THE ASYMMETRIC EFFECTS OF MONETARY POLICY ON STOCK PRICE BUBBLES

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ABSTRACT

Is the effect of US monetary policy on stock price bubbles asymmetric? We use a range of measures of excessive stock price variations that are unrelated to business cycle fluctuations. We find that the effects of monetary policy are asymmetric so responses to restrictive and expansionary shocks must be differentiated. The effects of restrictive monetary policy are more powerful than the effects of expansionary policies. We also find evidence that the asymmetric effect of monetary policy is state-contingent and depends on monetary, credit and business cycles as well as stock price boom-bust dynamics.

KEY WORDS
Non-linearity, Equity, Booms and busts, Federal Reserve.

JEL
E44, G12, E52.

The asymmetric effects of monetary policy on stock price bubbles *

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Abstract

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1. Introduction

Asset price bubbles are generally suspected to represent a threat to financial and macroeconomic stability. While asset prices play a key role in the transmission of monetary policy to the real economy, the very role of monetary policy regarding asset price bubbles remains uncertain and disputed. Policymakers may sometimes contemplate to use restrictive monetary policy to deflate asset price bubbles whereas expansionary policy may have adverse effects on financial stability (Smets, 2014). The debate on whether central banks should react to asset price bubbles hinges on the ability of monetary policy to affect asset price bubbles, which is ultimately an empirical question. This paper contributes to this issue and investigates whether the effects of restrictive and expansionary monetary policies on asset price bubbles are symmetric.

The identification of asset price bubbles is a notoriously major challenge. As illustrated by Brunnermeier (2008) and Martin and Ventura (2018), the theoretical literature has not reached a consensus on the definition and representation of bubbles. Thus, the empirical identification of asset price bubbles involves several approaches (see e.g. Shiller, 2000; Galí and Gambetti, 2015; Jordà et al., 2015). Our research question requires a measure of this unobserved and intrinsically elusive concept of asset price bubbles, and since there is no consensual measure, we choose to rely on various approaches. For the scope of this paper, we think of asset price bubbles as excessive asset price fluctuations that are unrelated to business cycle fluctuations. Based on this definition, asset price bubbles can arise from several approaches. The unifying point is that, whatever their source and nature, bubbles are excessive price movements and raise risks for financial and macroeconomic stability.[1]

While there is a large empirical literature dealing with the impact of monetary policy on asset prices (see e.g. Rigobon and Sack, 2004) or on asset price volatility (see e.g. Bernanke and Gertler, 1999) and with the correlation between monetary and asset price cycles (see e.g. Bianchi et al., 2016), the focus of this paper is on the causal effect of monetary policy on *excessive* stock price variations. On the theory side, Galí (2014) and Allen et al. (2017) reach contrasting theoretical predictions. The former suggests that interest rate hikes increase the bubble-component of stock prices whereas the latter suggest the opposite conclusion. Only a few papers - notably Galí and Gambetti (2015) and Evgenidis and Malliaris (2020) - have dealt with this issue empirically and these works all share the same background assumption that the effect of monetary policy is symmetric. This literature finds mixed results. In this paper we find no evidence of monetary policy influencing stock price bubbles. We test whether this may be the consequence of a potential aggregation bias such that, by accounting for asymmetric effects in the estimation procedure, the average effect can be more precisely estimated. The contribution of this paper is to relax the symmetry assumption and to assess whether the effect of monetary policy on excessive stock price variations is asymmetric. This paper also investigates whether the asymmetric impact of monetary policy is state-dependent, i.e. conditional on monetary, credit, business or stock market cycles.

As asset price bubbles are unobserved, all empirical approaches may fail to disentangle bubbly episodes from evolutions of fundamentals. This raises a concern about a potential misclassification of bubbles. During expansions (recessions), increasing (decreasing) future earnings expectations could be confounded with excessive price variations. To circumvent

[1] The plurality of approaches gives rise to various semantic wordings in the literature from "mispricing" to "boom-busts". We use the term "bubbles" for unwarranted movements from a benchmark value.

this issue, we use several bubble measures derived from different approaches based on different assumptions, and in which potential false positive may be different. In addition, we aim to define the bounds of the potential bias from misclassification by estimating the effect of monetary policy on the fundamentals – earnings – of stock prices.

We gather a broad set of measures of excessive stock price variations from several underlying models to provide robust evidence of the effect of monetary policy. We first consider the Cyclically-Adjusted Price Earnings (CAPE) ratio popularized by Shiller (2000) for the S&P500 which provides an *observable* proxy of over- or under-valuation of stock price markets. An alternative method is to use approaches that enable us to directly measure bubbles as excessive deviations of stock prices from an *estimated* benchmark value.[2] First, we resort to structural models à la Galí (2014) and Galí and Gambetti (2015). We estimate a discounted cash-flow model for the S&P500 to disentangle the fundamental value of stock prices from the bubble component. Second, because the discounted cash-flow model may overlook information stemming from financial and macroeconomic indicators, we compute a stock price bubble indicator as the residual from an equation where the benchmark value is the in-sample fitted value of a regression model accounting for a large information set. Third, we turn to statistical approaches in the spirit of Jordà et al. (2013, 2015). We compute bubbles as deviations from the deterministic trend of stock prices. Finally, for sake of parsimony, we uniquely combine these three measures of excessive price deviations in a composite indicator by estimating their first principal component.

Our empirical analysis proceeds in three steps. First, to measure the causal effect of monetary policy, we need an instrument for exogenous changes in the monetary policy stance. As a baseline, we follow Hanson and Stein (2015) and use the daily change in the 2-year nominal US sovereign yield on FOMC announcement days. Second, we estimate local projections à la Jordà (2005) to take advantage of the flexibility of this approach to analyze the potential asymmetric effects of restrictive and expansionary monetary shocks. Our contribution consists in introducing the possibility for such an asymmetric treatment of the relationship between monetary policy and asset price bubbles. We augment these local projections with an interaction term to estimate differentiated effects for expansionary and restrictive policies. Third, we investigate whether the asymmetric impact of monetary policy is state-dependent. Using triple interaction terms, we test whether the effect of monetary policy is conditional on monetary, credit, business or stock market cycles.

We estimate the dynamic impact of monetary policy over a 2-year horizon to account for its potentially slow and delayed effects on stock price imbalances. The OLS estimation is performed at the monthly frequency over the period from January 1986 to March 2019. Whereas we find no effect of monetary policy on stock price bubbles in a linear setting, we find that the effect of expansionary monetary policy is positive, whereas the effect of restrictive monetary policy is negative. However, while the effect of restrictive policies on stock price imbalances is always significant, there is less evidence of a significant effect of expansionary policies. The key finding of this paper is that the effects of monetary policy are asymmetric: the effects of restrictive monetary policy are more powerful than the effects of expansionary policies. The impact on the CAPE is almost twice bigger in absolute terms for restrictive policies. This suggests that the responses to restrictive and expansionary shocks

[2] We depart from Phillips et al. (2011), Phillips and Yu (2011), Homm and Breitung (2012) and Phillips et al. (2015) who propose recursive unit root tests to detect bubbles in real-time. These works focus on the explosive behaviour of bubbles to measure their starting date, but not their magnitude. Thus, they cannot be used to assess the effect of monetary policy, which may be expected to affect their magnitude rather than their starting date.

must be differentiated. We find that the average effect in the asymmetric setting is different from the linear effect. This finding provides supporting evidence of an aggregation bias supporting the need to resort to an asymmetric setting. Finally, we show that the effect of monetary policy on fundamentals displays no asymmetry so that the asymmetry pattern is specific to stock price bubbles.

The main result of this paper is robust to using the different measures of excessive stock price variations and the composite indicator, to different stock price indices (Dow Jones and NASDAQ), to alternative instrument for exogenous variation in the monetary policy stance (intraday surprises from Miranda-Agrippino and Ricco, 2020, for instance) and to different specifications of local projections. It is also robust to a measure of the CAPE that takes into account share buybacks and to a measure of analysts' expectations of future earnings that capture the forward-looking nature of stock prices. Finally, the frequency and magnitude of restrictive shocks is not different from those of expansionary ones, suggesting that the distribution of shocks does not drive the main result.

In addition, we provide evidence that the asymmetry is strong during tightening monetary cycles but disappears in loosening cycles. We show that the asymmetry is more pronounced and more precisely estimated during expansions than during slowdowns, consistent with the previous result if we assume that monetary policy is countercyclical. Though easing less or tightening more are both restrictive policies, these results suggest that they have not the same effect on asset price bubbles. We also find that the asymmetry is stronger during loosening credit condition cycles. This suggests that policymakers may have more traction on stock price bubbles when facing leveraged bubbles. Finally, we show that the asymmetry is stronger during booms than during busts. These findings enable us to characterize the asymmetry of monetary policy effects more precisely.

Because financial imbalances may generate a misallocation of capital, entail risks to financial stability, trigger deeper and longer recessions and impair the monetary policy transmission, the question of whether monetary policy is able to affect asset price bubbles crucially matters for policymakers.[3] The result that central banks have the ability to fuel and deflate stock price bubbles, although with different strengths, may contribute to shed some light on the debate about the behavior of central banks towards financial imbalances. According to the "leaning against the wind" approach, central banks should react to asset price imbalances (see e.g. Borio and Lowe, 2002, Cecchetti et al., 2003 and Woodford, 2012).[4] The benefits of the latter policy have been balanced against the costs of ex ante restrictive policies (see Bernanke and Gertler, 2001, or Svensson, 2016).[5] The state-dependence effect emphasized here has important implications for policymakers who consider tightening monetary policy to counter financial imbalances.[6] They may achieve similar outcomes by tightening monetary policy relatively less during tightening cycles or macroeconomic expansions. Correspondingly, the policy reaction needs to be stronger during monetary easing cycles or slowdowns.

[3] It is worth stressing 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).
[4] The latter also claim that price stability is not a sufficient condition to promote financial stability. Blot et al. (2015) confirm that there is no stable empirical link between price and financial stability.
[5] Gerlach (2010), Svensson (2012) and Collard et al. (2017) favor the use of macroprudential tools to deal with financial stability.
[6] The leaning against the wind strategy is asymmetric on that respect as it implies that central banks react systematically to positive bubbles. However, empirical evidence has emphasized a "Fed put" (Cieslak et al., 2019, and Cieslak and Vissing-Jorgensen, 2020) according to which central banks would react to asset prices but only when they plummet.

Our empirical investigation does not shed light on the reasons for this asymmetry, so we can only contemplate some speculative explanations. Non-linear effects of monetary policy on asset prices could be inherent to any standard model with rational agents, notably when they are not risk-neutral. Indeed, when investors are risk-averse (risk-lover), their utility functions are not linear but concave (convex). From a different perspective, Gennaioli and Schleifer (2010) analyze agents' beliefs formation and how limited recall of information retrieved from memory gives rise to the representativeness heuristic that induces agents to overestimate the probability of outcomes that are consistent with recent data. During a stock price boom, investors put too much weight on good news and a restrictive monetary decision could trigger the radical change in beliefs highlighted by Gennaioli et al. (2015). The asymmetric reaction of asset price bubbles could also be driven by the negativity bias - more weight is given to bad news compared to good news even though both news are of equal intensity. In the context of bubble processes, expansionary monetary policies are good news while restrictive policies are bad news. Finally, it seems reasonable to argue that central banks do not aim to exacerbate excessive stock price variations, but are rather indifferent or averse to excessive stock price variations. Under such an assumption, expansionary policies are unlikely to signal that policymakers are favorable to more stock price imbalances whereas restrictive monetary policies could be seen as policymakers being more averse to stock price imbalances. Restrictive monetary shocks may therefore be interpreted as a signal of policymakers' intentions to tame asset price bubbles.

The paper relates to the large literature exploring the relationships between monetary policy and asset prices, focusing on the causal effect of the former on the latter (see e.g. Swanson, 2015, Alessi and Kerssenfischer, 2019, Paul, 2019) or on the link between both cycles (see e.g. Detken and Smets, 2004, Bordo and Wheelock, 2007, Ahrend et al., 2008). The closest papers to ours are Basile and Joyce (2001), Galí and Gambetti (2015), Beckers and Bernoth (2016), Filardo et al. (2019) and Evgenidis and Malliaris (2020) that analyze the effect of monetary policy on excessive asset price variations. We differ from these contributions by highlighting the asymmetric effects with respect to restrictive and expansionary shocks. By doing so, we relate to the literature emphasizing asymmetries in the transmission of monetary policy to the real economy. Cover (1992), Angrist et al. (2018) and Tenreyro and Thwaites (2016) notably find that restrictive monetary policy has more powerful effects on macroeconomic variables than expansionary policy, while Garcia and Schaller (2002) and Tenreyro and Thwaites (2016) find that the effects of monetary policy are less powerful in recessions.

This paper is organized as follows. Section 2 assesses the asymmetric effects of monetary policy. Section 3 investigates the state-dependence of this effect. Section 4 examines the robustness of our main finding. Section 5 concludes.

