of monetary shocks affecting the housing price bubble indicator.

The lower panel of figure 3 plots the non-linear effects of monetary policy. For comparison, the linear response of bubble indicators is also represented together with the non-linear responses. The response of the stock bubble indicator to a restrictive shock is not significant whereas the response to an expansionary shock is positive. This effect is significantly different from zero from the 14th to the 20th month, and significantly different from the linear response after the 8th month. Monetary shocks account for around 3% of the variance of the stock bubble component.[22] The linear effect shown in the upper panel of Figure 3 therefore seems to stem from the assumption that the impacts of restrictive and expansionary shocks are symmetric. The asymmetric result however suggests that restrictive monetary policy has no effect on stock bubbles, but expansionary monetary policy would inflate them. The non-linear responses of the housing bubble indicator to restrictive and expansionary shocks are not significant.

The linear and non-linear responses show that the effects of monetary shocks on asset price bubbles are not symmetric. It calls for differentiating the responses to restrictive and expansionary shocks. Thus, when the monetary policy stance is measured by a shadow rate encompassing standard and unconventional monetary policy measures, restrictive monetary policy would neither be able to deflate stock price bubbles as proposed by the "leaning against the wind" literature nor inflate them as predicted by rational bubble models. However, expansionary monetary policy may contribute to increase stock market bubbles as illustrated by the responses of the bubble indicator.

4.3. Does the instrument matter?

The previous results are based on a proxy (the shadow rate) measuring the overall monetary policy stance. However, the effect of standard monetary policy may differ from the effect of unconventional monetary measures. In the policy debate, concerns have emerged about potential adverse effects of unconventional measures and notably quantitative easing policies, which may have fuelled asset price bubbles. This debate can be seen as a variation of the debate on the risks associated to periods of prolonged expansionary policy when policy rates remain "too low for too long". In reaction to this, it has been advocated that a "leaning against the wind" policy would help mitigating asset price bubbles. These issues can be dealt with by

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[22] We compute the variance decomposition using partial R² that indicates the fraction of the improvement in R² that is contributed by the excluded covariate.

disentangling the effects of shocks to the policy rate before July 2008 and shocks to unconventional policies since July 2008. To this end, we estimate equation (12) which accounts for the asymmetric effects of monetary policy, either with the policy rate or with a measure for unconventional monetary policy instead of the shadow rate.

We investigate the impact of interest rate policies by raising two questions. First, do expansionary interest rate shocks fuel stock bubbles? Second, are exogenous interest rate increases effective in deflating stock price bubbles as claimed by proponents of the "leaning against the wind"? To that end, we adjust equation (12) in two respects: first, we replace the shock to the overall stance of monetary policy, ϵ_t, by the exogenous shock to the conventional policy rate, ϵ_t^i, and second, we estimate equation (13) on a subsample ending in June 2008.[23]

$$y_{t+h} = \alpha_h + \beta_{I,h}(\epsilon_t^i.\Lambda_t) + \beta_{\epsilon h}\epsilon_t^i + \beta_{\Lambda,h}\Lambda_t + \sum_{i=1}^{2} \phi_{h,i}.y_{t-i} + \eta_{t+h} \qquad (13)$$

The upper left-hand side of Figure 4 plots the responses of the stock market bubble to restrictive shocks to the policy rate. Results suggest that restrictive interest rate shocks would still have no effect on the stock market bubble confirming the previous results when we consider the shadow rate as the instrument of monetary policy. The case for the "leaning against the wind" approach seems fragile as we find no evidence that restrictive monetary policy would deflate bubbles. It may also be noticed that adopting a more restrictive monetary policy stance to deflate bubbles would not trigger the adverse effects on stock market bubble emphasized by the rational bubble models *à la* Galí (2014).

The lower left-hand side of Figure 4 plots the responses of the stock market bubble to expansionary interest rate shocks. It provides an assessment of the risks associated with expansionary monetary policy in normal times when central banks use the policy rate as the main policy instrument. The stock market bubble responds positively to an expansionary interest rate shock. There seems to be risks associated with low interest rates but they would be circumvented to the stock market.

Finally, we analyse the effect of unconventional monetary shocks identified after July 2008. We adjust the estimation of equation (12) in two respects: first, we replace the exogenous shock to the overall stance of monetary policy, ϵ_t, by the exogenous shock to unconventional policies, ϵ_t^{Unc}, and second, we estimate equation (14) on a subsample starting in July 2008.

$$y_{t+h} = \alpha_h + \beta_{I,h}(\epsilon_t^{Unc}.\Lambda_t) + \beta_{\epsilon,h}\epsilon_t^{Unc} + \beta_{\Lambda,h}\Lambda_t + \sum_{i=1}^{2} \phi_{h,i}.y_{t-i} + \eta_{t+h} \qquad (14)$$

where Λ_t is a dummy variable for expansionary monetary policy shocks. The identification of shocks is performed with the method described in section 3 for the post-2008 subsample and considering unconventional policies only. The right-hand side of Figure 4 illustrates the effect of restrictive and expansionary shocks to the Federal Reserve's balance sheet. Restrictive monetary policy has no effect consistent with earlier evidence. The response of the stock market bubble to expansionary shocks indicates that unconventional measures do not feed bubbles. The response is even negative around 5 months and after 19 months indicating that the implementation of Federal Reserve's QE would have reduced the bubble component of stock prices, as suggested by the rational bubble model *à la* Galí (2014).[24] So the risk of a stock price bubble related to unconventional policies does not materialize in the data. The concerns

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[23] We estimate equation (13) for the stock bubble indicator only based on the absence of significant results for housing bubbles in the previous section. Estimates of equation (13) for the housing market are again not significant and available upon request from the authors. Further theoretical research may focus on making clearer why there is a different transmission of monetary policy to the bubble component of stock and housing markets.

[24] Although Gali (2014) considers the impact of conventional interest rate policies, the present finding suggests that his conclusion might apply to the effect of unconventional policies.

raised by Borio and Zabai (2016) about the side-effects of QE are not confirmed when focusing on the bubble component of stock prices.

4.4. An application to the euro area

We estimate the effects of monetary shocks on asset price bubbles in the EA using the same method for both the identification of bubble indicators and monetary shocks as for the US. Data are available from January 1999 to June 2016 for the EA. The stock price index is the Eurostoxx (listing the largest 295 firms). House prices data stem from a quarterly index for residential property prices calculated by the European Central Bank (ECB). Considering the cash-flow model, we use quarterly dividends paid by financial and non-financial corporations and quarterly rents received by households are available from Eurostat for the five biggest EA countries. The long-term interest rate – from benchmark government bonds – is used as the discount factor. The CISS is used to account for a time-varying risk-premium. We also identify the fundamental component with a model estimated on a large set of information, including macroeconomic and financial indicators such as: real disposable income, inflation, real GDP, industrial production, oil prices, European Commission confidence indicators for household and industry, a financial stress indicator (CISS), 3-month interbank interest rate, M3 and credit aggregates (credit counterparties of monetary aggregate). For the identification of monetary shocks, we use the ECB/Eurosystem staff macroeconomic projections and the ECB's SPF.

