Does Monetary Policy React to Asset Prices? Some International Evidence*

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Previous estimates of the monetary policy response to stock market fluctuations in the United States are found to be sizable and significant once the simultaneous interdependence between stock prices and interest rates is properly taken into account. We show that this result is not confirmed when we apply the analysis to other countries and when we consider an extended sample period including the past decade for the United States. We do not find any response in the European Union and in six inflation-targeting countries, with the exception of Australia. Moreover, we find that the response in the United States declines over time and becomes not statistically significant during the housing bubble period (2003–07).

JEL Codes: E44, G10.

1. Introduction

The recent financial crisis and the previous series of boom-bust cycles in asset prices show that fluctuations in financial variables can have a potentially large impact on the macroeconomic outlook. In fact, large swings in asset prices are often associated with strains in the financial sector and in the real economy (see Borio, Kennedy, and

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Prowse 1994 and Borio and Lowe 2002, among many others). While the main concern of central bankers—the inflation rate—is now low and stable in the main developed economies, financial instability has become a central issue in the formulation of the monetary policy strategy.[1]

A small but growing empirical literature takes a positive perspective and attempts to estimate the monetary policy response to stock market fluctuations. Early contributions either impose a zero interest rate response on impact (see Thorbecke 1997 and Neri 2004, among others, for VAR studies) or find evidence of a very small response, always statistically insignificant and sometimes even negative (see Bernanke and Gertler 2000 for a GMM estimation of Taylor rules). More recent contributions argue that previous studies fail to find a significant interaction between monetary policy and asset prices because they do not properly consider the simultaneous interdependence between interest rates and asset prices. This endogeneity problem is ruled out by assumption in early VAR contributions and is addressed only partially in the estimation of monetary policy rules by using instruments that are likely to be weak. When the endogeneity issue is addressed carefully, a positive and significant response of monetary policy to stock prices emerges, at least in the United States. This is so in Rigobon and Sack (RS, henceforth) (2003) in a VAR identified through heteroskedasticity, in Castelnuovo and Nisticò (2010) in an estimated dynamic stochastic general equilibrium (DSGE) model where monetary policy is allowed to respond to fluctuations in the stock market, in Bjørnland and Leitemo (2009) in a VAR identified using a combination of short-run and long-run restrictions, and also in Chadha, Sarno, and Valente (2004) in a Taylor-rule estimation with GMM where the authors carefully check the quality of the instruments and use an adjusted labor share

[1]The conventional wisdom is that central banks react to asset prices only to the extent that they affect expected inflation and the output gap (Bernanke and Gertler 2001). According to this view, a more aggressive response to asset prices has destabilizing effects because it is difficult to know whether a change in asset prices is due to fundamental factors or not. These claims have not stood unchallenged. Cecchetti et al. (2000), Bordo and Jeanne (2002), and Akram and Eitrheim (2008), among others, provide examples where a more proactive policy has stabilizing effects on the economy.

as the appropriate, and theoretically grounded, proxy for the output gap. Importantly, all these studies argue that the size of the estimated response is compatible with statements made by central bank officials. Therefore, the estimated response would not reflect an attempt to target asset prices, but just an indirect response to the impact that asset prices have on aggregate demand, in particular on consumption and investment decisions.

In this paper we complement previous studies by extending the analysis to several other countries and by using data for the past decade in the United States. Our main result is that there are good reasons to believe that the response to stock prices in normal times has been substantially lower than that found in previous studies.

Two kinds of considerations motivate our analysis. First, previous studies concentrate their attention on the U.S. economy. To the best of our knowledge, only a few contributions consider other countries. Chadha, Sarno, and Valente (2004) also estimate Taylor rules for the United Kingdom and Japan, whereas Bohl, Siklos, and Werner (2007) extend the RS (2003) analysis to Germany over the period 1985–98. In this paper we provide evidence for the European Union (EU) and six inflation-targeting countries (Australia, Canada, New Zealand, Norway, Sweden, and the United Kingdom). Second, none of the studies quoted above consider recent data. Bjørnland and Leitemo (2009) estimate their VAR over the period 1983–2002 using monthly data. Chadha, Sarno, and Valente (2004) consider the period 1979–2000 using quarterly data. RS (2003) use a sample of daily data from January 1985 to December 1999.[2] In our paper, instead, we use data until July 2007, when the current financial crisis begins.