2. The asymmetric effects of monetary policy

Do restrictive and expansionary monetary policies affect stock price imbalances differently? Answering this question raises several issues related to the identification of excessive stock price variations, the measure of monetary policy stance and the econometric model which is best suited to quantify these potential asymmetric effects. Our strategy builds on recent contributions in the macroeconomic literature and combines local projections of Jordà (2005) with market-based instruments for monetary policy shocks by Hanson and Stein (2015). The empirical strategy is first applied to the stock market excess valuation metric popularized by Shiller (2000). In a second step, we estimate stock price bubbles as excessive deviations of stock prices from an estimated benchmark value.

2.1. A muted linear effect

According to Campbell and Shiller (1998, 2005), the Cyclically-Adjusted Price Earnings (CAPE) ratio – defined as the ratio of the inflation-adjusted S&P500 price over average earnings from the previous 10 years – is a good indicator of future price adjustments.[7] An increase in the ratio is indeed generally followed by a decrease in stock prices rather than an increase in dividends, which would be the case under the efficient market hypothesis.[8] It may consequently be interpreted as an indicator of mispricing relative to earnings. Stock price data are monthly averages of daily closing prices and are deflated by the CPI-U (Consumer Price Index-All Urban Consumers).[9] This measure is widely used both in the academic literature and by practitioners in financial markets because of its simplicity and transparency.[10] Figure A in the Appendix plots the time series of CAPE.

We assess the dynamic impact of monetary policy on stock price valuations using Jordà (2005)'s Local Projection method. In linear stationary settings, the forecasting performance of VARs and local projections is quite similar (see Marcellino et al., 2006, and Kilian and Kim, 2011). However, because a linear autoregressive representation of the data generating process of time series used in this paper may be misleading, the robustness of local projections to model misspecification makes them an appealing procedure to recover dynamic responses. The relevance of local projections, which can easily account for non-linearities, is even more pronounced since the research question focuses on investigating potential asymmetries. Considering that some 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:

$$y_{t+h} = \alpha_h + \beta_h \epsilon_t + \phi_h . X_t + \eta_{t+h} \tag{1}$$

where y_{t+h} is the stock price excess valuation measure of Shiller (2000) at the horizon h, ϵ_t is the exogenous instrument used to measure the causal effect of monetary policy, and X_t is a vector of control variables that includes a lag of the dependent variable and a lag of the exogenous instrument for the monetary policy stance.[11]

The estimation of causal effects of monetary policy with local projections requires to use externally identified exogenous innovations as instruments for monetary policy shocks. Following Hanson and Stein (2015), we use the daily change in the 2-year nominal US sovereign yield on FOMC announcement days to quantify surprises in the FOMC monetary stance.[12] This measure has the advantage of being observed compared to monetary shocks estimated from a first step that would be generated regressors and may bias standard errors. As is common in the macro literature, we sum daily monetary surprises to the monthly

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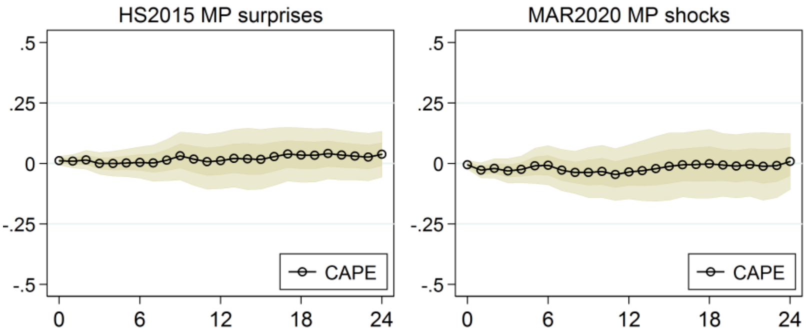
[7] Shiller (http://www.econ.yale.edu/~shiller/data.htm) explains how annual data is transformed: "Monthly dividend and earnings data are computed from the S&P four-quarter totals for the quarter (…), with linear interpolation to monthly figures".
[8] Stock prices increase when markets participants expect higher dividends.
[9] See Table A in the Appendix for data description and sources.
[10] For robustness purposes, we also use the total return CAPE (TR-CAPE) which takes into account corporate payout policies and share buybacks, and a measure of expected earnings 12 months ahead by brokers and analysts which takes into account the forward-looking nature of stock prices.
[11] Including the VIX index as an additional control to proxy for the time-varying risk-premium does not change the results.
[12] Standard high-frequency methodologies, such as Kuttner (2001), use changes in the price of futures contracts to identify monetary surprises. However, unconventional measures may lead to underestimate the true monetary policy surprise, which are not reflected in the short-term interest rate. In section 4, we show estimates with Kuttner (2001) surprises on a smaller subsample, and Miranda-Agrippino and Ricco (2020) monetary shocks for robustness purposes. The main result of Figure 1 holds in all cases.

frequency. As conventional and unconventional decisions are announced on FOMC statement days, our measure captures surprises related to all monetary policy measures. In addition, Gürkaynak et al. (2005) show that a large part of the news embedded in FOMC announcements is about the monetary policy stance over the next quarters. Our measure therefore captures these signals about the likely future policy path (i.e. formal or informal forward guidance). Figure B in the Appendix plots the time series of monetary shocks.

When private agents' and central bank information sets are different, monetary policy surprises are not akin to monetary policy shocks. The former are shocks to private agents' information sets whereas the latter are shocks to the policy instrument. Changes in interest rate around policy announcements may not only signal the policy stance but also provide information on policymakers' view of the future state of the economy such that monetary surprises would then be contaminated by 'signaling' or 'information effects' (Melosi, 2017, and Nakamura and Steinsson, 2018). Jarocinski and Karadi (2019) and Miranda-Agrippino and Ricco (2020) provide evidence of these information effects, and construct monetary policy shocks corrected for this information transfer related to policymakers' views of the future state of the economy.[13] We use Miranda-Agrippino and Ricco (2020)'s monetary shocks to assess whether the response of stock price imbalances to monetary surprises is driven by the 'information effect' or by pure monetary shocks.

Figure 1 – The linear effect of monetary policy



Note: Estimates are based on equation (1). Both estimations are run over the sample January 1986 - March 2019. The dependent variable is Shiller's (2015) CAPE ratio. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015) on the left panel and are replaced by Miranda-Agrippino and Ricco (2020)'s shocks on the right panel. Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

We estimate equation (1) with OLS from January 1986 to March 2019, with heteroskedasticity robust standard errors.[14] We set h equal to 24 periods such that Figure 1 plots the dynamic responses of Shiller (2000)'s CAPE measure to monetary policy over 2 years. We find that monetary surprises have no effect on the CAPE ratio at all horizons over our sample. This is also true for information-corrected monetary shocks. Whereas since Bernanke and Kuttner (2005) the literature has provided ample evidence that monetary policy affects negatively stock prices, recent work on stock price imbalances suggest mixed results. Galí (2014) makes the point theoretically that "rational bubbles" respond to interest rate rises by growing faster. Galí and Gambetti (2015) provide empirical evidence supporting this view, but Beckers and Bernoth (2016) find opposite results. A potential reason for the null effects displayed in

[13] See Cieslak and Schrimpf (2019) and Lakdawala and Schaffer (2019) for alternative ways of controlling for information shocks.
[14] Estimations start in January 1986 to be consistent with regressions for bubble indicators (see below).

Figure 1 may be that the average effect would be more precisely estimated in a non-linear framework. Our hypothesis is that there might be an aggregation bias at work because bubble processes are not constant over time, such that the effects of monetary policy on stock price imbalances are not linear. We test this hypothesis in the next section by allowing for asymmetric effects.

2.2. Allowing for asymmetric effects

In order to investigate if expansionary and restrictive monetary surprises have different effects, Equation (1) is rewritten to account for non-linearities. Such asymmetric impacts of monetary policy have been considered for the transmission to the real economy (see Weise, 1999, Lo and Piger, 2005, Tenreyro and Thwaites, 2016, Angrist et al., 2018). The literature on the effect of monetary policy on stock price imbalances has so far made the assumption that the impact is linear and we aim to relax this assumption and explore whether a linear framework has missed a key feature of the dynamics of imbalances and biased the outcomes. Equation (1) is therefore augmented with a latent variable and interaction terms:

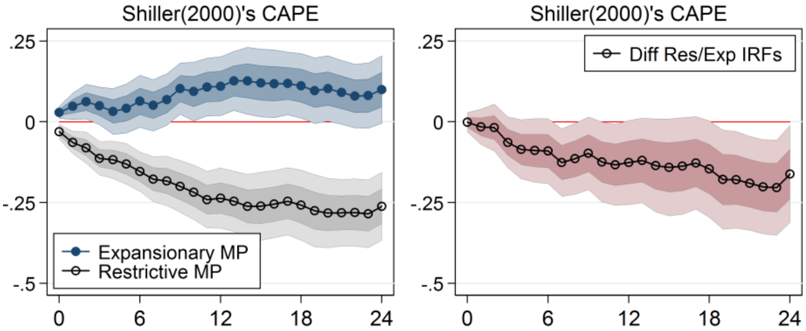
$$y_{t+h} = \alpha_h + \beta_{\Pi,h}(\epsilon_t.\mathrm{I}_t) + \beta_{\epsilon,h}\epsilon_t + \beta_{\mathrm{I},h}\mathrm{I}_t + \phi_{X,h}X_t + \phi_{\Pi,h}(X_t.\mathrm{I}_t) + \eta_{t+h} \tag{2}$$

where $\mathrm{I_t}$ is a dummy variable that equals 1 for expansionary monetary shocks and 0 for restrictive monetary shocks. $\epsilon_t.\mathrm{I}_t$ is the interaction term between monetary shocks and the latent dummy variable. This specification enables us to single out the potential asymmetric effects of restrictive ($\beta_{\epsilon,h}$) and expansionary ($\beta_{\epsilon,h} + \beta_{\Pi,h}$) monetary policy on stock price imbalances. Because the effect of control variables on the dependent variable may also change with the sign of monetary shocks, we include an interaction term $X_t.\mathrm{I}_t$ between the vector of controls and the latent variable. Based on the result presented in Figure 1, the exogenous instrument for monetary policy in the baseline specification is the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015).[15] Equation (2) is estimated with OLS from January 1986 to March 2019, with heteroskedasticity robust standard errors. Table B in the Appendix provides the estimated parameters of this equation. Monetary surprises (β_ϵ) have a negative and significant effect on the CAPE index at all horizons. The interaction term (β_Π) is significant and positive at all horizons too.

Figure 2 plots the dynamic responses of Shiller (2000)'s CAPE measure to monetary policy over 2 years. The effect of expansionary monetary policy on the CAPE ratio is positive, whereas the effect of restrictive monetary policy is negative. However, the effect of restrictive monetary surprises is significantly larger than the effect of expansionary ones. The peak effect of expansionary monetary policy (an exogenous one-S.D. decreases in the policy rate) would increase the CAPE by 0.12 standard deviation (see the blue swathes on the left panel of Figure 2). At the opposite, an exogenous one-S.D. increases in the policy rate would decrease the CAPE by 0.28 standard deviation at maximum (see the grey swathes). To give an idea of the magnitude of the effects, the mean and standard deviation of the CAPE is 24 and 7 respectively over our sample, while a one-S.D. of monetary surprises corresponds to 5 basis points. The model explains 75% of the variance of the CAPE index at 12-month ahead (see Table B in the Appendix). The contribution of monetary surprises accounts for around 10% of the variance, suggesting that monetary policy has sizable effect on mispricing even if it does not account for the bulk of misalignments.

[15] We also report the responses of the CAPE ratio to Miranda-Agrippino and Ricco (2020)'s monetary shocks in the Appendix.

Figure 2 – The asymmetric effect of monetary policy



Note: Estimates are based on equation (2) over the sample January 1986 - March 2019. The dependent variable is Shiller's (2015) CAPE ratio. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

The asymmetric response of the CAPE index is confirmed by the right panel of Figure 2 that shows the difference between the effects of restrictive and expansionary monetary policies. The difference is negative and significant. A key consequence of this finding is that the responses to restrictive and expansionary monetary policy must be differentiated. The main result of this paper is that – when allowing for non-linearities – the effects of US monetary policy on stock price bubbles are found to be asymmetric: restrictive monetary policies have larger effects than expansionary policies.

In addition, given that expansionary and restrictive surprises occur at almost equal frequencies and magnitudes (see Table C in the Appendix), the linear effect should be the average of restrictive and expansionary monetary policies. However, the linear effect displayed in Figure 1 does not correspond to the average of the effects of restrictive and expansionary monetary policies displayed in Figure 2. In addition, the average effect of monetary policy surprises in a non-linear set-up is negative and significant as reported in Table B in the Appendix. This suggests that using a linear framework produces an aggregation bias and reinforces the rationale for using a non-linear framework to explore the question of the effect of monetary policy on stock price imbalances.