Figure 5 plots the linear and non-linear bubble responses in the EA. The linear responses of both stock and housing bubble indicators are presented in the upper-panel and are not significant. The lower panel illustrates the decomposition between restrictive and expansionary shocks. Restrictive shocks do not influence asset price bubbles whereas expansionary shocks affect positively stock and housing price bubbles. Evidence using EA data confirms that the effects of monetary shocks on asset price bubbles are asymmetric. They also show that restrictive monetary policies would not be able to deflate asset price bubbles. Finally, these results confirm that expansionary shocks may fuel bubbles.

4.5. Sensitivity analysis

We assess the sensitivity of our baseline results to alternative assumptions at each step of the empirical methodology. First, we test whether the results are driven by the model-averaging feature of the PCA, which may aggregate individual bubble components with different reactions to monetary shocks. We test this hypothesis by estimating the effect of monetary shocks on each individual bubble component – the "structural", "econometric" and "statistical" approaches. More precisely, we estimate equations (11) and (12) separately for ϵ_t^{OLS}, the rolling-sum of v_t^{OLS} and the dummy variable. Second, we test whether alternative measures of the PCA bubble indicator, as described in section 2.4, react differently to monetary shocks. Third, we test whether our results are sensitive to the identification of monetary shocks as described in section 3.2. Fourth, we examine whether alternative specifications of the Local Projection model alter the results. We modify equations (11) and (12) in four ways: without any controls, with the contemporaneous bubble indicator of the other asset, with forward values (at the same horizon *h*) of the bubble indicator of the other asset, and with the level of the policy variable and lags of the monetary shock (as proposed by Ramey, 2016).

Robustness tests for the linear and the non-linear specification are available in Figures A and B in the appendix. These figures plot the responses of all alternatives described above together with the baseline response with its corresponding confidence interval for each market. We find that the alternative responses do not differ significantly from the baseline response. The main

results of this paper are strongly confirmed. First, the effects of monetary policy are asymmetric. Second, restrictive monetary policy is not able to deflate asset price bubbles whereas expansionary policies do fuel stock market bubbles. These two results are robust to individual bubble measures, alternative PCA bubble indicators, different procedures to estimate monetary shocks and several specifications of the local projection equation.

5. Conclusion

Because financial stability is now an objective for central bank policymakers, the issue of whether monetary policy may help achieving this goal is crucial and notably hinges on the effect of monetary policy instruments on asset prices. Yet, it must be reminded that the reaction of asset prices is also part of the transmission channel of monetary policy. Consequently, central banks need to know if asset price movements are desirable or when monetary policy has negative side-effects. This paper deals with this issue and assesses the impact of monetary policy shocks on asset price bubbles. To this end, we propose a novel approach to identify bubbles on stock and housing markets based on a range of bubble models usually used in the literature. As none of existing models have reached consensus, we develop a new bubble indicator based on a model-averaging approach.

A key message of this paper is that the response of bubbles is not symmetric. The responses to restrictive and expansionary shocks in empirical analyses should be differentiated. This calls for more theoretical research on the transmission channels of monetary policy to bubbles to account for these asymmetries. The main result of the paper is that restrictive policies are not able to deflate asset price bubbles whereas expansionary policies inflate them. Besides, the effect of expansionary monetary policy also depends on the policy instrument. Our results suggest that the risk that quantitative easing would inflate asset price bubbles does not materialize in the data. However, interest rates cuts – standard expansionary monetary policy – would have significant and positive effects on stock price bubbles.

From a policy perspective, central banks should be aware that there is a risk for lower policy rate to inflate stock price bubbles. However, due to the asymmetric response of bubble components, monetary policy would not be the right tool to try to deflate them. Financial instability may not be seen only through asset *price* bubbles, so that there may be a role for monetary policy through its effect on credit to dampen financial risks together with macroprudential policies.

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Figure 1. Asset price bubbles

Stock market

Housing market
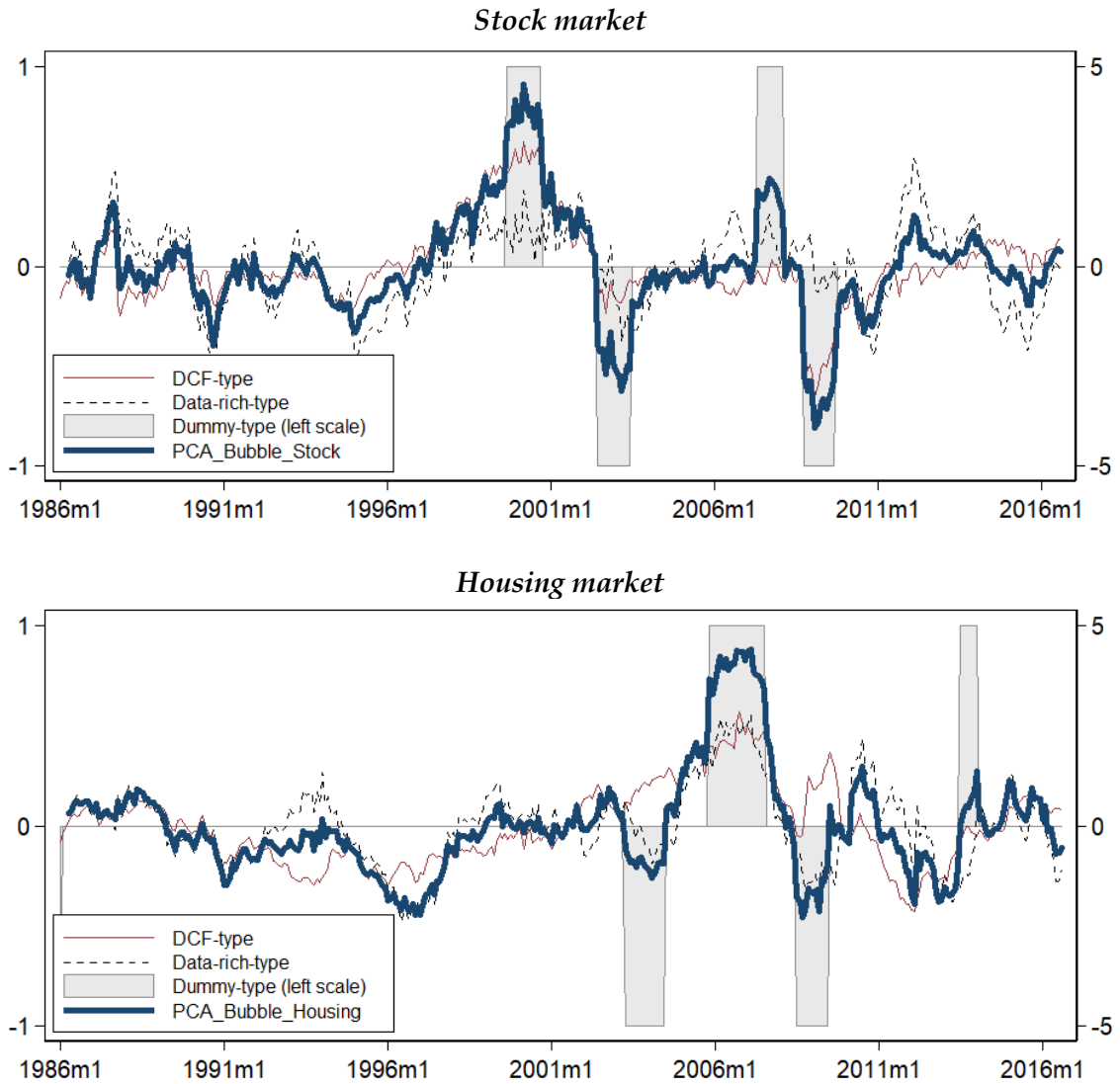
Note: authors' estimations described in section 2. The thin red line corresponds to the individual bubble series of model 1 (the structural one). and econometric models. The thin black line corresponds to the individual bubble series of model 2 (the data-rich one). The grey area corresponds to the dummy variable of model 3 (the statistical approach). The thick blue line plots the first principal component of the three preceding series. All continuous series are normalized. The right scale is expressed in standard deviation of the corresponding series.