We follow the approach proposed by RS (2003) and we use a VAR identified through heteroskedasticity. This approach can obtain reliable estimates from a sample of just a few years of data, since it relies on data at daily frequency. This is especially important given our focus on the past decade, which was not included in previous studies.

Three sets of results bring us to the conclusion that the interest rate response to stock prices in the past two decades has been

[2] Castelnuovo and Nisticò (2010) use quarterly data until 2007, but their sample starts in 1954. Therefore, their results are not directly comparable with ours.

substantially lower than what was found in previous state-of-the-art studies. First, when we apply the RS (2003) procedure to the EU and six inflation-targeting economies, we are able to identify a significant, albeit small, response only in Australia. This highlights the peculiarity of the U.S. case when compared with the international evidence. Second, we show that the sizable response estimated by RS (2003) for the period 1985–99 in the United States relies heavily on the effects of the stock market crash of 1987. When we reestimate the same model over the period 1988–99, the estimated response is substantially reduced, although still statistically significant. Third, when we estimate the same model over the so-called bubble period (2003–07), the monetary policy response becomes statistically insignificant.

The fact that the response of monetary policy to stock prices has declined over time in the United States is a robust feature of our model. Interestingly, Taylor (2007) identifies a substantial deviation of U.S. monetary policy from the prescription dictated by a Taylor rule responding to inflation and the output gap over the period 2003–07. A lower response to asset prices is consistent with the Taylor view that U.S. monetary policy over the period 2003–07 was more expansionary than in the previous period.

This paper has the following structure. In section 2 we present the RS methodology and we replicate the RS main result. In section 3 we provide a cross-country analysis and we extend the evidence for the United States. In section 4 we discuss some critical issues related to the identification-through-heteroskedasticity approach. Concluding remarks are contained in section 5.

2. The Model

VAR models are the most frequently used tool to measure the interactions between macroeconomic variables. As we are interested in interest rates and stock prices, the structure of the simplest VAR is the following:

$$A \begin{bmatrix} i_t \\ s_t \end{bmatrix} = C(L) \begin{bmatrix} i_{t-1} \\ s_{t-1} \end{bmatrix} + B \begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix},$$

where i_t is the three-month Treasury-bill interest rate, s_t is stock returns, A is a 2×2 matrix that describes contemporaneous relationships among the variables, $c(L)$ is a finite-order lag polynomial, and ε_t and η_t are structural disturbances. B is a 2×2 matrix in which non-zero off-diagonal elements allow some shocks to affect both endogenous variables.

We use the three-month Treasury-bill rate rather than the federal funds rate. While the federal funds rate is changed every six weeks or so, the three-month Treasury-bill rate adjusts daily, reflecting expectations of future variations in the federal funds rate, and is closely monitored by the central bank. RS (2003) and Bohl, Siklos, and Werner (2007) carefully motivate this choice. In this paper we follow the literature to facilitate results comparability.

The usual assumptions to achieve identification in this kind of model are to impose a triangular form to matrix A (Cholesky decomposition) and a diagonal structure to matrix B. In this way the model is exactly identified. However, a triangular matrix A implies that one of the two variables does not react contemporaneously to the other. This assumption, which is reasonable in other contexts, is clearly inappropriate in this case. In our application, each shock to one of the variables has an immediate effect on the other in the financial markets.

RS do not impose a triangular structure on matrix A, and they build an identification procedure relying on the heteroskedasticity that is present in the data and that is usually not considered in VAR studies. It is crucial to use daily data to exploit fully the heteroskedasticity present in the data. In fact, heteroskedasticity diminishes considerably in lower-frequency data.

Realizations of interest rates and stock returns can be seen as the intersection between two schedules. The first is the reaction function of asset prices to changes in the interest rate (supposed to be downward sloping because an increase in the interest rate lowers the discounted value of future dividends, i.e., the value of the asset). The second is the reaction of the interest rate to the evolution of the stock market. The objective of the procedure is to estimate the slope of this second schedule. Because of heteroskedasticity, endogeneity, and unobservability of a common shock z_t, introduced below, OLS estimates are biased. Thus, we look for a variable (an instrument) that shifts the stock market curve without affecting the monetary policy response. An increase in the variance of the stock market shock changes the covariance between stock returns and the interest rate, and this change plays the role of an instrument.