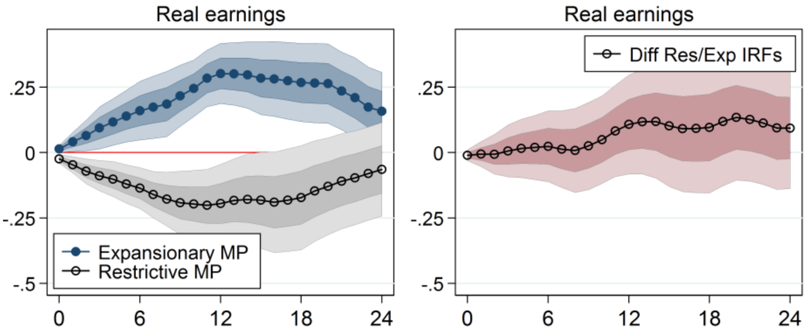
2.3. The risk of misclassification

Since asset price bubbles are unobserved, any empirical approach is exposed to a risk of misclassification that may either provide a signal of bubble when there is none or miss to detect ongoing bubbles. When the economy grows faster (slower), potential misclassification risks intensify as increasing (decreasing) expected earnings and bubbles may be confounded and generate false positives. In the case of estimating an average effect, misclassifying a non-bubble as a bubble would bias the estimated coefficient towards the response of fundamentals. To deal with this issue, we use several alternative measures of price deviations from fundamentals that each yields different potential misclassifications. If our main result is specifically driven by a misclassification resulting from the CAPE, it may not appear with alternative measures. We explore this avenue in the next section.

In our asymmetric setting where tightening monetary policy is more powerful than easing, we might worry that misclassification would not be equally shared between periods of tightening and easing and would happen more often during monetary tightening. If so, the coefficient on monetary tightening would be more biased toward the response of fundamentals than the coefficient on monetary easing. The asymmetry would purely arise from the misclassifications happening when monetary tightening occurs.

The CAPE could be prone to misclassification because it is a backward-looking measure.[16] By construction, it increases when current prices rise faster than the 10-year average of earnings. In monetary tightening cycles, the economy is exhibiting strength and the CAPE is expected to be high because prices reflect expectations about the future economy's strength. Yet the CAPE measure, based on past earnings, would classify this rise as a bubble. However, if these expectations are grounded, the CAPE rises might generate false positives and the estimated effect of tightening monetary policy might be capturing decreases in the fundamental component of stock prices.

Figure 3 – The asymmetric effect on fundamentals



Note: Estimates are based on equation (2) over the sample January 1986 - March 2019. The dependent variable is real earnings of S&P500 firms. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

Estimating the responses of the "fundamental" component to monetary policy helps bound any potential bias of bubble responses to tightening. If the response of the fundamentals to monetary policy is more negative than the response of the bubble component, then the misclassification might be an issue. However, if the fundamentals response is less negative than the effects of restrictive surprises on the bubble component, the bias goes the other way and the asymmetry evidenced earlier could even be understated.

We estimate equation (2) with real earnings of S&P500 (not the 10-year average) as the dependent variable.[17] Data are retrieved from Shiller's website. Compared to the CAPE responses, Figure 3 shows that the asymmetry is not significant. The effect of expansionary policy on earnings is positive while restrictive monetary policy reduces earnings, both with similar magnitudes. Consequently, the asymmetry is likely to be related to the bubble component rather than to fundamentals or simply asymmetric stock price behaviour.

2.4. Exploring the scope of asset price bubble representations

There is no consensus on the most appropriate way to identify bubbles empirically or a unique definition of them. According to Brunnermeier (2008): "*Bubbles are typically associated with dramatic asset price increases followed by a collapse. Bubbles arise if the price exceeds the asset's*

[16] Another way to tackle the misclassification issue is to use an alternative price-earnings ratio based on current (not the 10-year average) earning expectations 12-month forward that is not prone to this criticism. See Table H in the Appendix.
[17] Table C in the Appendix shows the effect of monetary policy on the fundamental-component of stock prices, computed as the fitted value of the S&P500 index based on real earnings.

fundamental value" whereas Shiller (2014) emphasizes the role of psychology and investors' emotions.[18] These definitions reflect theoretical controversies as bubbles arise in many distinct models and the intrinsic difficulty for representing the benchmark value for asset prices.[19] Bubbles are defined as a deviation from an equilibrium value which is related to some expected variables (dividends for instance in the discounted cash-flow model for stock prices). Yet, as emphasized by Gürkaynak (2008), testing for bubbles entails a joint-hypothesis and the rejection of the null hypothesis may indicate either a rejection of the bubble hypothesis or a rejection of the model used to infer the benchmark value.

The literature on asset price bubbles has provided several potential measures of imbalances. In the rest of the paper, we aim to show that our result holds for many representations of stock price bubbles and are not specific to the CAPE. Galí and Gambetti (2015) resort for instance to a "structural model" where the bubble component of stock prices is estimated from a VAR model including notably dividends, the nominal interest rate and the GDP deflator. The empirical estimation enables us to identify the response of fundamental value from the discounted cash-flow model and the bubble as the difference between stock prices and the estimated fundamental value.[20] Conversely, Jordà, Schularick and Taylor (2012, 2015) do not identify the fundamental value from a structural model. They rely on the dynamic properties of bubbles characterized by excessive increases of asset prices followed a sharp fall. The bubble is identified when there is a significant deviation – beyond one standard deviation – of the asset price from a statistical trend followed by a fall of the price exceeding 15%. The estimated trend captures the benchmark value.

In this paper, our approach is practical and agnostic. We acknowledge that these approaches provide different insights on stock price bubbles. Considering that stock prices (P_t) are the sum of a benchmark value (F_t) and a bubble component (B_t), we resort to three alternative methods to estimate (F_t) and infer the bubble component of stock prices.

First, under a "structural" approach, the bubble component is captured as the deviation from the fundamental value derived from the estimation of the discounted cash-flow model (see Fama, 2014). Second, we use a larger set of information to estimate the benchmark value, which may provide a better fitted value in-sample. This approach is called "data-driven". Third, we resort to a "statistical" approach to identify deviations from a statistical trend.

In a Discounted Cash-Flow (DCF) model, the fundamental value of stock prices, under full information and rational expectations, is the sum of expected discounted future cash-flows:

$$F_t = \sum_{i=1}^{\infty} \left(\frac{1}{1+\rho}\right)^i E_t(D_{t+i}) \tag{3}$$

[18] See Shiller (2014), page 1487: *"news of price increases spurs investor enthusiasm which spreads by psychological contagion from person to person, in the process amplifying stories that might justify the price increases and bringing in a larger and larger class of investors, who, despite doubts about the real value of an investment, are drawn to it partly through envy of others' successes and partly through a gambler's excitement."*

[19] See Brunnermeier and Oehmke (2013) and Scherbina (2013) for surveys. The literature can be broadly separated in two strands. Blanchard and Watson (1982), Diba and Grossman (1988), Galí (2014), Allen and Gale (2000, 2004) Gruen et al. (2005), Christiano et al. (2010), Allen, Barlevy and Gale (2017) start from a rational bubble framework, whereas Abreu and Brunnermeier (2003), Ofek and Richardson (2003), Kindleberger (2005) and Schiller (2015) and Bordalo et al. (2017) give some role to behavioural dynamics.

[20] The response of the fundamental component of stock prices is inferred from the impulse reaction functions of dividends, the nominal interest rates and the GDP deflator.

where D_t stands for the cash-flow (here dividends for stock prices) and ρ is the discount factor proxied by the long-term interest rate. Assuming risk-neutral agents, a constant discount factor and constant cash-flows, equation (3) becomes:

$$F_t = D/\rho \tag{4}$$

Equation (4) can be used to identify deviations as the difference between the current price and its estimated fundamental 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 (5) relating the asset price to the current cash-flow, the discount factor and the risk-premium with OLS:

$$P_t = \alpha_0 + \alpha_1 . D_t + \alpha_2 . \rho_t + \alpha_3 . \phi_t + \kappa_t \tag{5}$$

where P_t is the log of the stock price index – the S&P 500 – deflated by the consumer price index. D_t is the associated cash flow of this given asset (real smoothed dividend), ρ_t is the time-varying discount factor captured by the real long-term sovereign interest rates, and ϕ_t a proxy for the time-varying risk-premium, measured by the VIX – the Chicago board of trade volatility index – which is often used as a proxy for uncertainty and market appetite for risk. κ_t is the residual and stands for the deviation of the stock price from fundamentals. Data on dividends are available at a quarterly frequency.[21] The fitted value of equation (5) is an unconstrained estimation of the DCF model and has the same interpretation as the CAPE indicator. A rise in the current stock price relative to its fitted value – $\widehat{P_t} = \widehat{\alpha_0} + \widehat{\alpha_1} . D_t + \widehat{\alpha_2} . \rho_t + \widehat{\alpha_3} . \phi_t$ – signals a deviation – here a positive κ_t – of the stock price relative to the fundamental value. It differs from the CAPE indicator since we use the current dividends instead of the ten-year average of real earnings, $\widehat{\alpha_0}$ and $\widehat{\alpha_1}$ are not constrained respectively to zero and one, and we also account for changes in the discount and the risk-premia factors.[22]

An alternative approach to proxy the benchmark value is to account for all available information available at time (t). To that end, we estimate a model where the stock price index is explained by a large set of macroeconomic and financial variables. Thus, the OLS estimate provides the best in-sample prediction of the stock price conditional to a given information set. The bubble indicator stems from a "data-driven" approach and is the unexplained component of equation (6):

$$P_t = \beta_0 + \beta(L) . P_t + \beta_1 . M_t + \beta_2 . F_t + v_t \tag{6}$$

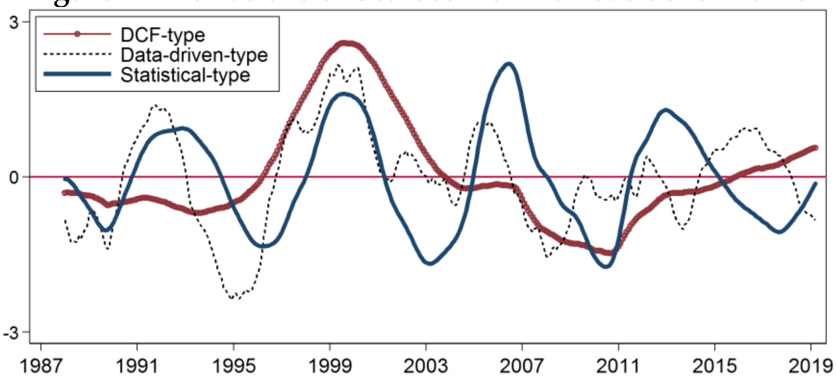
M_t and F_t are vectors of macroeconomic and financial variables covering a large set of information, which may not be captured by current dividends but that would yet provide information on future dividends and hence contribute to influence the benchmark value for stock prices. Variables included in equation (6) are dividends, real interest rate and VIX but also real GDP, inflation, M2 (deflated by the CPI), the outstanding amount of credits (deflated by the CPI), the Conference Board consumer and ISM firm confidence indicators, the house price (deflated by the CPI), the oil price (deflated by the CPI) and real disposable income. For all explanatory variables, 3 lags are included in the specification.[23] Lags of the endogenous variable are also included in the estimation. The residuals capture the

[21] Quarterly data are linearly interpolated to monthly frequency.
[22] We also assess the robustness of our main result with a ten-year moving average of real dividends.
[23] Specifications with leads have also been tested but do not change the result and the residual dynamics.

component of the asset price unrelated to macroeconomic and financial fundamentals. The residuals from equation (5) and mostly from equation (6) may suffer from an endogeneity bias if asset price bubbles boost macroeconomic and financial variables. In that case, a positive bubble would push up the growth of credit and economic activity increasing the correlation with the observed current stock price. The residuals would be smaller - introducing a bias in the measure of the bubble. There is consequently a trade-off between the information which is accounted for - and provides a better evaluation of the benchmark value for stock prices - and the bias introduced to capture the bubble component. The measure provided by the data-driven approach would thus be a lower estimation of bubbles. Yet, as we show after, the response of the stock price bubble to monetary policy surprises is still significant in this worst case scenario.

We also consider a model corresponding to the "statistical" approach where bubbles are measured by the 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. Bordo and Jeanne (2002), Detken and Smets (2004), Goodhart and Hofmann (2008), Alessi and Detken (2011), Bordo and Landon-Lane (2013), and Jordà et al. (2013, 2015) use deviation from a trend to identify bubbles. Our third model relies on the estimation of deviations from the one-sided Christiano-Fitzgerald (CF) trend/cycle decomposition.[24]

Figure 4 – Deviations of S&P500 from various benchmarks



Note: The DCF-type measure is estimated with equation (5), the data-driven measure is estimated with equation (6), and the statistical measure is computed using the Christiano-Fitzgerald filter.

So far, asset price bubbles are defined as the difference between the asset price and its benchmark (the fundamental, the fitted value or a statistical trend). However, these deviations are static and likely to be short-lived. Such static deviations could result from anomalies in financial markets more than bubbly processes that are not seen as one-period events. They are likely to be small (Filardo, 2004) and irrelevant for macroeconomic stability. Each single (one-period) deviation may not matter for policy makers but successive positive or negative deviations may signal a persistent imbalance. We therefore compute a moving average of these deviations - over 49 months, from 24 months before to 24 months after - to capture dynamic and persistent deviations from the respective benchmarks over a medium-term sample.[25]

[24] 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. The CF filter is designed to exclude cycle component below 12 months.
[25] As a robustness check, we show that the results hold without smoothing the residuals.