Figure 2. Monetary shocks
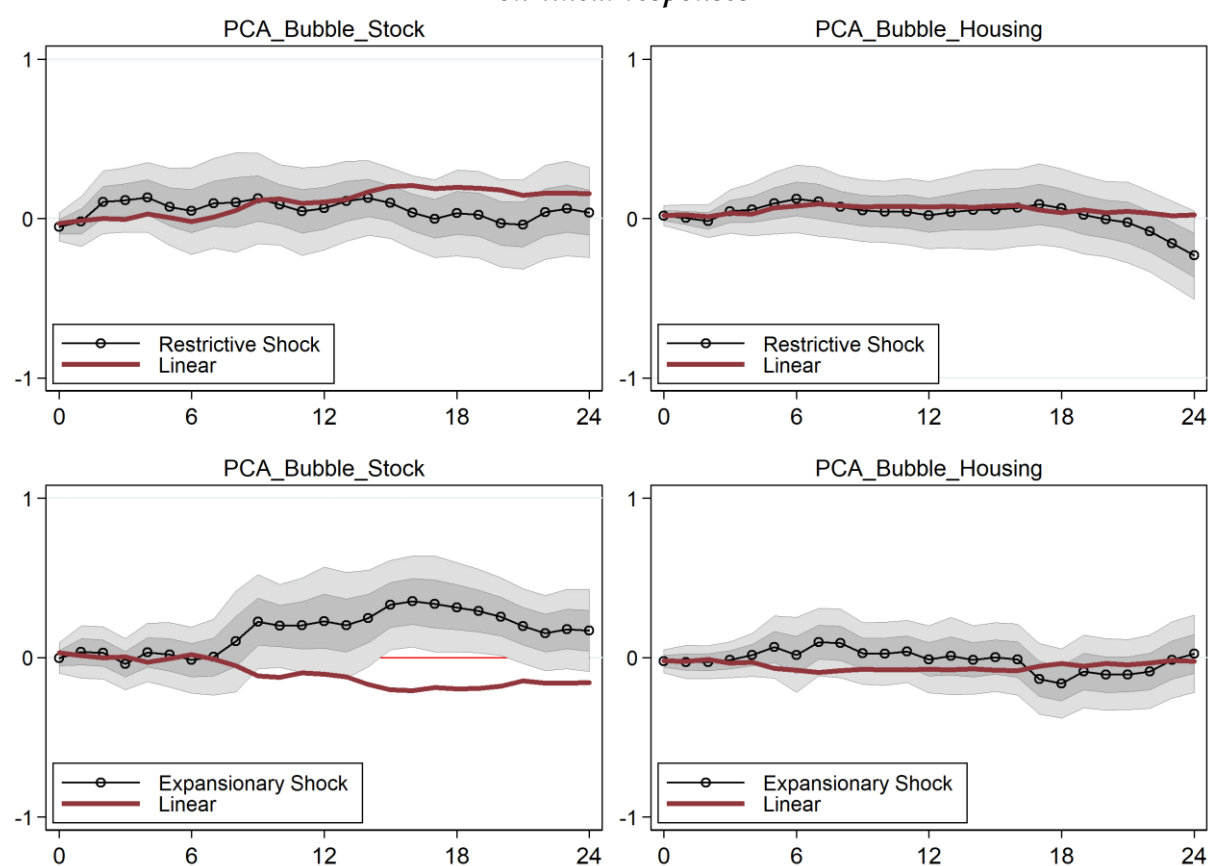
Overall monetary stance (shadow rate)
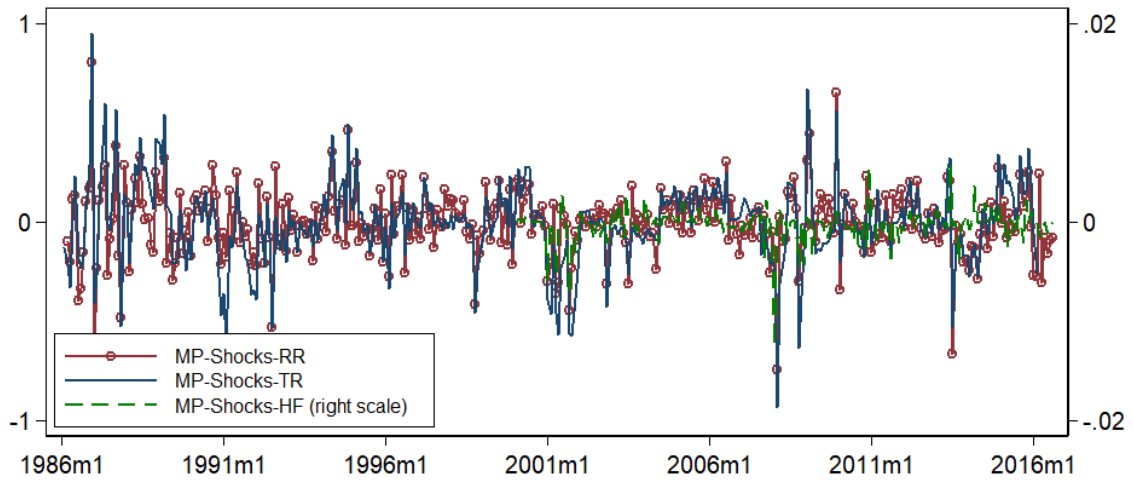
Note: Romer and Romer (RR), high-frequency event study (HF) and Taylor rule-type monetary shocks are computed by estimating equations (9)-(10), (12) and (13) respectively as described in section 3. The unit on both the left and right scale is in percentage points.