RS estimate the following VAR:

$$i_t = \beta s_t + \theta x_t + \gamma z_t + \varepsilon_t$$
$$s_t = \alpha i_t + \phi x_t + z_t + \eta_t,$$

where i_t is the three-month interest rate (T-bill), s_t is the daily return on a stock market index, x_t includes five lags of the two endogenous variables and some macroeconomic shocks (measured as monthly releases of some macro indicators and subtracting the value expected by market participants; see RS 2003), and z_t represents some unobserved shocks affecting both i_t and s_t. The common shock takes into account any macroeconomic shock not included in x_t or shifts in agents' risk preferences.[3] ε_t is a monetary policy shock, and η_t is a stock market shock.

The structure of the model is quite rich, but our objective is very simple. We want to estimate the coefficient β that measures the response of the interest rate to the stock market return.

The assumption on the correlation structure of the shocks is the following: the shocks ε_t and η_t and the unobserved shock z_t are supposed to be orthogonal, and at this stage all three can be heteroskedastic. Note that orthogonality of ε_t and η_t does not imply that disturbances are uncorrelated: In fact, the presence of z_t induces correlation.

The system cannot be estimated directly, because of the endogeneity problem discussed above and because z_t is an unobservable variable, but we can write it in reduced form:

$$\begin{bmatrix} i_t \\ s_t \end{bmatrix} = \Phi x_t + \begin{bmatrix} \upsilon_t^i \\ \upsilon_t^s \end{bmatrix},$$

where

$$\upsilon_t^i = \frac{1}{1-\alpha\beta}((\gamma+\beta)z_t + \beta\eta_t + \varepsilon_t)$$
$$\upsilon_t^s = \frac{1}{1-\alpha\beta}((1+\alpha\gamma)z_t + \eta_t + \alpha\varepsilon_t).$$

[3]The impact of z_t on s_t is normalized to one.

In the VAR literature, it is common practice to recover the estimates of structural-form parameters from reduced-form residuals. Given the structure of correlations specified above, the covariance matrix of reduced-form residuals is as follows:

$$\Omega = \frac{1}{(1-\alpha\beta)^2} \times \begin{bmatrix} (\beta+\gamma)^2\sigma_z^2 + \beta^2\sigma_\eta^2 + \sigma_\varepsilon^2 (1+\alpha\gamma)(\beta+\gamma)\sigma_z^2 + \beta\sigma_\eta^2 \\ + \alpha\sigma_\varepsilon^2 (1+\alpha\gamma)^2\sigma_z^2 + \sigma_\eta^2 + \alpha^2\sigma_\varepsilon^2 \end{bmatrix}.$$

By estimating the model in reduced form, we obtain a consistent estimate for the covariance matrix of reduced-form residuals. Unfortunately, the covariance matrix provides only three moments, which are not enough to achieve identification. The maximum number of parameters that can be identified is three, but in matrix Ω we have six unknowns: α, β, γ, σ_z^2, σ_η^2, σ_ε^2. Hence, we do not have enough restrictions to recover the structural-form parameters.

Still, heteroskedasticity can help in our task if we can identify different regimes for the covariance matrix of the reduced-form residuals. The additional regimes provide new restrictions and may enable us to identify the parameters of the structural form. Unfortunately, for each new regime indexed by the subscript i, we add three new equations but also three new unknowns: $\sigma_{i,z}^2$, $\sigma_{i,\eta}^2$, $\sigma_{i,\varepsilon}^2$. Nevertheless, if we assume that the monetary policy shock ε is homoskedastic (thus σ_ε^2 is constant across regimes), we add three equations and only two unknowns for each regime. With three regimes we have nine equations and ten unknowns $(\alpha, \beta, \gamma, \sigma_\varepsilon^2, \sigma_{1,z}^2, \sigma_{1,\eta}^2, \sigma_{2,z}^2, \sigma_{2,\eta}^2, \sigma_{3,z}^2, \sigma_{3,\eta}^2)$, but this is enough to achieve partial identification, and in particular we can estimate the parameter β.