We estimate these three models for the US stock prices. Due to the availability of the VIX indicator, the sample period spans from January 1986 to March 2019. The stock price index is the S&P500, deflated by the CPI.[26] Figure 4 shows the time-series of these smoothed deviations. The dotcom bubble is clearly identified by all models whereas a bubble is identified in 2007 but only with the statistical model. It has to be stressed that all indicators suggest an under-valuation of the stock prices in 2008-2009. There is a rebound in 2012-2013 identified with the statistical model but not with other models. Regarding the 2017-2019 period, only the DCF-type model points a small bubble relative to previous peaks. It is worth signalling that the DCF model provides a picture close to the CAPE indicator. Both indicators suggest a major boom in 2000, smaller ones in 1987 and 2017-2018 and an under-valuation of the US stock prices in 2008-2009. These alternative measures of bubbles show that the classification of an asset price increase (decrease) as a positive (negative) bubble crucially hinges on the identification approach. Regarding the effect of monetary policy on stock price bubbles, these differences are important to ensure that our conclusions are not driven by one specific model.

2.5. Monetary policy and stock price bubble indicators

Equation (2) is estimated with the three measures of stock price imbalances. All models confirm that the effect of monetary policy is asymmetric (Figure 5). While expansionary shocks are followed by a non-significant increase in the size of the bubble, the response to restrictive monetary policy is negative and significant. Those results confirm that monetary policy would be more powerful in reducing the size of bubbles than in fueling them. It is interesting to note that the asymmetry is less pronounced when we consider the data-driven approach although the magnitudes are not statistically different between the three models. This suggests that a potential misclassification bias is not responsible for this result.

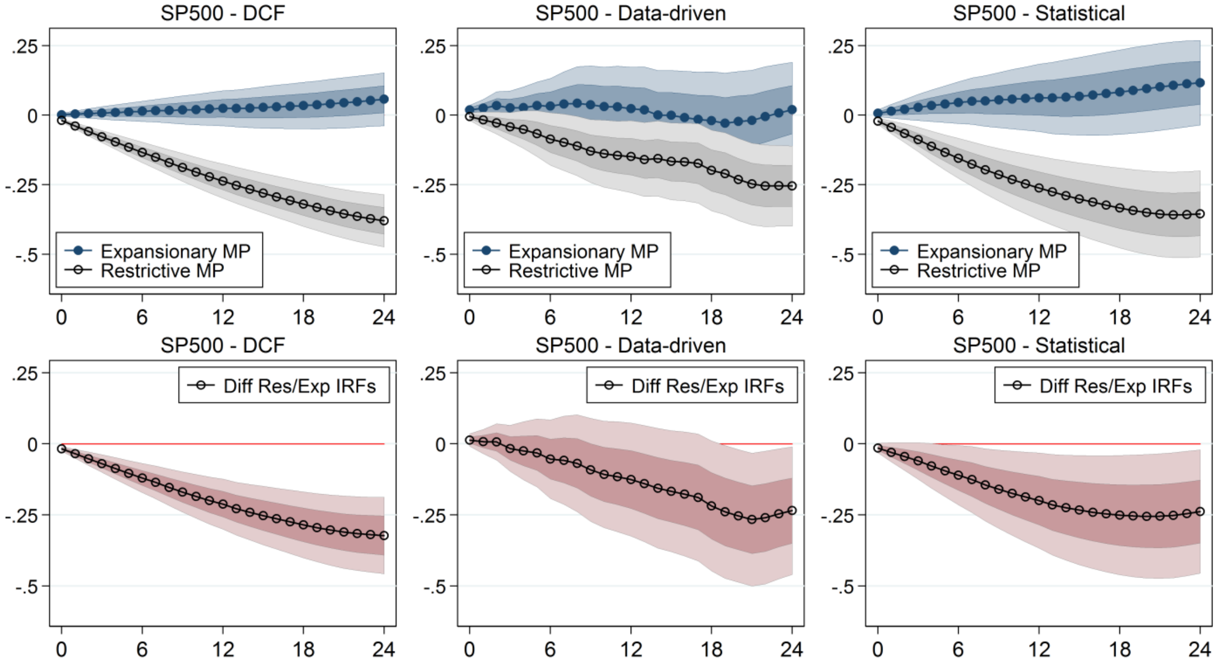
The DCF and statistical approaches yield representations of the benchmark value centered either on the fundamental of the asset (DCF) or on the data generating process of the asset price (statistical) such that the bubble may account for the effect of various remaining factors. Conversely and by construction, the data-rich approach encompasses additional factors in the estimation of the benchmark value (the fitted value) such that the bubble is not explained by those macroeconomic and financial variables but would potentially result from investors' heuristics such as herding, myopia, etc. Shutting down the role of macroeconomic and financial variables in the determination of stock price bubbles reduces the asymmetry of monetary policy. It suggests that the asymmetric effect of monetary policy may be partially related to the influence of these variables, but is still nonetheless significant. We investigate this topic in the next section.

For the sake of simplicity and parsimony, we summarize the information provided by the three models in a composite indicator. To that end, we compute the first component from a principal-component analysis (PCA). Estimating the first principal component enables us to extract the common denominator of the three models used above. This first principal component has an eigenvalue of 1.79 and explains 60% of the variance of the three stock price imbalance series. The composite index for the bubble component of the S&P500 is shown in Figure D in the Appendix.[27] Three peaks – 1992, 2000, 2007 – and four troughs –

[26] The same exercise is performed with the Dow Jones and the NASDAQ indices (see robustness).
[27] Figure D also plots the time series of the composite index for the Dow Jones and the NASDAQ.

1990, 1996, 2004 and 2010 – emerge since 1986. The PCA measure confirms that there is weak evidence of a bubble for the S&P 500 over the recent period (at least until March 2019).

Figure 5 – Impulse response functions *(upper panel)* and difference between expansionary and restrictive responses *(lower panel)*

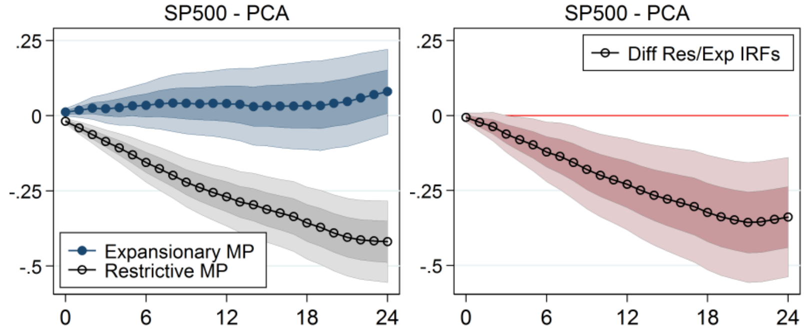


Note: Estimates based on equation (2) over the sample January 1986 - March 2019. The dependent variables are the three proxies for stock price imbalances. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

Figure 6 plots the dynamic responses of the stock price bubble composite measure to expansionary and restrictive monetary surprises estimated with equation (2). The asymmetric effect of monetary policy on stock price bubbles is confirmed and only the response to restrictive policies is significant. The parameter estimates are displayed in Table B in the Appendix. From the current month to the 24th month after the shock, the effects of restrictive monetary decisions are significant and the interactions term is always positive and significant. The contribution of monetary policy surprises to the variance of the bubble indicator varies between 6 and 12% which is sizeable and close to the CAPE estimates. Finally, the repeated evidence of the asymmetric effect of monetary policy suggests that a potential misclassification bias is not responsible for this result.

This asymmetry might be related to different tentative explanations on which the existing empirical literature may shed light. On the one hand, restrictive monetary policy may have more impact because it signals the intentions of the Federal Reserve to tame asset price bubbles. Although the Federal Reserve has not adopted a "leaning against the wind strategy" (Friedrich et al., 2019), Oet and Lyytinen (2017) have shown from a textual analysis of FOMC meetings that policymakers are concerned by financial stability and that it could be an element explaining monetary policy decisions. In addition, Bjørnland and Leitemo (2009) illustrate interdependencies between stock prices and the federal feds funds rate: a shock to stock prices is followed by an adjustment of the monetary policy stance. However, this explanation is challenged by the fact that the FOMC reaction function seems to be asymmetric but in the other direction: Furlanetto (2011) and Ravn (2012) find that the FOMC would be more prone to cut interest rates when stock prices plummet than to increase them in case of a boom.[28]

Figure 6 – The response of our composite index for stock price imbalances



Note: Estimates based on equation (2) over the sample January 1986 - March 2019. The dependent variable is the composite index for stock price imbalances. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

On the other hand, restrictive monetary surprises may be interpreted as a bad news from market investors and turn down their expectations of future price dynamics. The finance literature has highlighted asymmetric patterns in stock returns (see e.g. Braun et al., 1995, Veronesi, 1999, Bekaert and Wu, 2000) while some contributions related to behavioural finance emphasized the role of negativity bias to explain the (economic) decisions of individual agents (see e.g. Laakkonen and Lanne, 2009, and Akhtar et al., 2011). In financial markets, downward movements ("bad news") are followed by higher market volatility than upward movements ("good news").

3. Investigating the state-dependence of the asymmetric effect

In the previous section, we have provided evidence that the effect of monetary policy on stock price bubbles is asymmetric with restrictive monetary policies having larger effects than expansionary policies. We explore further this result by investigating whether this asymmetric effect is state-dependent. More precisely, we assess whether the response of stock price bubbles change during monetary policy tightening (respectively expansionary) cycles, during credit booms (respectively periods of credit busts), during periods of economic expansion (respectively economic slowdown) and during periods of stock price booms (respectively periods of busts).

To do so, equation (2) is augmented with a second latent variable such that equation (7) has a triple interaction term and the simple interaction terms between our 3 variables of interest – our exogenous instrument for monetary shocks ϵ_t, the latent dummy variable for expansionary and restrictive monetary policy I_t, and the latent dummy variable for the state of the economy, $\mathrm{S}_{j,t}$:

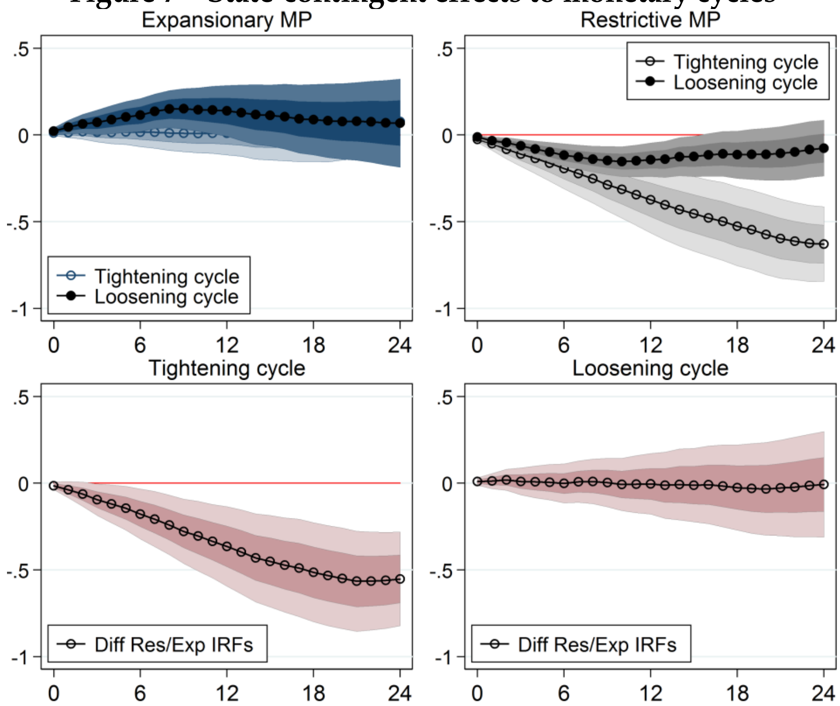
$$\begin{aligned} y_{t+h} = \alpha_h + \beta_{\mathrm{T},h}(\epsilon_t.\mathrm{I}_t.\mathrm{S}_{j,t}) + \beta_{\Pi1,h}(\epsilon_t.\mathrm{S}_{j,t}) + \beta_{\Pi2,h}(\epsilon_t.\mathrm{I}_t) + \beta_{\Pi3,h}(S_{j,t}.\mathrm{I}_t) \\ + \beta_{\epsilon,h}\epsilon_t + \beta_{\mathrm{I},h}\mathrm{I}_t + \beta_{\mathrm{S},h}\mathrm{S}_{j,t} \\ + \phi_{\Pi1,h}(X_t.\mathrm{S}_{j,t}) + \phi_{\Pi2,h}(X_t.\mathrm{I}_t) + \phi_{X,h}X_t + \eta_{t+h} \end{aligned} \tag{7}$$

[28] Those results are consistent with the Fed's put emphasized by Cieslak and Vissing-Jorgensen (2020).

where y_{t+h} is the first principal component measure of stock price imbalances at different horizons h, $S_{j,t}$ the latent dummy variable for the state of the economy will take different forms according to the monetary, credit, business or stock price cycles $j=\{M,C,B,S\}$.[29] The controls are interacted with both the sign of monetary shocks I_t and the state of the economy $S_{j,t}$. Equation (7) is estimated with OLS over a sample from January 1986 to March 2019. Heteroskedasticity and autocorrelation robust Newey-West standard errors are computed.