Figure 3. Linear and non-linear bubble responses

Linear responses to a restrictive (positive) shock to the overall monetary stance
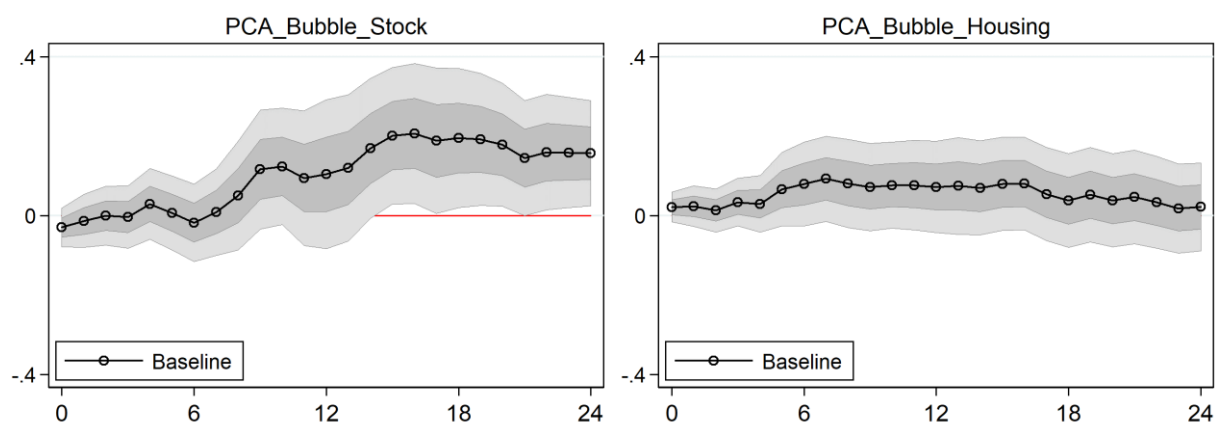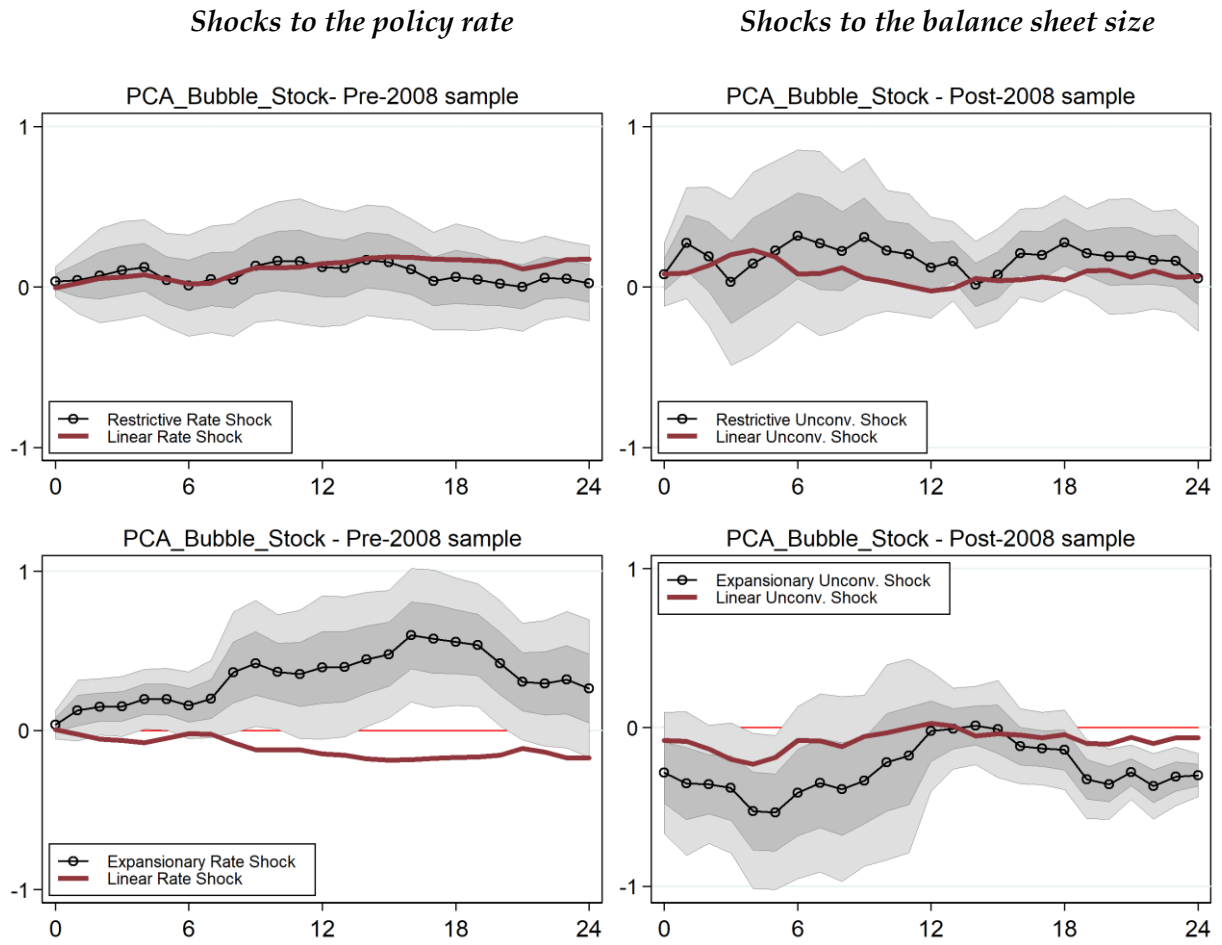

Non-linear responses

Note: Linear and non-linear responses correspond to estimates of equations (11) and (12) respectively. Shaded area represents the 1 and 2 standard errors confidence interval of the response of the baseline shock (MP-Shocks-RR). The y axis is expressed in terms of standard deviation of the bubble indicators. The monetary shock corresponds to a 1 standard deviation increase in the shock series.