The assumption that σ_ε^2 is constant is not very restrictive because it does not imply that i_t is homoskedastic. In fact, the variance of the interest rate is also composed of $\sigma_{i,z}^2$ and $\sigma_{i,\eta}^2$, which change through time.[4] The other essential assumption to achieve identification is

[4]The assumption of homoskedastic monetary shocks implies that we have to limit our attention to stable monetary regimes. For example, the RS procedure is not accurate (and delivers unstable results) for the case of Japan, where the central bank was forced to change its strategy many times in the past twenty years to fight against the zero lower bound.

that the parameters α, β, and γ are constant across regimes. This is common practice in the VAR literature, also when heteroskedasticity is not considered.

RS (2003) show that with three regimes, one solution of the following quadratic equation is a consistent estimator for β:

$$a\beta^2 - b\beta + c = 0,$$

where $a = \Delta\Omega_{31,22}\Delta\Omega_{21,12} - \Delta\Omega_{21,22}\Delta\Omega_{31,12}$, $b = \Delta\Omega_{31,22}\Delta\Omega_{21,11} - \Delta\Omega_{21,22}\Delta\Omega_{31,11}$, and $c = \Delta\Omega_{31,12}\Delta\Omega_{21,11} - \Delta\Omega_{21,12}\Delta\Omega_{31,11}$.[5]

With four regimes, we have overidentifying restrictions that allow us to estimate β by GMM.

An advantage with this model is that many assumptions are testable. In fact, if the model is correctly specified, we should find the same results for β under any three regimes, since the parameter β is supposed to be constant in the sample period. If not, the parameters are unstable across regimes, the assumption of homoskedasticity for the monetary policy shock is not correct, or there are non-linearities that are not captured in the RS specification.

Thus far, we have proved that with at least three regimes we are able to consistently estimate the parameter β. To determine the regimes, we estimate the VAR in reduced form and take the residuals. The heteroskedasticity of the shocks allows us to identify four regimes: regime 1 where both shocks have low volatility, regime 2 where the interest rate shock has low volatility and the stock market shock has high volatility, regime 3 where both shocks have high volatility, and regime 4 where the interest rate shock has high volatility and the stock market shock has low volatility. We split the observations into the four regimes according to the following criterion: one observation is considered to have high variance if the thirty-day rolling variance of the residual is more than one standard deviation over the average of the series. RS admit that this approach is arbitrary, but at least two arguments can justify this choice:

(i) As shown in Rigobon (2003), the estimates are consistent even if the regimes are badly specified. The estimates are

[5] $\Delta\Omega_{31,22}$ is the (2,2) element of matrix $\Delta\Omega_{31}$. $\Delta\Omega_{31} = \Omega_3 - \Omega_1$.

not consistent only if the misspecification is so large that the system fails to meet the following order condition:[6]

$$\Omega_{11,i}\Omega_{12,j} - \Omega_{11,j}\Omega_{12,i} \neq 0$$

for regimes i and j with $i \neq j$.

This condition has an intuitive explanation. It fails when two covariance matrices are proportional, i.e., relative variances are constant across regimes. In this case, some moment conditions are not independent and heteroskedasticity cannot be helpful (for a proof of this result and more details, see Rigobon 2003).[7]

(ii) The same criterion is largely used in the literature to identify periods of excessive volatility in asset markets (Bordo and Jeanne 2002).

The last step is to compute the distributions of the estimated coefficients. The distributions are calculated by bootstrap by using the asymptotic distribution of the covariance matrices. We simulate 1,000 draws for each Ω_i. For each covariance matrix, we estimate β using different subsets of regimes. In the end, we obtain 1,000 estimates and we are able to compute the distributions.

In our first experiment, we replicate the results of RS (2003). The sample period is January 1985–December 1999 and the data are daily. Data on three-month interest rates (T-bill rate for the United States) and stock market indexes (S&P 500 for the United States) are taken from Datastream and Bloomberg.

RS (2003) include in the variable x_t some observable macroeconomic shocks measured as the difference between the released value and the expected value of five monthly macroeconomic indicators.

[6]In this case, the parameters are not even identified because this condition is the equivalent of the rank condition that is tested in the identification literature once the order condition (number of equations equal to number of unknowns to achieve just-identification) is satisfied.

[7]Rigobon (2003) proves that the estimates are consistent also when the windows of the heteroskedasticity are wrongly specified (but the number of regimes is correct) and when there are more regimes in the data than the ones assumed in the model. Of course, if the regimes are badly specified, the differences across regimes are lower and the power of the model is reduced.