The first question we explore is whether monetary cycles influence the asymmetric effects of monetary policy on stock price bubbles. It has indeed been shown by Detken and Smets (2004), Bordo and Wheelock (2007), Ahrend et al. (2008) and Bianchi et al. (2016) that asset price booms generally arise when monetary policy is loosening. However, this correlation does not entail a causal effect of monetary policy decisions. Consequently, we assess whether the effect of monetary policy – both expansionary and restrictive decisions – is different when the stance of monetary policy is easing or tightening.

Figure 7 – State-contingent effects to monetary cycles



Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (7) over the sample January 1986 - March 2019. The dependent variable is the composite index for stock price imbalances. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

To do so, we estimate a standard Taylor rule by regressing the policy rate on contemporaneous and lagged values of inflation, real GDP growth and unemployment. The residuals are autocorrelated and show large swings above and below zero that can be matched to periods during which the policy rate was above the fitted value of this simple model or below (see Figure E in the Appendix). Although this representation of the reaction function is simplified, it gives some insight of whether monetary policy was in a tightening

[29] In this section, equation (7) is estimated with the composite indicator for the sake of parsimony.

or loosening cycle. The variable $S_{M,t}$ therefore equals 1 when residuals are positive, so during tightening cycles, and 0 when residuals are negative, so during loosening cycles.[30]

The response of stock price bubbles may not only reflect the asymmetry to expansionary and restrictive shocks but also the environment in which these shocks are implemented. An expansionary shock can arise from a stronger decrease in the policy stance than expected or from a smaller increase in the policy stance than expected. The upper panels of Figure 7 present the effects of expansionary and restrictive monetary surprises during both phases of monetary cycles. We observe that the effects of expansionary monetary surprises (top left) are significant only in loosening cycles but the effects are not significantly different between loosening and tightening cycles. At the opposite, the effects of restrictive monetary surprises (top right) on stock price bubbles are negative and significant in both tightening and loosening cycles. We also find that the effect of restrictive monetary surprises is much stronger in tightening cycles than in loosening cycles. A tighter than expected monetary policy stance during a tightening cycle may convey a stronger signal on policymakers' preferences and therefore on the outlook of asset prices, such that it contributes to a greater reduction in the size of the bubble.

The bottom panels of Figure 7 show the difference between the effects of restrictive and expansionary monetary surprises for both phases of monetary cycles. The asymmetric effects of monetary policy are especially strong during tightening cycles (bottom left): expansionary policies have no effects on stock price bubbles (upper left) but restrictive policies have strong negative effects (upper right). At the opposite, there is no evidence of such asymmetry during loosening cycles (bottom right). This result suggests that stock price bubbles are much more affected in periods during which policy tightens. This result does not seem to be driven by a potentially different frequency and standard deviation of expansionary and restrictive monetary shocks in both states (see Table C in the Appendix).

The second question we explore is whether business cycles influence the asymmetric effects of monetary policy on stock price bubbles. It has indeed been shown that monetary policy would have a different effect on the output during periods of expansion and slowdown. The conclusion remains yet unclear as Weise (1999) or Lo and Piger (2005) find that the response of output to monetary policy shocks is stronger when output growth is low while Tenreyro and Thwaites (2016) find that monetary policy has less effect during recessions. To assess whether the effect of monetary policy on stock price bubbles during periods of expansion and slowdown, we compute an output gap measure using the one-sided Christiano-Fitzgerald trend/cycle decomposition filter on the level of real GDP (see Figure E in the Appendix for the time series of the output gap). The variable $S_{B,t}$ equals 1 when the output gap is positive and 0 when the output gap is negative.

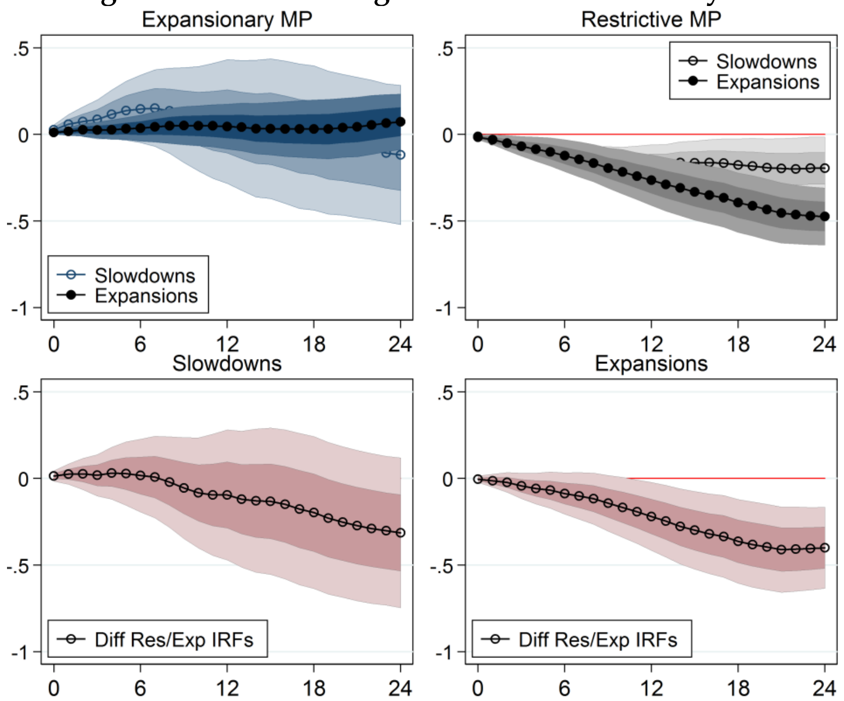
The upper panels of Figure 8 plot the effects of expansionary and restrictive monetary surprises during both phases of business cycles. We observe that the effects of expansionary monetary surprises (top left) are not significant. It is noteworthy that the effect of expansionary policies during slowdowns is much more uncertain (the confidence intervals are much larger than for any other specifications) suggesting a limited ability of monetary policy to affect stock price bubbles during these episodes. At the opposite, the effects of restrictive monetary surprises (top right) on stock price bubbles are negative and significant in both slowdown and expansion cycles. Again, the point estimates of the effect of restrictive

[30] Those cycles are illustrated by Figure E in the Appendix.

monetary surprises are larger in expansions than in slowdowns but they are not statistically different one from the other.

The bottom panels of Figure 8 show the difference between the effects of restrictive and expansionary monetary surprises for both phases of business cycles. The asymmetric effects of monetary policy are confirmed in expansions only: restrictive monetary surprises have more impact on stock price bubbles than expansionary ones. This result suggests that restrictive monetary policy has more traction on stock price bubble processes during expansions. This is consistent with Tenreyro and Thwaites (2016) who find that the effects of monetary policy are less powerful in recessions. As for the effect of monetary surprises during monetary cycles, those results are not driven by the frequency and standard deviation of monetary shocks during economic expansions and slowdowns (see Table C in the Appendix).

Figure 8 – State-contingent effects to business cycles

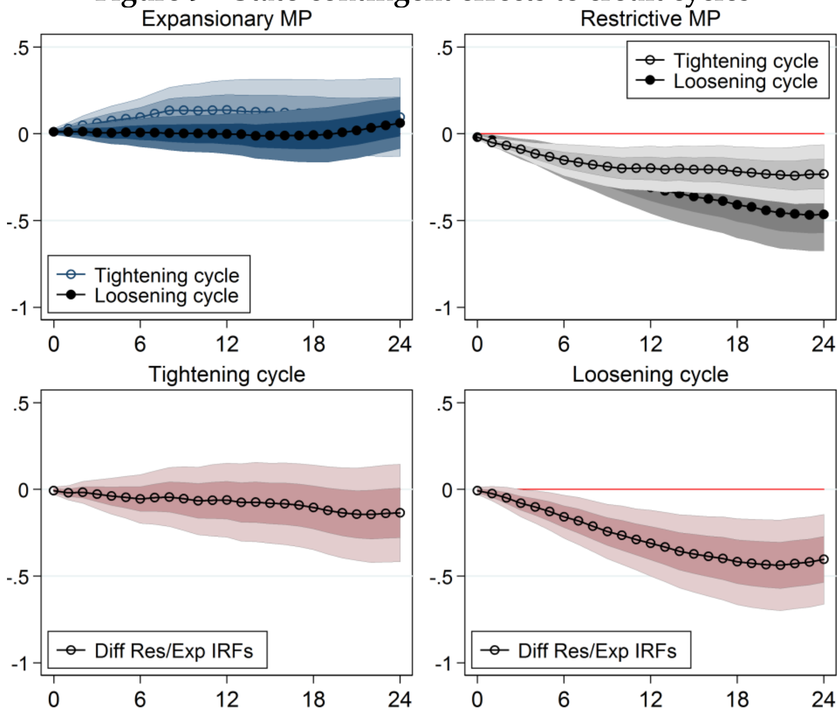


Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (7) over the sample January 1986 - March 2019. The dependent variable is the composite index for stock price imbalances. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

The third question we explore is whether credit cycles influence the asymmetric effects of monetary policy on stock price bubbles. Jordà et al. (2015) have notably emphasized the joint dynamic of bubbles and credit. It is therefore crucial to assess whether the effect of monetary policy is amplified during credit booms. To capture credit cycles, we compute the year-over-year growth rate of loans – including notably commercial and industrial loans, real estate and consumer loans – of all commercial US banks and the variable $S_{C,t}$ equals 1 when the credit growth rate is above its sample average, so during period of loosening credit conditions, and 0 when the credit growth rate is below its sample average, so during cycles of tightening credit conditions. The sample average of credit growth represents the steady state level of credit growth (see Figure E in the Appendix for the time series of the detrended credit growth).

The upper panels of Figure 9 display the effects of expansionary and restrictive monetary surprises during both phases of credit cycles. The effects of expansionary monetary surprises (top left) are not significantly different from zero and not significantly different between loosening and tightening credit cycles. At the opposite, the effects of restrictive monetary surprises (top right) on stock price bubbles are negative and significant in both tightening and loosening credit cycles. The point estimate of the effect of restrictive monetary surprises is larger in loosening cycles than in tightening cycles but both are not statistically different.

Figure 9 – State-contingent effects to credit cycles



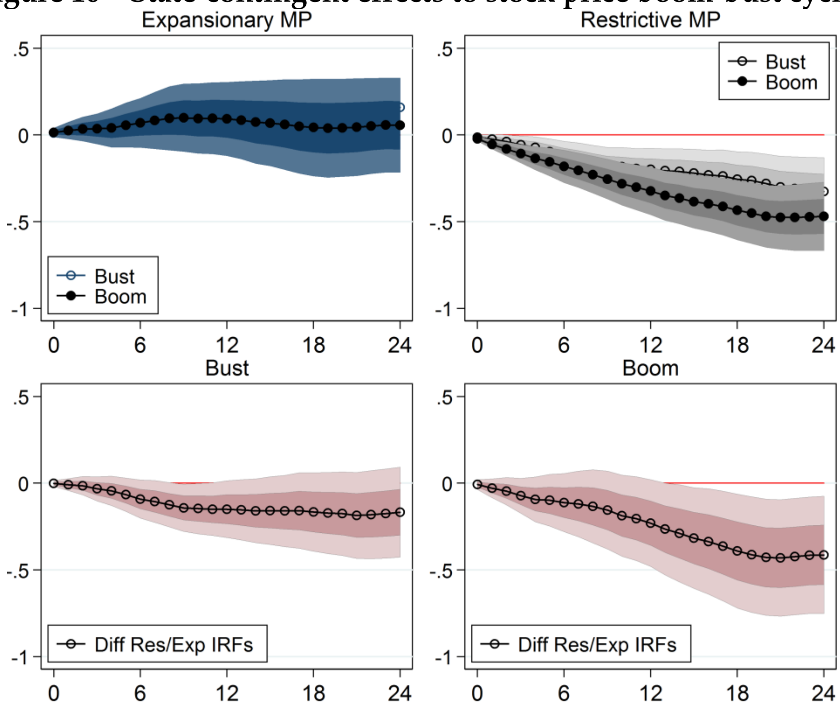
Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (7) over the sample January 1986 - March 2019. The dependent variable is the composite index for stock price imbalances. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

The bottom panels of Figure 9 show the difference between the effects of restrictive and expansionary monetary surprises for both phases of credit cycles. The asymmetric effect of monetary policy is confirmed in loosening credit cycles only: restrictive monetary surprises have more impact on stock price bubbles than expansionary policies. This finding suggest that policymakers may have more traction on stock price bubbles when they face leveraged bubbles, possibly because they can affect both the stock price bubble process and the funding conditions of investors. It should be added that those results are not driven by the distribution of shocks during the cycles. There is no striking difference in the frequency of surprises between credit booms and credit busts and standard-deviations are also close (see Table C in the Appendix).

The fourth question we explore is whether stock price boom-bust cycles influence the asymmetric effects of monetary policy on stock price bubbles. The latent dummy variable $S_{S,t}$

equals 1 when our first principal component measure of stock price imbalances is positive and 0 when our PCA measure is negative. The upper panels of Figure 10 present the effects of expansionary and restrictive monetary surprises during both boom and bust phases of stock price cycles. The effects of expansionary monetary surprises (top left) are not significantly different from zero and not significantly different between booms and busts. At the opposite, the effects of restrictive monetary surprises (top right) on stock price bubbles are negative and significant in both boom and bust phases. This result suggests that the effects of restrictive policies are more potent whatever the dynamics on stock markets.