Figure 4. Non-linear stock bubble responses to different policy instruments

Shocks to the policy rate   *Shocks to the balance sheet size*

Note: Left-hand side and right-hand side responses correspond to estimates of equation (13) and (14). Shaded area represents the 1 and 2 standard errors confidence interval around the non-linear response. The y axis is expressed in terms of standard deviation of the bubble indicators. The monetary shock corresponds to a 1 standard deviation increase in the shock series.
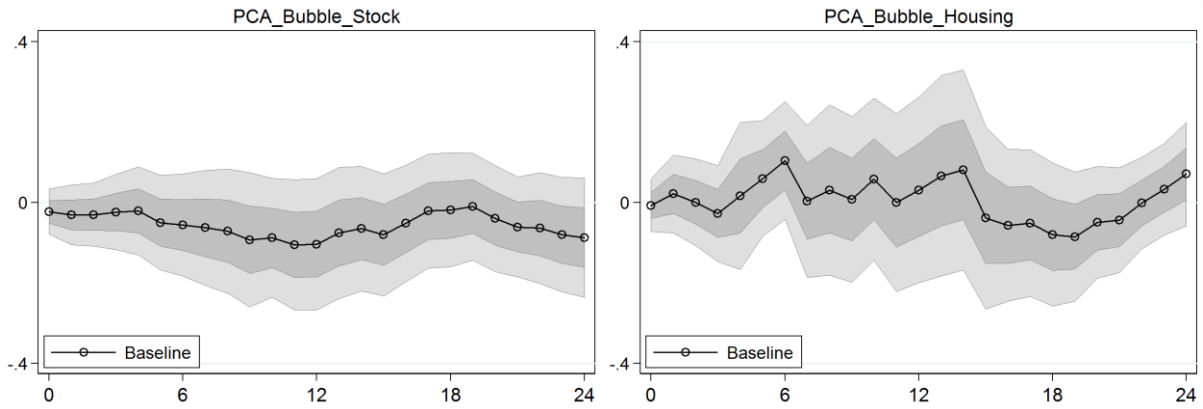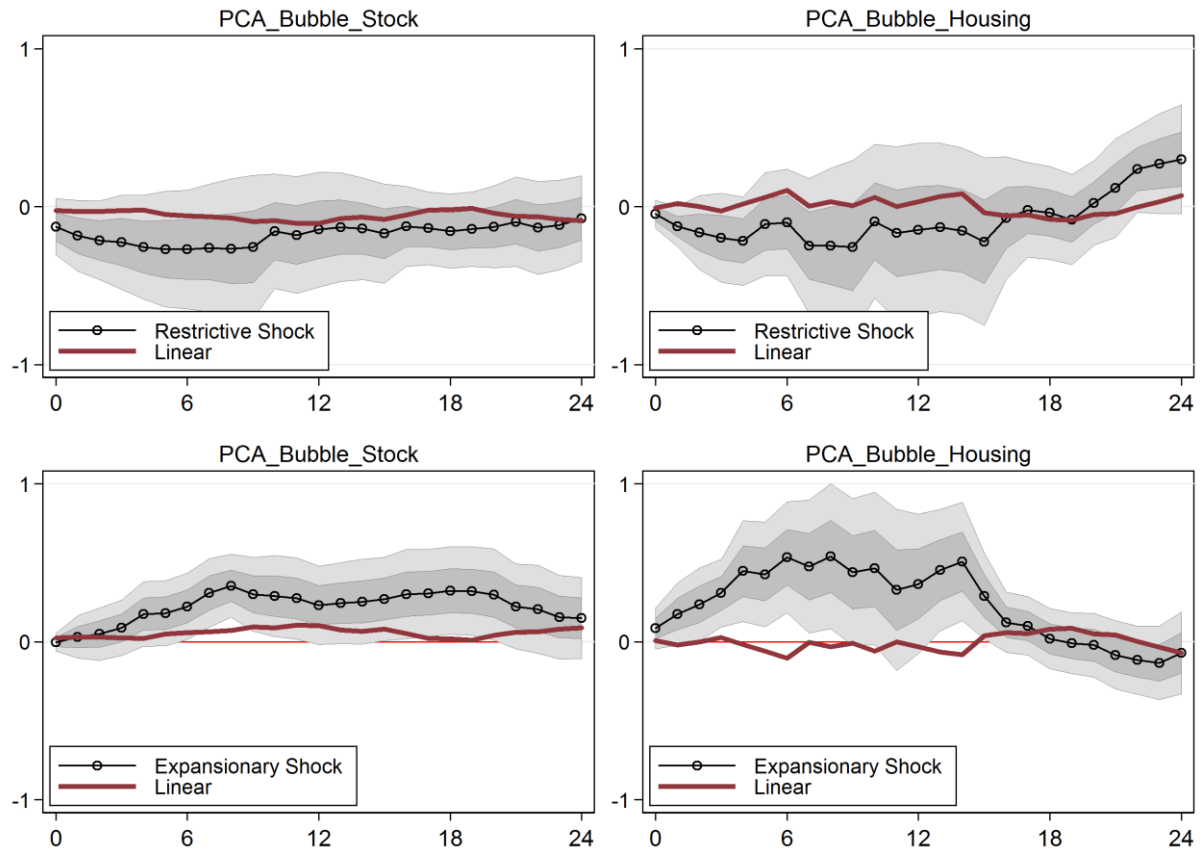
Figure 5. Linear and non-linear bubble responses in the euro area

Linear responses to a restrictive (positive) shock to the overall monetary stance

Non-linear responses

Note: Linear and non-linear responses correspond to estimates of equations (11) and (12) respectively. Shaded area represents the 1 and 2 standard errors confidence interval of the response of the baseline shock (MP-Shocks-RR). The y axis is expressed in terms of standard deviation of the bubble indicators. The monetary shock corresponds to a 1 standard deviation increase in the shock series.

Table 1. Range of individual bubble models & PCA estimation

ID	Approach	Estimation Method	Bubble identification
model 1	Cash-flow	OLS	Residuals
model 2	Data-rich	OLS	Rolling 3y sum of residuals
model 3	Statistical	Christiano-Fitzgerald filter	1.5 standard deviation deviation from trend

p-values for autocorrelation and white noise tests			
Stock market			
		ε^{OLS}	v^{OLS}
Cumby-Huizinga test for autocorrelation		0.00	0.63
Portmanteau test for white noise		0.00	0.85
Housing market			
		ε^{OLS}	v^{OLS}
Cumby-Huizinga test for autocorrelation		0.00	0.50
Portmanteau test for white noise		0.00	0.00

Correlation structure between individual bubble series			
Stock market			
	model 1	model 2	model 3
model 1	1		
model 2	0.33	1	
model 3	0.54	0.20	1
Housing market			
	model 1	model 2	model 3
model 1	1		
model 2	0.47	1	
model 3	0.15	0.41	1

Principal component/correlation (Rotation: unrotated - Obs = 365)			
	Eigenvalue	Proportion	KMO stat
PCA_Bubble_Stock	1.74	0.58	0.57
PCA_Bubble_Housing	1.70	0.57	0.54

PCA scoring coefficients (eigenvectors)			
Variable	model 1	model 2	model 3
PCA_Bubble_Stock	0.65	0.46	0.61
PCA_Bubble_Housing	0.55	0.66	0.51

Correlation (PCA_Bubble_Stock,PCA_Bubble_Housing) = **0.13**

Note: The KMO stat is the Kaiser-Meyer-Olkin measure of sampling adequacy.

APPENDIX
NOT FOR PUBLICATION

Table A. Data sources and Description

	Name	Description	Source	Frequency
Asset prices				
Stock	sp500	S&P 500	Datastream	Monthly
Housing	housep	Shiller Index	Shiller	Monthly
Cash-flow model				
Dividends	divid_rsa	Paid dividends by corporations	BEA	Quarterly
Rents	rent	Rents received by households	BEA	Quarterly
Discount factor	txlg	10 year treasury bond interest rates	Datastream	Monthly
Risk premium	vix	Volatility Index	CBOE	Monthly
Data-rich information				
Income	rdb	Real disposable income	BEA	Quarterly
Real GDP	gdp	Real GDP	BEA	Quarterly
IndPro	ipi	Industrial production	BEA	Monthly
Oil prices	oil	Oil prices	Datastream	Monthly
Inflation	inf	Inflation	BEA	Monthly
Confidence indicators	csind & cscons	Confidence indicators for consumers and firms	Conference Board & ISM	Monthly
Financial stress	kcfsi	Kansas City Financial indicator	FRED	Monthly
Monetary Aggregate	m2	M2	Datastream	Monthly
Credit Aggregate	credit	Credits granted by commercial banks	Datastream	Monthly
Monetary policy				
Policy rate	fedfunds	Effective FFR	Federal Res.	Monthly
Target rate	fedtarget	FFR target	Federal Res.	Monthly
shadow rate	wu&xia	Shadow rate	Wu-Xia (2015)	Monthly
shadow rate	krippner	Shadow rate	Krippner (2016)	Daily
Unconventional measures	unconv	Table H.4.1 Fed's total assets	Federal Reserve	Monthly

Note: All nominal variables are deflated by the CPI.