Table 1. Estimates for the United States, 1985–99

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0183	0.0180	0.0185	0.0177	0.0191
Std. Dev. of Distribution	0.0457	0.0093	0.0055	0.0055	0.0058
Mass Below Zero	0.7%	2.3%	0.1%	0.2%	0.2%

Table 2. Results of RS (2003), the United States, 1985–99

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0210	0.0214	0.0210	0.0273	0.1402
Std. Dev. of Distribution	0.0052	0.0058	0.0052	0.314	3.8122
Mass Below Zero	0.0%	0.0%	0.0%	1.4%	1.4%

The role of these shocks in the model is negligible, and in fact we are even able to reproduce the results for the United States without them. In our specification, the variable x_t only consists of five lags of the two endogenous variables.

The results (shown in table 1) are consistent with the assumptions of the model. The estimates of the coefficient β are almost identical across regimes and are very similar to the ones found by RS (2003), which we present in table 2.

The reader may notice that the quality of estimates involving both regimes 3 and 4 in the RS specification is significantly lower than the others. In regimes 3 and 4, the residuals in the interest rate equation exhibit high volatility. Apparently, regimes with high stock market volatility are more useful for identifying the parameters. This fact is confirmed in all our estimates throughout the paper.

The main result of the RS paper is that by employing an appropriate identification procedure, the reaction of monetary policy to stock price movements is positive and significant. A point estimate of 0.018 means that a 10 percent rise in the S&P 500 index increases the three-month interest rate by 18 basis points. RS argue that this result is very plausible and corresponds to the impact of stock prices on aggregate demand, mainly through the wealth effect on consumption.

Table 3. Estimates for the European Union, 1999–2006

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	−0.0003	0.0005	0.0005	0.0062	0.0063
Std. Dev. of Distribution	0.0193	0.0006	0.0006	0.0070	0.0086
Mass Below Zero	18.9%	19.5%	18.4%	17.6%	18%

3. Results

In this section we present our results, showing that the monetary policy response to stock prices in normal times has been substantially lower than that found in previous studies. First, we present evidence for the EU and six inflation-targeting countries. Second, we extend the analysis for the United States.

3.1 International Evidence

We estimate our model with data for the EU and six inflation-targeting countries (Australia, Canada, New Zealand, Norway, Sweden, and the United Kingdom).

We present the results for the EU in table 3. The sample period is January 1999–April 2006:

The European Central Bank (ECB) reaction turns out to be not significant. The size of the sample is smaller than for the other countries, due to the introduction of the euro in 1999, but is sufficient to derive some evidence on European monetary policy. In 2002, ECB President Trichet said explicitly that "it is clearly not opportune to introduce asset prices in the central bank's reaction function," and our result confirms this view.[8] A topic of discussion is whether we should find some sign of indirect reaction. Trichet recognizes that recent changes in asset prices have influenced private spending more than past swings because of the more widespread stock ownership in a number of industrialized countries. He states, however, that this

[8]President Trichet's speech is available at www.banque-france.fr/gb/instit/telechar/discours/sp230402.pdf.

Table 4. Estimates for Inflation-Targeting Countries Based on All Regimes

	Australia 93–07	Canada 95–07	New Zealand 94–07	Norway 01–07	Sweden 93–07	United Kingdom 93–07
Mean	0.0058	0.0000	−0.0035	−0.0046	0.0009	0.0010
Std. Dev.	0.0023	0.0017	0.0040	0.0039	0.0006	0.0009
Mass Below Zero	0.3%	51.3%	80.9%	88.9%	9.8%	13.9%

impact is still low compared with the United States. And in fact our results do not show any sign of indirect reaction.

We now turn our attention to several inflation-targeting countries: Australia, Canada, New Zealand, Norway, Sweden, and the United Kingdom. Since our procedure requires stability in the monetary policy framework, we decided to run estimates over a period of stable inflation, according to the classification in Roger (2009). New Zealand was the first country to adopt inflation targeting in 1990 and achieved a stable inflation rate in 1993. Canada adopted inflation targeting in February 1991 and was able to stabilize inflation in early 1995. Our estimation period starts in February 1994 for New Zealand (due to data availability) and in January 1995 for Canada. According to Roger (2009), the remaining four countries adopted inflation targeting when inflation was already under control. Therefore, we will start our estimation period at the beginning of the inflation-targeting regime in January 1993 for Australia, Sweden, and the United Kingdom and in March 2001 for Norway.[9]