Figure 10 – State-contingent effects to stock price boom-bust cycles



Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (7) over the sample January 1986 - March 2019. The dependent variable is the composite index for stock price imbalances. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

The bottom panels of Figure 10 show the difference between the effects of restrictive and expansionary monetary surprises for both phases of stock price cycles. The asymmetric effects of monetary policy are confirmed in both booms and busts: restrictive monetary surprises have more impact on stock price bubbles than expansionary ones. However, the asymmetry is more pronounced during booms (bottom right) than during busts (bottom left). These results may signal that restrictive surprises would have more ability to reverse markets' expectations during asset price booms suggesting a role for central banks to tame bubbles. This result supports the hypothesis of the negativity bias explanation of the non-linearity. Here again, the estimates seem not driven by the frequency and standard deviation of monetary shocks (see Table C in the Appendix).

Overall, the main result of Figure 6 is confirmed: restrictive shocks have significant effects in all situations but not expansionary policies. The asymmetry occurs during tightening policy cycles, loosening credit conditions cycles and economic expansions, and is more pronounced during stock price booms. This suggests that the effect of monetary policy on stock price bubbles works more powerfully when policy is countercyclical.

Restrictive monetary policy is more powerful during tightening monetary cycles and economic expansions. As a consequence, it implies that easing less is not equivalent to tightening more. Two restrictive surprises may not have the same effect on stock price bubbles. Monetary policy can be tightened during loosening monetary cycles if the interest rate is less decreased than expected. During tightening cycles, a restrictive surprise occurs if the interest rate increases more than expected. We show that restrictive policy would be more effective to reduce the size of stock price bubbles during tightening cycles. This result has important implications for policymakers aiming to engage in a leaning against the wind strategy. In such a case, monetary policy is expected to ease less or tighten more – compared to a situation when the central bank only focus on price and output – when financial imbalances are identified. According to our results, the central bank can achieve the same reduction of stock price imbalances with a softer reaction during tightening cycles. During loosening cycles, the reaction should be more important to achieve the same reduction of the bubble. If we represent monetary policymaking by an augmented Taylor rule augmented with financial imbalances, our results imply that the coefficient associated to stock price imbalances could change along the monetary cycle. The response to stock price imbalances could be lower during tightening cycles and higher during easing cycles. This asymmetry should be taken into account when assessing the cost-benefit analysis of the leaning against the wind strategy.

4. Robustness analysis

The previous developments emphasize the asymmetric effect of monetary policy on stock price bubbles. There is not only a difference between expansionary and restrictive monetary policy but we have also illustrated that the effects of monetary policy depend on the monetary, credit and business cycles. The sensitivity of these results can be assessed to show that they are not driven by the hypotheses regarding the sample period, the stock price index that is considered, the specifications of the underlying models for bubbles' identification, the identification of monetary policy shocks and the specification of the local projection estimations.

4.1. Monetary policy and stock price bubbles before the zero lower bound

All estimations so far have been performed over a sample covering the financial crisis and the zero lower bound (ZLB) period. During this period, central banks resorted to a larger set of policy instruments and implemented expansionary measures. However, the ZLB period may be too short to infer the dynamic effect of monetary policy on stock price bubbles. Besides, monetary policy was mainly expansionary during the period so that it would be hazardous to disentangle the asymmetric effects of these non-standard tools. Conversely, it is worth restricting the analysis to the pre-ZLB period when the Federal Reserve was using changes in the Fed Funds rate target as its sole policy instrument.[31]

To that end, we estimate equation (2) over a restricted sub-sample, which spans the time period from January 1986 to June 2008 and should capture the effect of monetary policy

[31] The question of the risks associated with low interest rates has been raised by Del Negro and Otrok (2007), Taylor (2009), Dokko et al. (2011) and Juselius et al. (2016) among others.

during "normal" times when the Fed Funds rate is the unique indicator of the monetary policy stance. We use the daily change in the Federal funds futures as in Kuttner (2001) together with the daily change in the 2-year rate since it captures the surprise in the policy decision on the day of FOMC decisions but also the surprise component of announcements that provide information on the future likely path of policy.

Impulse responses are displayed on Figure F in the Appendix and show that interest rate decisions had asymmetric effects on stock market bubbles during the pre-crisis period. With the change in 2-year rates, expansionary policy has no significant impact on bubbles. Increase in the policy rate would conversely decrease the size of bubbles and the asymmetry between positive and negative surprises is significant. Turning to the response to a change in the Federal Fed funds rate, we still find a significant difference between positive and negative surprises. It is consequently not clear whether the low interest rate policy during the early 2000's has been responsible for the stock market boom but the rise in interest rate after 2004 may have contributed to the decrease of the bubble after 2007.

4.2. Alternative stock price indices

The previous estimations have been performed with the S&P500 as a benchmark for the US stock price index. We first check that the results also hold for the other major stock price index: the Dow Jones and the NASDAQ. The structural, data-driven and statistical models have been estimated with these stock price indices and we have computed the composite indicator with the PCA. The identification of stock price booms and busts with the Dow Jones and the NASDAQ provides information extremely close to the picture that emerged with the S&P 500 (Figure D in the Appendix).

We do not observe differences in the responses of those stock price indices as we still find that restrictive monetary policy has more impact on stock price bubbles than expansionary decisions (see Figure G). The magnitude of the response is also slightly weaker for both shocks – compared to response of the S&P500.

4.3. Alternative bubble indicators

As the measure of the bubble component of stock prices is a key challenge, we have tested for alternative measure of the bubble and for alternative models' identification. As pointed out by Bunn and Shiller (2014), the level of the CAPE ratio may be influenced by the corporate payout policy and notably by share buybacks. To account for the bias, we use the total return CAPE (TR-CAPE) index available from Shiller. Besides, it may also be stressed that stock prices are forward-looking so that, we would better take into account the expected earnings instead of current ones. The price-earnings ratio has also been calculated with earnings expected at 12 months by brokers and analysts. In addition, we have estimated equation (5) with the 10-year moving average of dividends to strip out cycles in the spirit of the CAPE. The response of these indicators is displayed in Figure H and shows that the response is significantly different for expansionary and restrictive monetary policy. The impact of expansionary policy is significant with the TR-CAPE index and with the indicator based on the 10-year moving average of dividends but it is almost totally muted for expansionary shocks when we consider the price to forward-earnings ratio. The difference between the responses to expansionary and restrictive policy is significant in all three cases.

Beyond the measure of the bubble, the identification of the bubble component in the DCF, data-driven and statistical models may also matter. On the one hand, we have estimated error-correction models for equations (4) and (5) to capture the possibility that stock prices may be a combination of a long-run trend and short-run dynamics. For the statistical model, we have used the Hodrick-Prescott filter instead of Christiano-Fitzgerald. The response of monetary policy is computed for a composite indicator stemming from a principal component analysis. Moreover, in the baseline scenario, we have considered medium-term deviations identified in each model to account for persistent deviations of stock prices. Consequently, the residuals from the DCF and data-driven models and the cycle component from the Christiano-Fitzgerald have been transformed and smoothed with a moving-average on 49 months. To avoid such a transformation, we have computed the PCA from raw residuals and the cycle component. The results are reproduced in Figure I. They show that the asymmetry between the effects of expansionary and restrictive shocks is significant.

4.4. Alternative measures of monetary policy shocks

The developments on the inference of causal effects of monetary policy has highlighted several approaches to identify monetary policy shocks from narrative approaches to high-frequency instruments as used in the baseline estimation with the monthly sum of daily changes in the 2-year nominal US sovereign yield on FOMC announcement days. Figure J illustrates the response of our composite measure of stock price bubbles to alternative measures of monetary policy shocks.

First, we have computed the daily change in the 1-year nominal US sovereign yield on FOMC announcement days instead of the daily change in the 2-year. The "path factor" of Gürkaynak et al. (2005) is estimated with contracts up to one year. Second, we have extended the window and considered a two-day change in the the 2-year nominal US sovereign yield on FOMC announcement days. The implicit assumption is that the full reaction to a policy announcement might not be instantaneous, particularly for longer term horizons. This could be because investors are uncertain about the implications of the released news and update their beliefs as asset prices and volumes, and media reports reveal others' beliefs. Thus, it could take some time for markets to process the information content of a policy decision, a statement or an economic report.

Another concern – mentioned earlier - is that monetary surprises may be confounded by some implicit release of central bank private information (Romer and Romer, 2000). Therefore, we also use the monetary shock series estimated by Miranda-Agrippino and Ricco (2020). They propose to combine the high-frequency approach with the narrative approach and identify shocks from the regression of market-based surprises on their own lags and central bank's forecasts. The impulse responses obtained with the local projection indicate that responses to positive shocks on the interest rate are significant (Figure J). A restrictive monetary would always deflate stock price bubbles. The difference between the responses is significant reinforcing the results according to which the effect of monetary policy on stock price bubbles is asymmetric.

4.5. Alternative specifications of local projections

Finally, we assess whether the results are sensitive to the specification of the local projections and consider additional control variables in equation (2). First, we include three lags of the dependent variable – the stock price bubble – and of the exogenous instrument for the monetary policy stance. Second, we include a financial stress indicator – the VIX – in order to control for the effect of a potential central bank put (see Cieslak and Vissing-Jorgensen, 2020) – the link between monetary policy and stock price imbalances may be more pronounced during financial crises. In both cases, restrictive monetary policy has more effect on the bubble indicator than expansionary monetary policy (Figure K).

5. Conclusion

This paper relaxes the common assumption in the literature that the effect of monetary policy on stock price bubbles is linear. We estimate the dynamic impact of monetary policy over a 2-year horizon using local projections. As a measure of stock price imbalances, we use Shiller (2000)'s measure of over- and under-valuations and our own estimates of stock price bubbles derived from structural, data-driven and statistical approaches.

We find that restrictive policies have always a significant effect on stock price imbalances. The evidence is less supportive of a significant effect of expansionary policy suggesting that monetary policy would be able to reduce the size of stock price bubbles but not necessarily to inflate them. The key finding is that the effects of monetary policy are asymmetric: the effects of restrictive monetary policy are more powerful than the effects of expansionary policies. We also find that the asymmetry is stronger during tightening monetary cycles but disappears in loosening cycles. Besides, we find that the asymmetry is more pronounced during loosening credit condition cycles and during expansions than during slowdowns. We also find that the asymmetry is stronger during booms than busts.

Finally, the results suggest that restrictive surprises have more effect when they are implemented during tightening cycles and economic expansions than restrictive surprises implemented during loosening cycles and economic slowdowns. This effect is important to assess the cost-benefit advantages of the leaning against the wind strategy and should be accounted for in future research.

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APPENDIX
FOR ONLINE PUBLICATION

Table A. Data sources and description

	Description	Source	Frequency
	Asset prices		
Stock prices	S&P 500	Datastream	Monthly
Stock prices	Dow Jones	Datastream	Monthly
Stock prices	NASDAQ	Datastream	Monthly
CAPE	Cyclically-Adjusted Price Earning ratio	Shiller (2000)	Monthly
Monetary surprises	2-year nominal yield (daily variation)	Datastream	Daily
Monetary surprises	Futures-based Fed Funds Rate surprises (daily variation)	Kuttner (2001)	Daily
Monetary surprises	Projection of intraday (30min window) surprises in the price of the fourth Fed Funds futures (FF4) on Greenbook forecasts.	Miranda-Agrippino and Ricco (2020)	Monthly
Monetary surprises	5-year nominal yield (daily variation)	Datastream	Daily
Dividends	Paid dividends by corporations	BEA	Quarterly
Discount factor	10 year treasury bond interest rates	Datastream	Monthly
Risk premium	Volatility Index	CBOE	Monthly
Income	Real disposable income	BEA	Quarterly
Real GDP	Real GDP	BEA	Quarterly
IndPro	Industrial production	BEA	Monthly
Oil prices	Oil prices	Datastream	Monthly
Inflation	Inflation	BEA	Monthly
Confidence indicators	Confidence indicators for consumers and firms	Conference Board & ISM	Monthly
Financial stress	Kansas City Financial indicator	FRED	Monthly
Monetary Aggregate	M2	Datastream	Monthly
Credit Aggregate	Credits granted by commercial banks	Datastream	Monthly

Note: All nominal variables are deflated by the CPI.