Table B. Descriptive statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
sp500_r	368	6.36	0.47	5.36	7.11
housep_r	368	4.16	0.16	3.96	4.52
divid_rsa	368	4.30	0.44	3.43	4.96
txlg_r	368	2.59	1.70	-1.85	6.03
rent_r	368	3.40	0.82	1.62	4.59
rdb_r	368	7.25	0.20	6.88	7.58
ipi	368	-0.17	0.20	-0.56	0.06
gdp	368	8.29	0.23	7.86	8.63
inf	368	2.64	1.37	-1.96	6.38
m2_r	368	8.21	0.12	8.01	8.49
credit_r	368	7.98	0.33	7.46	8.54
csind	368	52.04	4.85	33.10	61.40
cscons	368	92.51	25.61	25.30	144.70
oil_r	368	2.96	0.54	1.82	4.11
vix	368	2.96	0.36	2.33	4.18
wu&xia	359	3.43	3.23	-2.99	9.85
krippner	368	3.05	3.63	-5.37	9.85

Table C. Properties of estimated monetary shocks

Descriptive statistics					
Variable	Obs	Mean	Std. Dev.	Min	Max
mpshock_rr	366	0.00	0.18	-0.74	0.80
mpshock_tr	361	0.00	0.21	-0.93	0.95
mpshock_hf	200	0.00	0.00	-0.01	0.01
prshock_rr	267	0.00	0.14	-0.65	0.47
uncshock_rr	97	0.00	0.04	-0.14	0.16

Correlation					
	mpshock_rr	mpshock_tr	mpshock_hf	prshock_rr	uncshock_rr
mpshock_rr	1				
mpshock_tr	0.84	1			
mpshock_hf	0.25	0.33	1		
prshock_rr	0.81	0.68	0.21	1	
uncshock_rr	-0.15	-0.12	-0.03	.	1

Shapiro-Francia normality test					
Variable	Obs	W'	V'	z	Prob>z
mpshock_rr	366	0.94	15.78	5.93	0.00
mpshock_tr	361	0.95	13.47	5.59	0.00
mpshock_hf	200	0.79	34.51	7.32	0.00
prshock_rr	267	0.97	5.73	3.68	0.00
uncshock_rr	97	0.87	11.56	4.83	0.00

| Autocorrelation test		Predictability of exogenous shock series			
	AR(1) coef.		F-stat	p-value	Adjusted R^2
mpshock_rr	0.00	mpshock_rr	0.69	0.73	-0.01
mpshock_tr	0.45***	mpshock_tr	3.81	0.00	0.07
mpshock_hf	0.04	mpshock_hf	2.24	0.02	0.06
prshock_rr	0.01	prshock_rr	2.74	0.00	0.06
uncshock_rr	-0.03	uncshock_rr	0.49	0.89	-0.06