Table B. Parameter estimates for Figures 2 and 6

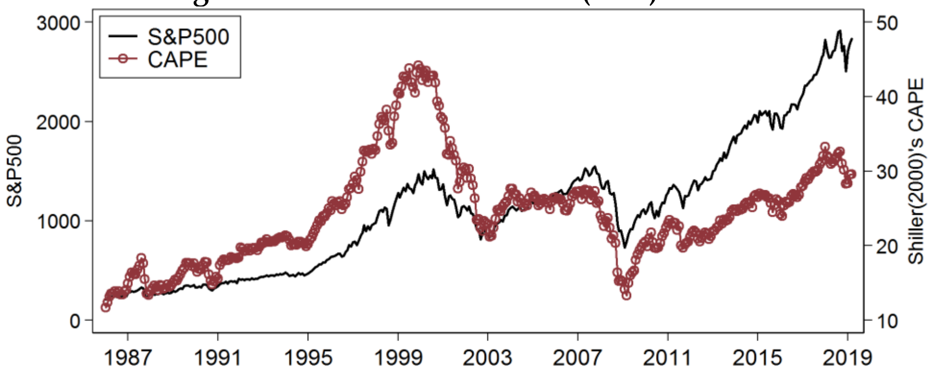
	CAPE					PCA_Stock				
	F0	F6	F12	F18	F24	F0	F6	F12	F18	F24
$I_t \times \epsilon_t$	0.060***	0.218***	0.346***	0.370***	0.362***	0.030***	0.190***	0.310***	0.391***	0.499***
	[0.02]	[0.05]	[0.07]	[0.07]	[0.07]	[0.01]	[0.05]	[0.08]	[0.10]	[0.10]
ϵ_t	-0.031**	-0.154***	-0.236***	-0.258***	-0.262***	-0.019***	-0.156***	-0.270***	-0.357***	-0.419***
	[0.01]	[0.03]	[0.05]	[0.06]	[0.05]	[0.00]	[0.03]	[0.05]	[0.06]	[0.07]
I_t	0.005	-0.099**	-0.147**	-0.258***	-0.337***	-0.014	-0.140**	-0.273**	-0.425***	-0.512***
	[0.02]	[0.05]	[0.06]	[0.08]	[0.10]	[0.01]	[0.07]	[0.11]	[0.13]	[0.13]
Constant	0.062*	0.475***	0.925***	1.328***	1.800***	0.023***	0.169***	0.307***	0.443***	0.572***
	[0.03]	[0.09]	[0.13]	[0.16]	[0.19]	[0.01]	[0.04]	[0.06]	[0.08]	[0.09]
X_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	398	392	386	380	374	374	368	362	356	350
R^2	0.99	0.88	0.75	0.62	0.49	0.99	0.81	0.52	0.29	0.23
Partial R^2	0.04	0.09	0.11	0.11	0.10	0.04	0.06	0.07	0.09	0.12

Note: Standard errors in brackets. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Estimates are based on equation (2) over the sample January 1986 - March 2019. The dependent variable is Shiller's (2015) CAPE ratio (left panel) and the first principal component of the three approaches to measure bubbles (structural, data-rich and statistical). Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015).

Table C. Descriptive statistics for monetary surprises

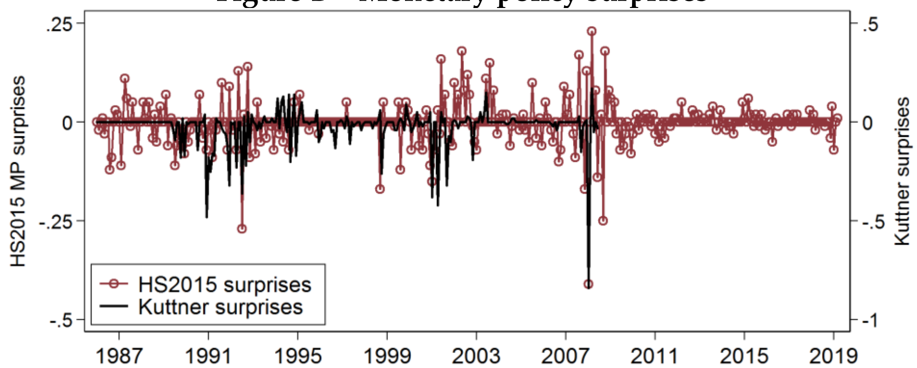
		Obs	Mean	Std. Dev.	Min	Max
Overall						
All		399	0.00	0.05	-0.41	0.23
Restrictive		92	0.05	0.05	0.01	0.23
Expansionary		104	-0.06	0.06	-0.41	-0.01
Monetary cycles						
Tightening	Restrictive	49	0.05	0.05	0.01	0.23
	Expansionary	53	-0.07	0.07	-0.41	-0.01
Loosening	Restrictive	43	0.06	0.05	0.01	0.18
	Expansionary	51	-0.04	0.04	-0.27	-0.01
Credit cycles						
Tightening	Restrictive	40	0.05	0.04	0.01	0.18
	Expansionary	55	-0.05	0.04	-0.27	-0.01
Loosening	Restrictive	52	0.06	0.05	0.01	0.23
	Expansionary	49	-0.07	0.07	-0.41	-0.01
Business cycles						
Slowdowns	Restrictive	42	0.05	0.05	0.01	0.18
	Expansionary	52	-0.05	0.03	-0.15	-0.01
Expansions	Restrictive	50	0.05	0.05	0.01	0.23
	Expansionary	52	-0.07	0.07	-0.41	-0.01
Stock price cycles						
Busts	Restrictive	44	0.05	0.05	0.01	0.23
	Expansionary	54	-0.06	0.07	-0.41	-0.01
Booms	Restrictive	48	0.05	0.04	0.01	0.18
	Expansionary	50	-0.06	0.05	-0.27	-0.01

Figure A – S&P500 and Shiller (2000)'s CAPE



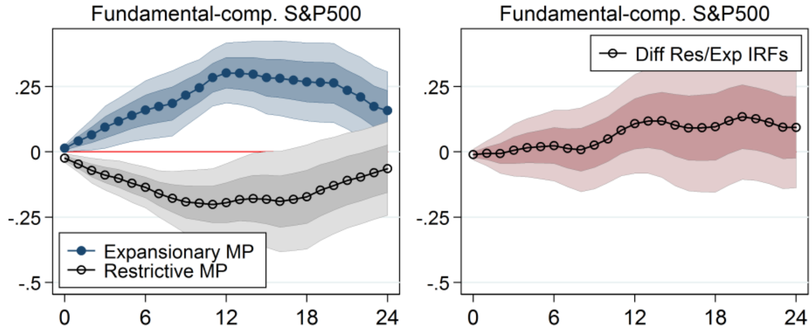
Note: This figure presents the time series of S&P500 (left scale) and Shiller (2000)'s CAPE (right scale).

Figure B – Monetary policy surprises



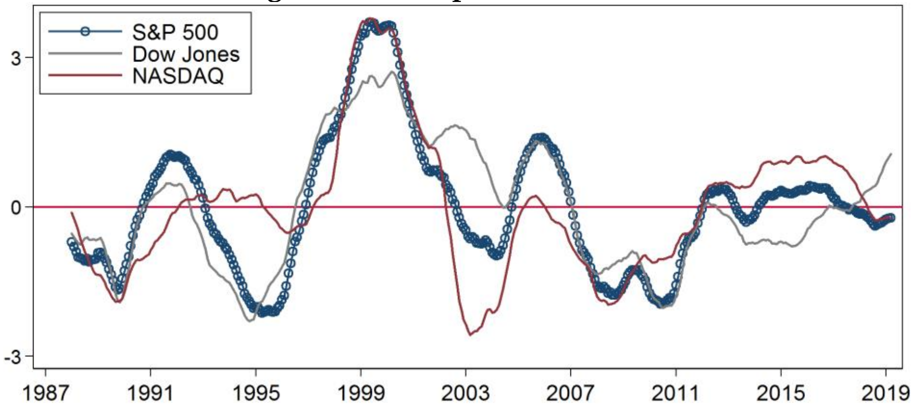
Note: Monetary shocks are computed as the daily change in the 2-year nominal US sovereign yield on FOMC announcement days (Hanson-Stein surprises) and as the daily change in the Federal funds rate futures (Kuttner surprises).

Figure C – The effect of monetary policy on the fundamental-component of SP500 prices



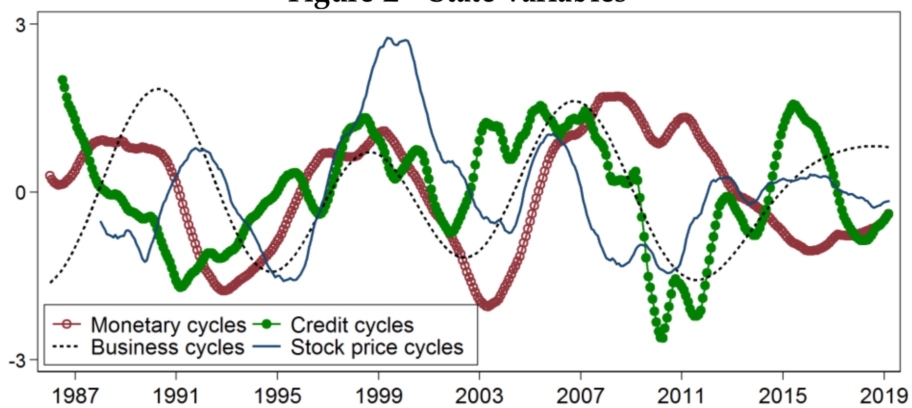
Note: Estimates are based on equation (2) over the sample January 1986 - March 2019. The dependent variable is the fitted value of the S&P500 index based on real earnings. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

Figure D – Composite indicators



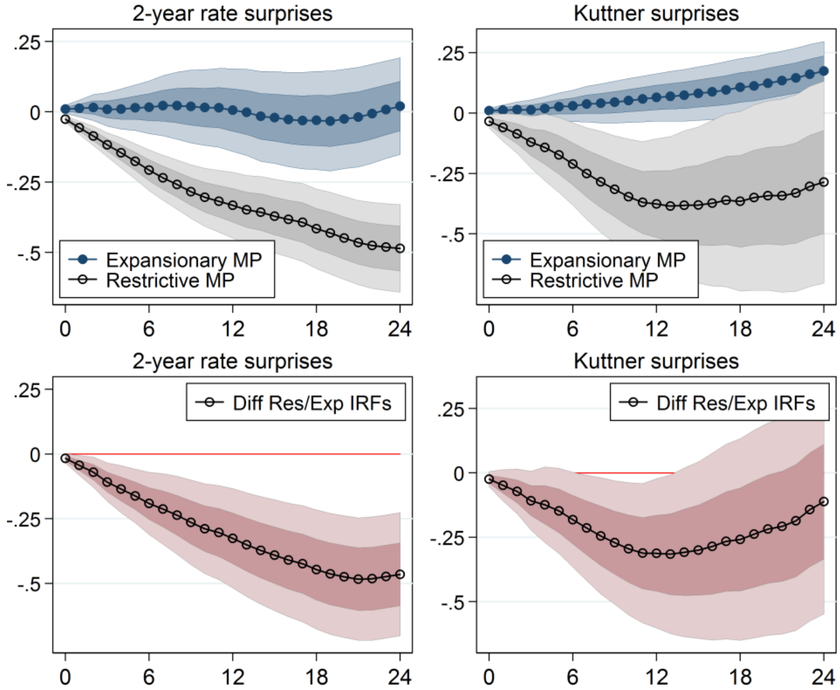
Note: Composite bubble indicators are computed as the first principal component from a principal-component analysis (PCA) of the three individual bubbles measures (the structural, data-driven and statistical approach) for the S&P500, the Dow Jones and the NASDAQ.

Figure E – State-variables



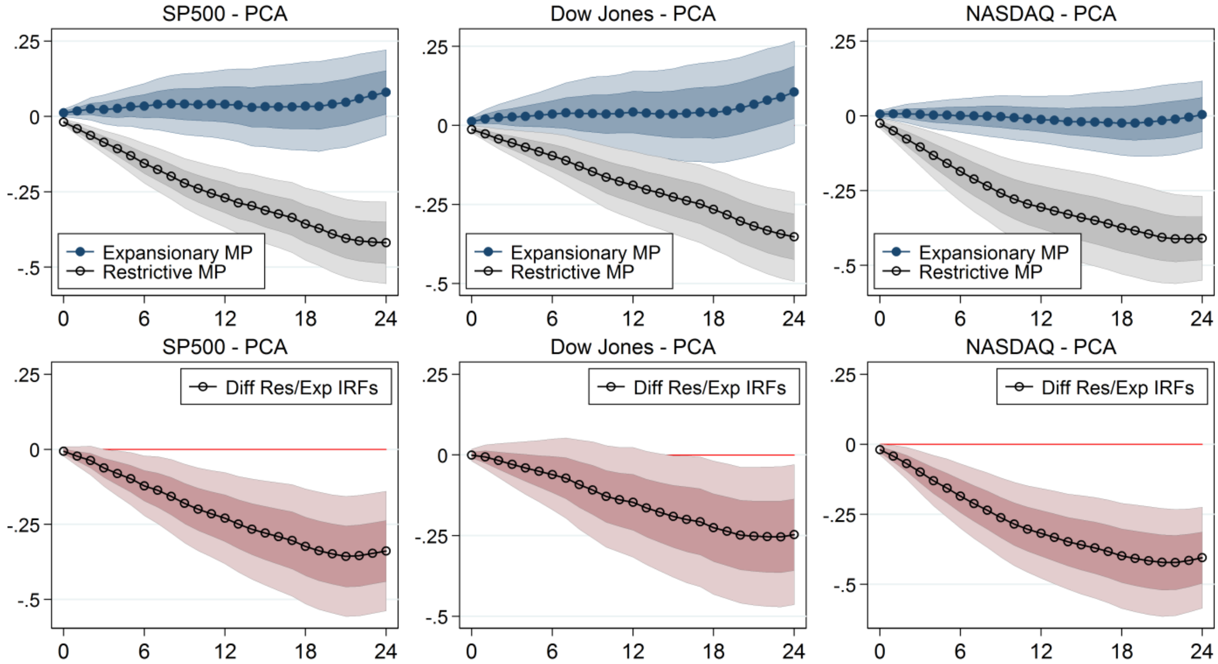
Note: Monetary cycles are estimated as the residuals of a standard Taylor rule regressing the policy rate on contemporaneous and lagged values of inflation, real GDP growth and unemployment. Credit cycles are computed as the detrended annual credit growth rate of loans of all US commercial banks. Business cycles are estimated using Christiano-Fitzgerald decomposition on the level of real GDP. Stock price cycles are based on our composite indicator computed from a PCA of the three approaches to measure bubbles (structural, data-rich and statistical).