Note: The vector of variables for predictability tests includes lagged values of inflation, ipi, gdp, shadow, ffr, oil, m2, vix, and bonds.
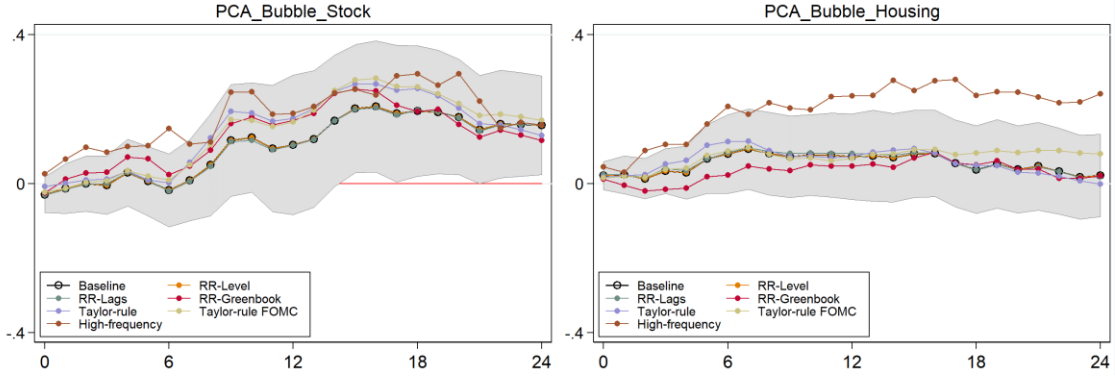
Figure A. Robustness tests for the linear specification
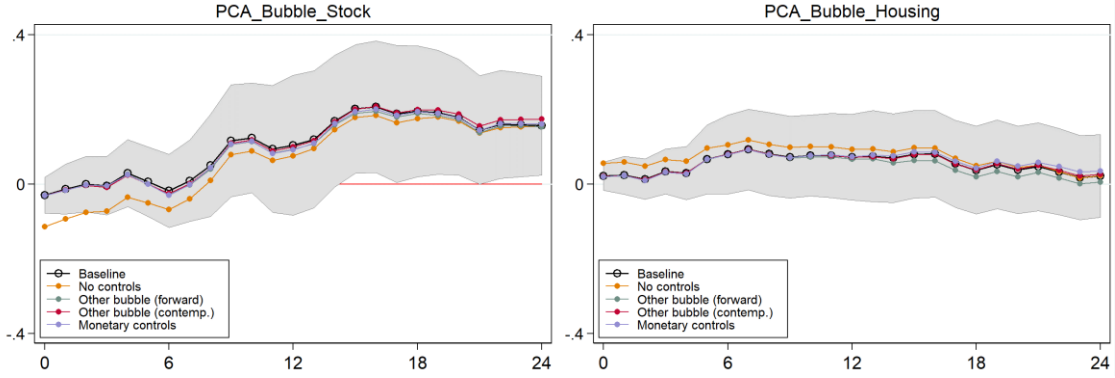
Bubble models decomposition
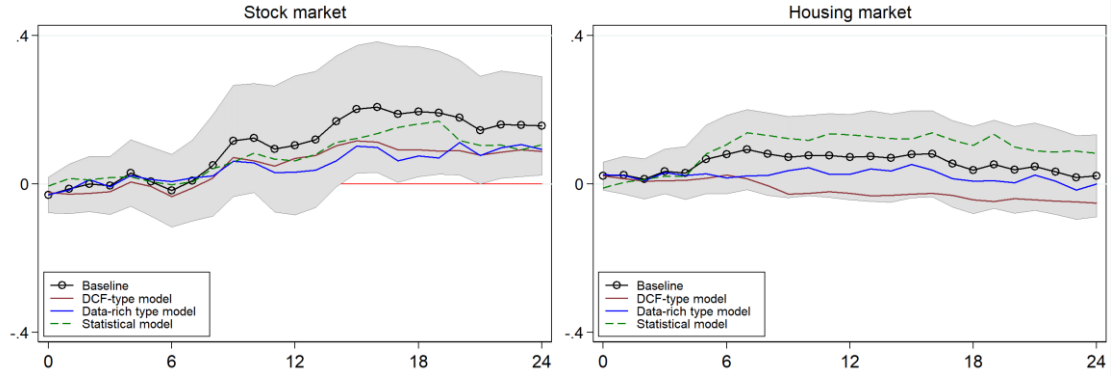
Alternative PCA
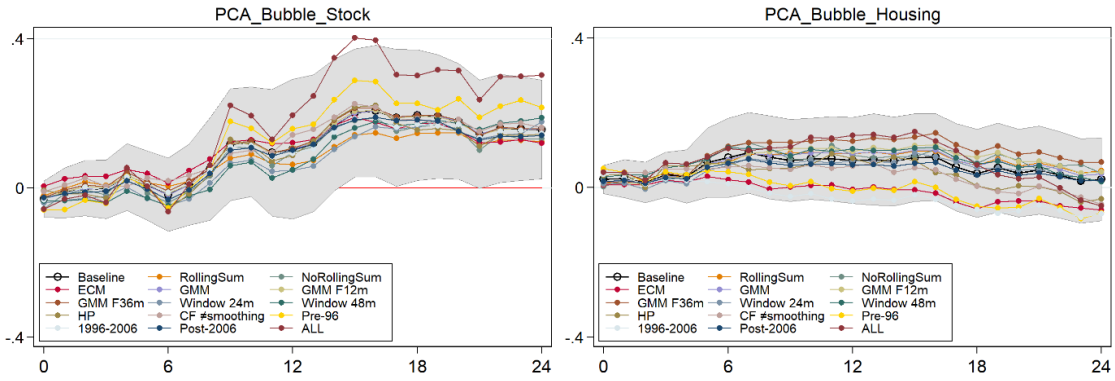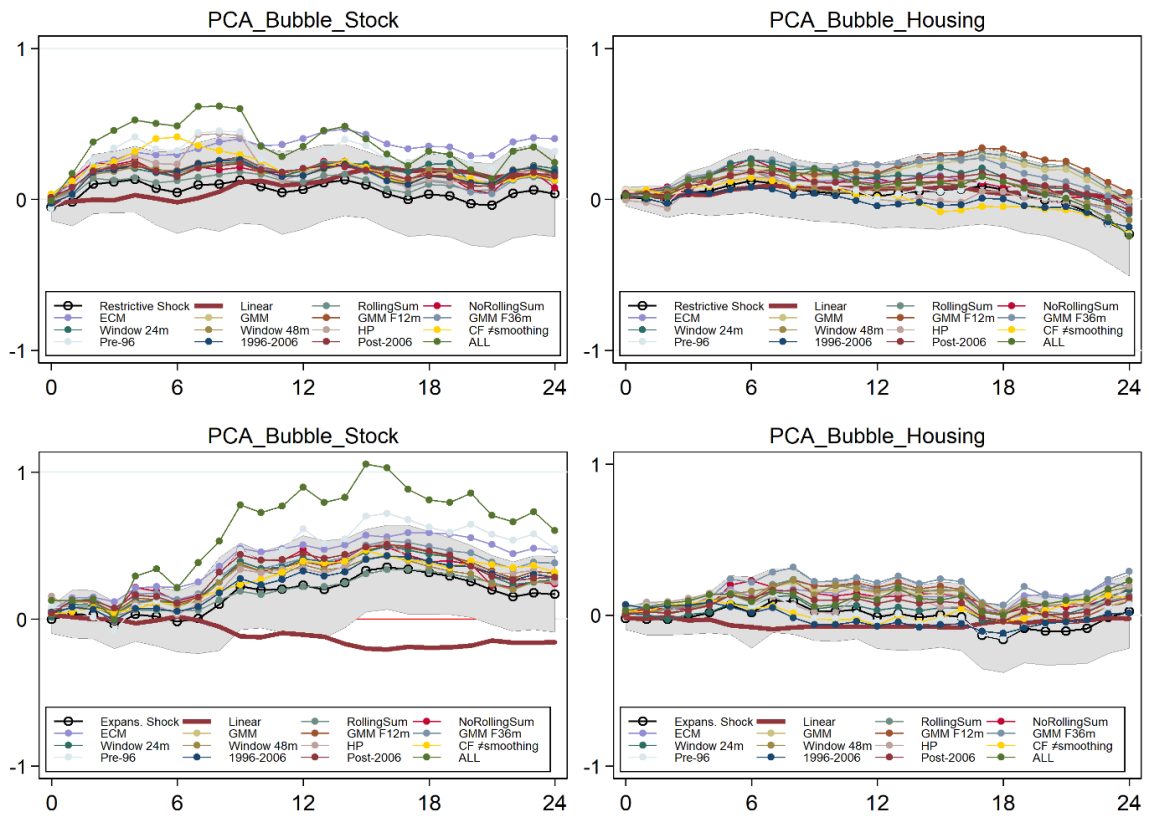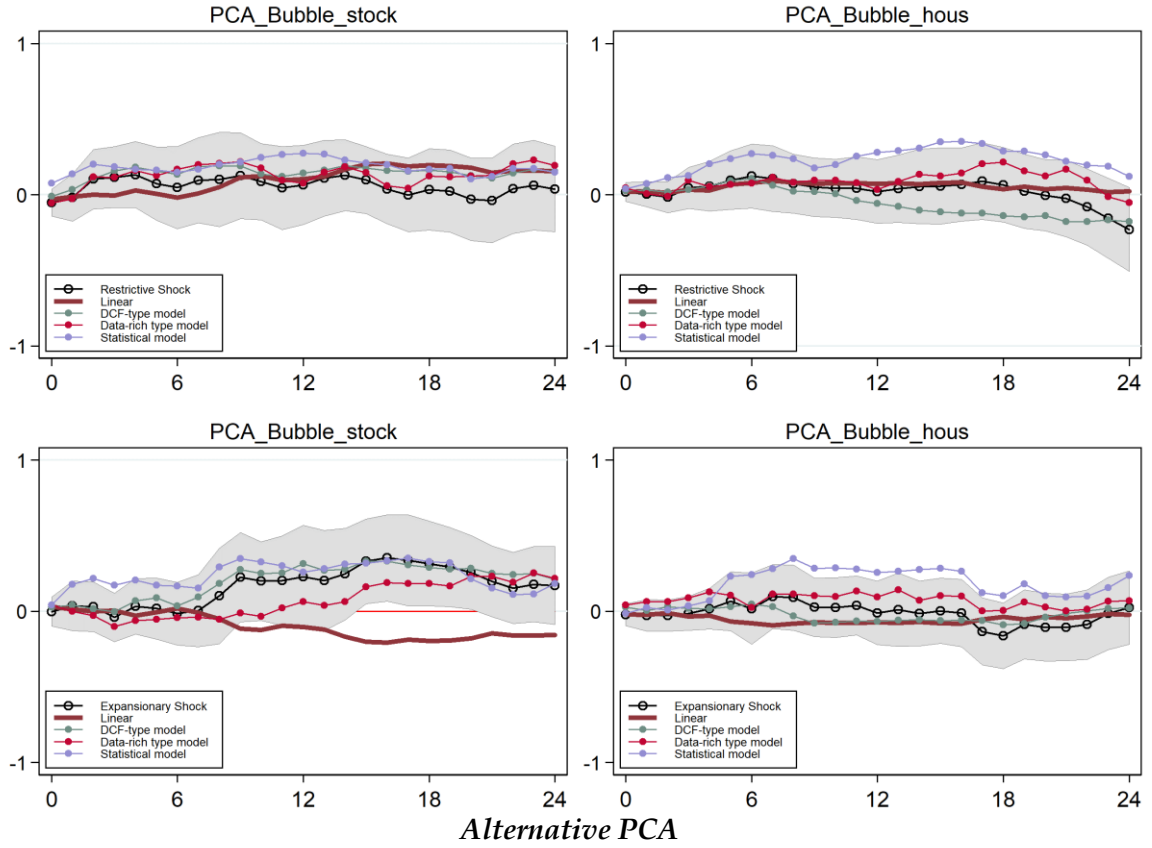

Alternative monetary shocks

LP specification

Note: Linear responses correspond to estimates of equation (11). Shaded area represents the 2 standard errors confidence interval of the response of the baseline shock (MP-Shocks-RR). The y axis is expressed in terms of standard deviation of the bubble indicators. The monetary shock corresponds to a 1 standard deviation increase in the shock series.
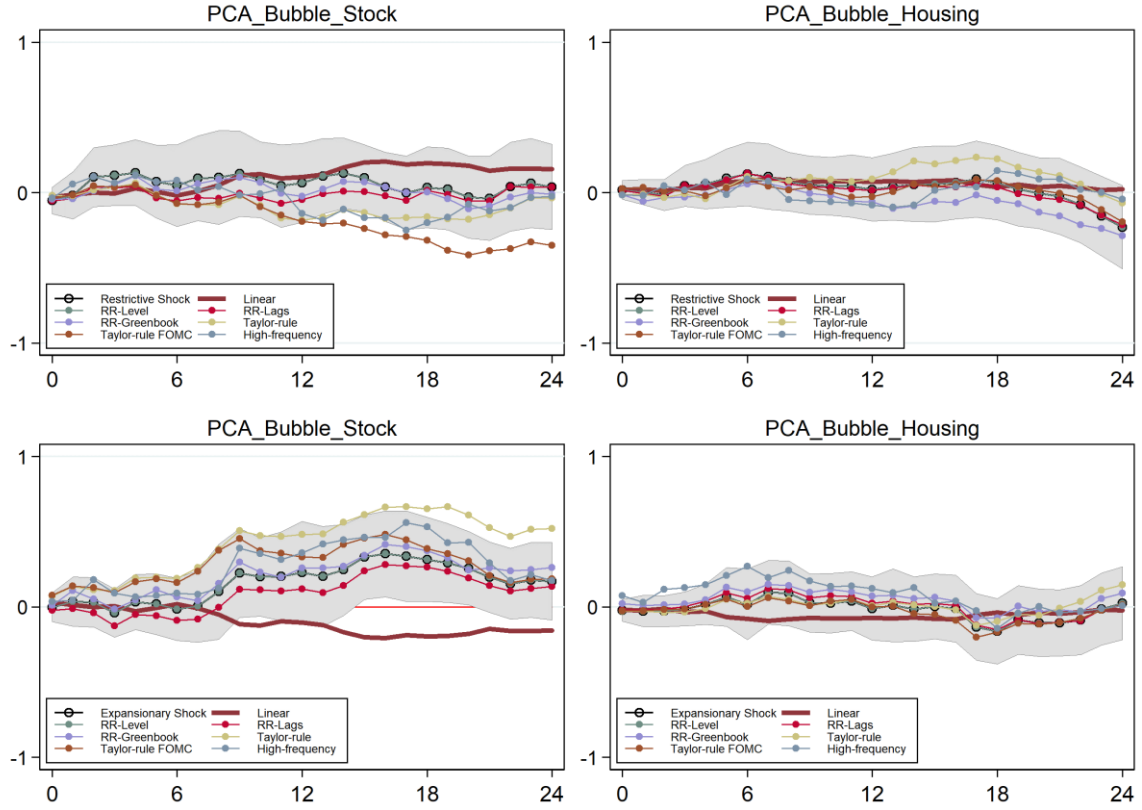
Figure B. Robustness tests to the non-linear specification

Bubble models decomposition

PCA_Bubble_stock

PCA_Bubble_hous

PCA_Bubble_stock

PCA_Bubble_hous

Alternative PCA

PCA_Bubble_Stock

PCA_Bubble_Housing
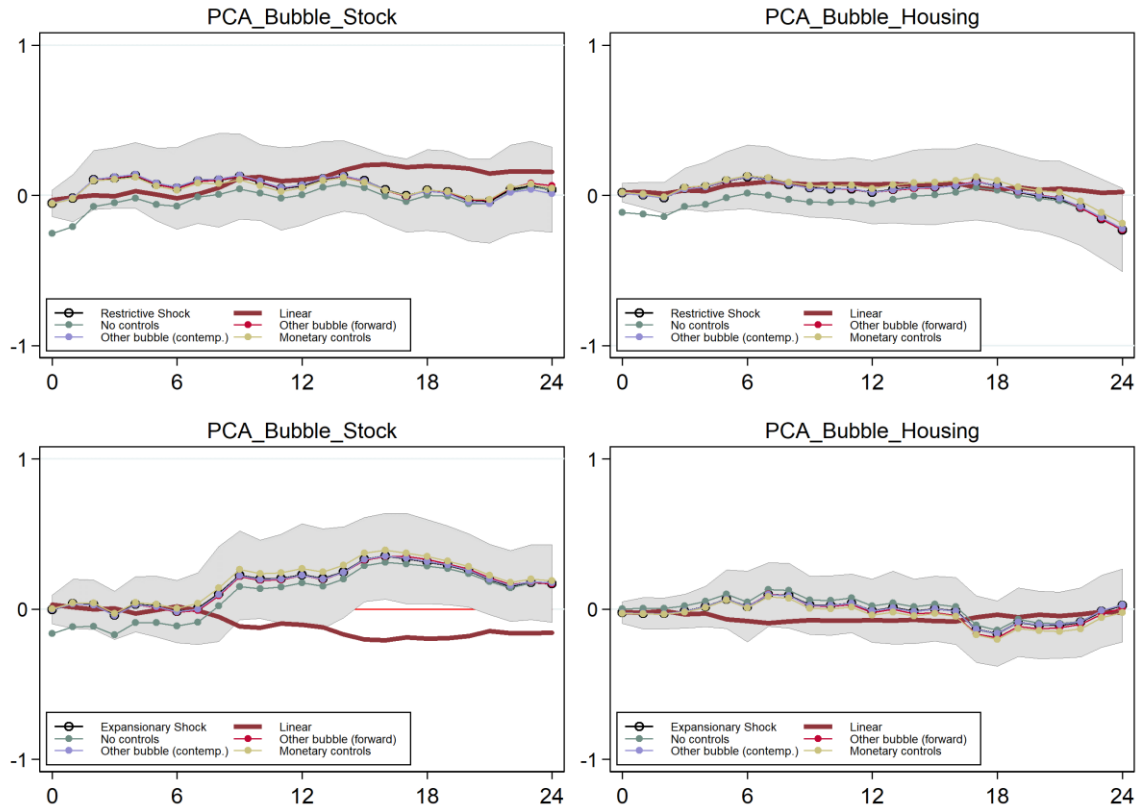

PCA_Bubble_Stock

PCA_Bubble_Housing

Monetary shocks

LP specification

Note: Linear and non-linear responses correspond to estimates of equations (11) and (12) respectively. Shaded area represents the 2 standard errors confidence interval of the response of the baseline shock (MP-Shocks-RR). The y axis is expressed in terms of standard deviation of the bubble indicators. The monetary shock corresponds to a 1 standard deviation increase in the shock series.