Figure F – Asymmetric effects of conventional monetary policy



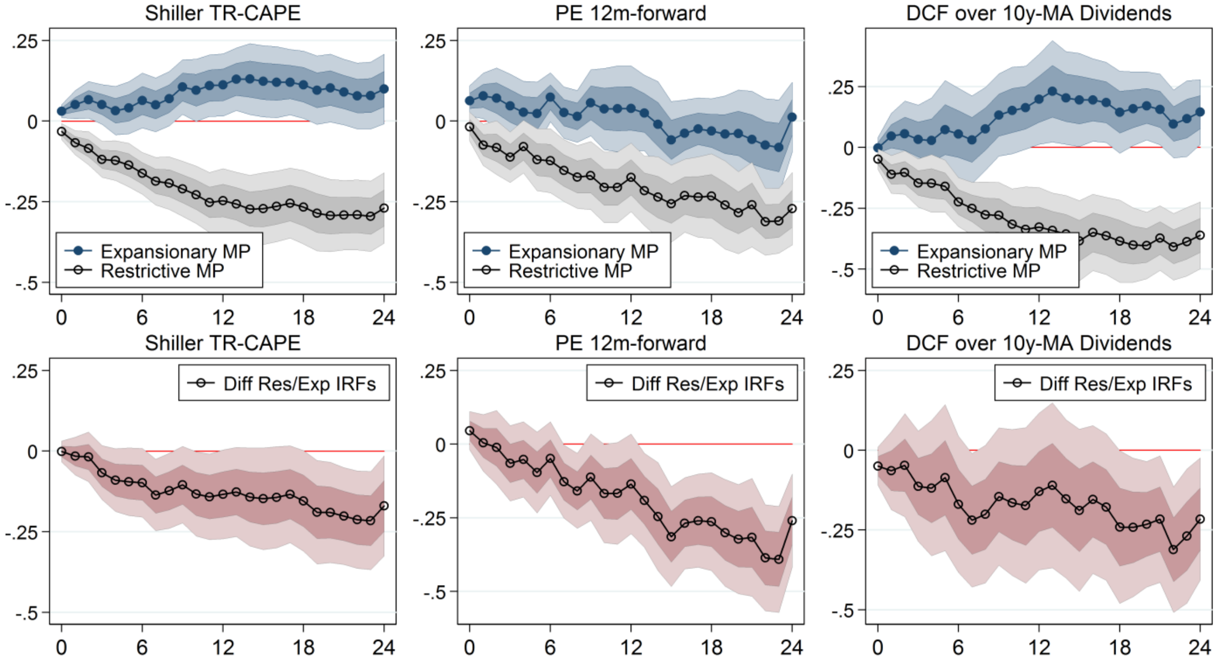
Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (2) over the sample January 1986 – June 2008. The dependent variable is the composite index for stock price imbalances. Monetary surprises are computed as the daily change in nominal 2-year interest rates or in Fed Funds rate futures on the day of the policy decision following Hanson and Stein (2015) and Kuttner (2001) respectively. Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

Figure G – Alternative stock price indices



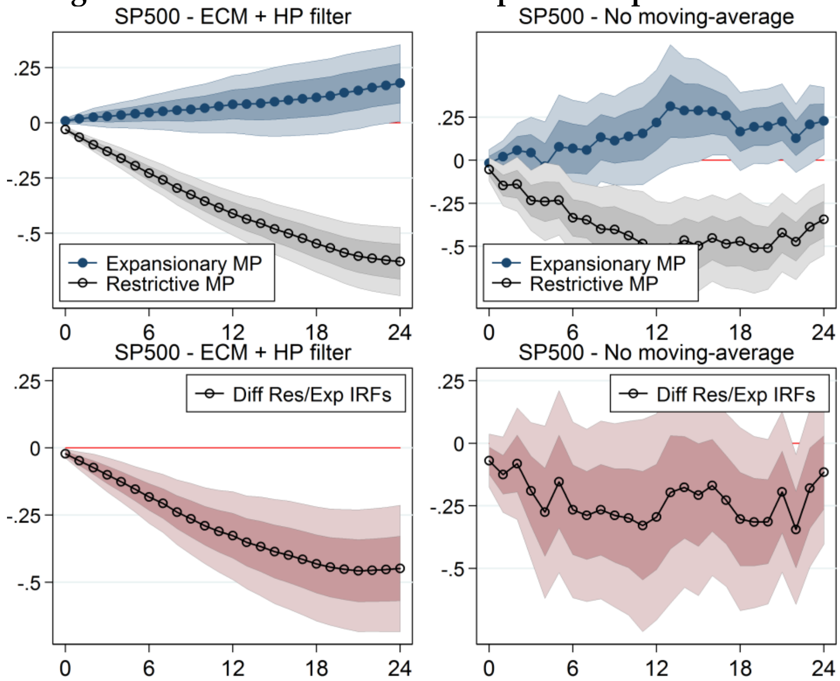
Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (2) over the sample January 1986 - March 2019. The dependent variable is the composite index for stock price imbalances. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

Figure H – Alternative bubble representations



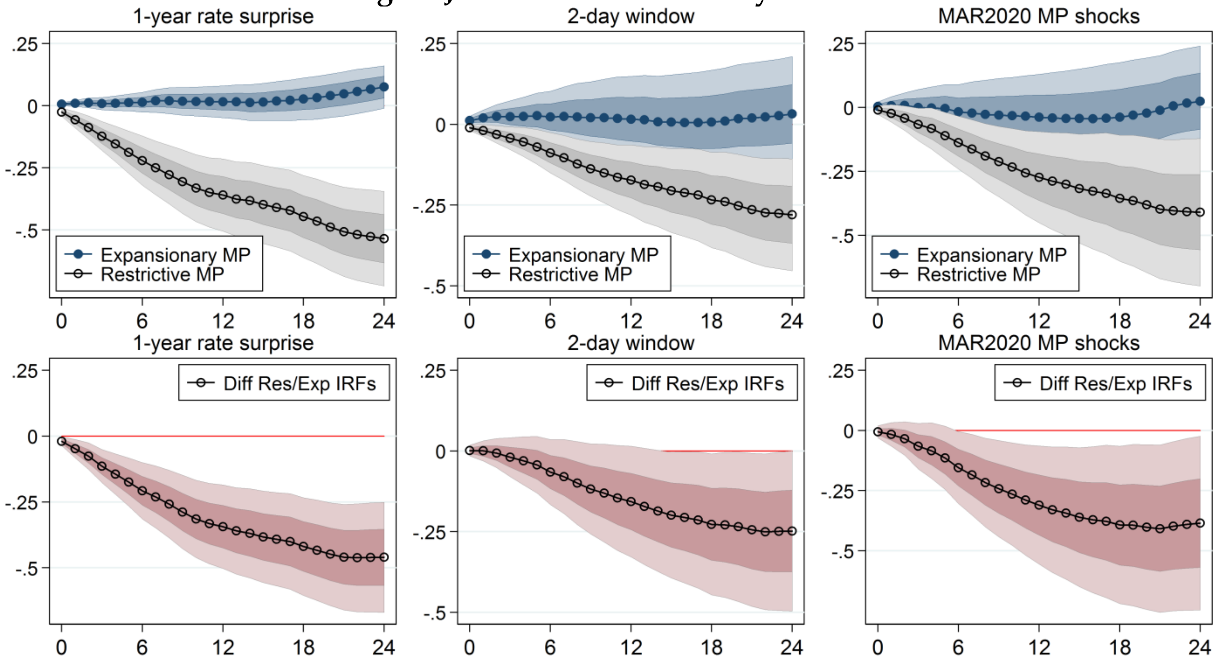
Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (2) over the sample January 1986 - March 2019. The dependent variable is the CAPE measure incorporating the effects of share buybacks (left panel), the 12-month forward PE (middle) and the residuals of a DCF model based on the 10-year moving-average in dividends (right). Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

Figure I – Alternative bubble computation specifications



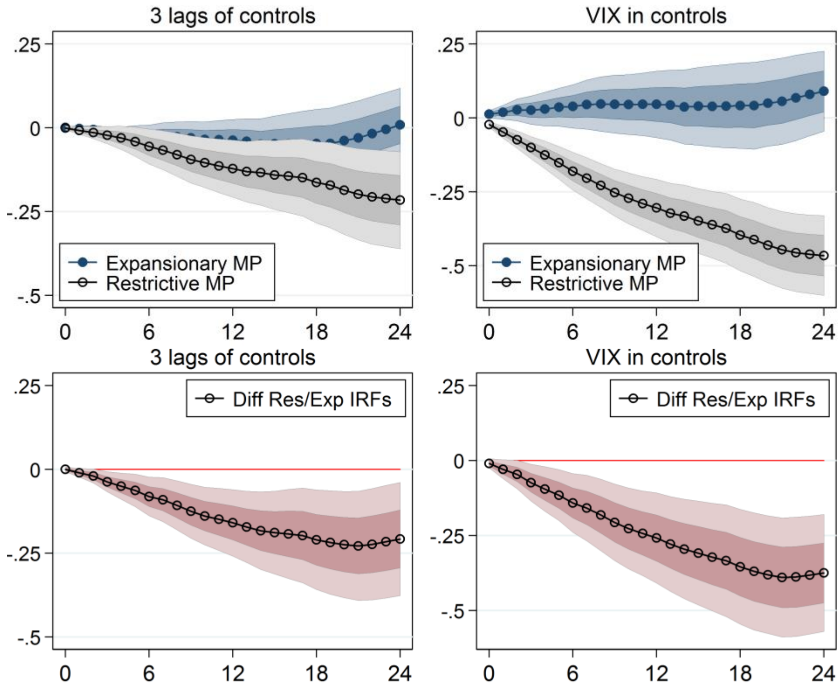
Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (2) over the sample January 1986 - March 2019. The dependent variable is two alternative measures of the composite index for stock price imbalances based on error-correction models for the DCF and data-driven models and the HP filter for the statistical model (left panel) and without the moving-average smoothing of residuals or the cycle component (right panel). Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

Figure J – Alternative monetary shocks



Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (2) over the sample January 1986 - March 2019. The dependent variable is the composite index for stock price imbalances. Monetary surprises are computed as A. the daily change in nominal 1-year interest rates on the day of the policy decision following Hanson and Stein (2015), B. the two-day change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015), and C. the monetary shocks of Miranda-Agrippino and Ricco (2020). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

Figure K – Alternative LP specifications



Note: This figure shows impulse response functions in the upper panel and the difference between expansionary and restrictive responses in the lower panel. Estimates based on equation (2) augmented with additional lags or more Newey-West terms for the correction of standard errors, over the sample January 1986 - March 2019. The dependent variable is the composite index for stock price imbalances. Monetary surprises are computed as the daily change in nominal 2-year interest rates on the day of the policy decision following Hanson and Stein (2015). Shaded areas represent confidence intervals for 1 and 2 robust standard errors.

ABOUT OFCE

The Paris-based Observatoire français des conjonctures économiques (OFCE), or French Economic Observatory is an independent and publicly-funded centre whose activities focus on economic research, forecasting and the evaluation of public policy.

Its 1981 founding charter established it as part of the French Fondation nationale des sciences politiques (Sciences Po), and gave it the mission is to “ensure that the fruits of scientific rigour and academic independence serve the public debate about the economy”. The OFCE fulfils this mission by conducting theoretical and empirical studies, taking part in international scientific networks, and assuring a regular presence in the media through close cooperation with the French and European public authorities. The work of the OFCE covers most fields of economic analysis, from macroeconomics, growth, social welfare programmes, taxation and employment policy to sustainable development, competition, innovation and regulatory affairs.

ABOUT SCIENCES PO

Sciences Po is an institution of higher education and research in the humanities and social sciences. Its work in law, economics, history, political science and sociology is pursued through ten research units and several crosscutting programmes.
Its research community includes over two hundred twenty members and three hundred fifty PhD candidates. Recognized internationally, their work covers a wide range of topics including education, democracies, urban development, globalization and public health.
One of Sciences Po’s key objectives is to make a significant contribution to methodological, epistemological and theoretical advances in the humanities and social sciences. Sciences Po’s mission is also to share the results of its research with the international research community, students, and more broadly, society as a whole.

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