The only country that has reacted significantly to stock prices under the inflation-targeting regime is Australia (table 4). Notice, however, that the size of the response is rather low and much smaller than in the United States. The response is even smaller, and significant only at the margin, in Sweden and the United Kingdom. We do not detect any response in Canada, whereas we find a somewhat negative response in New Zealand and Norway. We have not found

[9]Very few observations enter the high-volatility regimes in Norway, Sweden, and the United Kingdom. In these cases, we lower the window to one-half of a standard deviation instead of one standard deviation.

interesting subsample dynamics, as is the case for the United States (cf. below). On average, our results square well with previous estimates by Bohl, Siklos, and Werner (2007) for Germany and Chadha, Sarno, and Valente (2004) for the United Kingdom. It seems that stock prices played a limited role in the monetary policy strategy in inflation-targeting open economies.

According to the RS (2003) interpretation, the estimates across countries should be proportional to the financial wealth effect. Slacalek (2009) estimates the magnitude of the financial wealth effect in several countries. He finds significant effects in the United States, Australia, Canada, Sweden, and, to some extent, also in the United Kingdom. The estimates for other European countries are much lower and statistically not significant, in keeping with Altissimo et al. (2005) and Paiella (2007). In some cases our international evidence on the monetary policy response to stock prices is consistent with the above results: this is the case for the United States, Australia (although our estimated response is quantitatively very small), and the EU. However, we do not detect a meaningful response in Canada, Sweden, and the United Kingdom. Therefore, our results do not offer material for a systematic analysis of the cross-country variation in the monetary policy reaction. Instead, our results strongly highlight the peculiarity of the U.S. case in the international context, given that it is only in the United States that we find a non-negligible response, at least until 2003 (cf. below).[10] Our evidence seems to justify the fact that the literature focuses to such an extent on the United States and that only a few contributions consider other countries.

3.2 More Results for the United States

We now turn our attention to the United States and we extend the RS analysis until the beginning of the recent financial crisis.

A first legitimate question is whether the stock market crash of 1987 drove the RS result to some extent. It may very well be that such an unusual event required a large policy response. To analyze

[10] Assenmacher-Wesche and Gerlach (2008) study how the financial structure influences the impact of monetary policy on asset prices. They conclude that it plays a limited role. Our cross-country analysis for the opposite relationship (the impact of asset prices on monetary policy) suggests a similar result.

Table 5. Estimates for the United States, 1988–99

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0085	0.0090	0.0083	0.0362	0.0412
Std. Dev. of Distribution	0.0024	0.0024	0.0024	0.0145	0.0216
Mass Below Zero	0%	0%	0%	0.2%	2%

Table 6. Estimates for the United States, 1988–2003

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0080	0.0078	0.0080	0.0310	0.0157
Std. Dev. of Distribution	0.0016	0.0017	0.0017	0.0227	0.0524
Mass Below Zero	0%	0%	0%	6.5%	31.2%

this issue, we start our estimation in January 1988 instead of January 1985. The results shown in table 5 are considerably different from the evidence provided above (cf. table 1).

The estimated coefficient is still positive and strongly significantly different from zero. However, its magnitude is reduced from 0.018 to 0.0085 in the estimate using four regimes. Such a lower response is confirmed when we extend the sample until February 2003 (table 6). Results are rather stable and indicate an interest rate response much lower than the one estimated by RS (2003) for the period 1985–99.

So far we have disregarded the period from March 2003 to July 2007 because this is the period when stock and real estate prices surged, the so-called bubble period that preceded the current deep crisis. Taylor (2007) argues that during this period U.S. monetary policy was more expansionary than in the recent past: "During the period from 2003 to 2006 the federal funds rate was well below what experience during the previous two decades of good macroeconomic performance would have predicted."

On the basis of this argument, we believe that the period 2003–07 deserves a separate investigation to see whether we observe variations in the response to stock prices.

Table 7. Estimates for the United States, 2003-07

	Regimes All	Regimes 1,2,3	Regimes 1,2,4	Regimes 1,3,4	Regimes 2,3,4
Mean of Distribution	0.0025	0.0018	0.0024	0.0162	0.0101
Std. Dev. of Distribution	0.0047	0.0056	0.0047	0.0641	0.0284
Mass Below Zero	28%	35%	28.7%	44.7%	30.2%

In fact, our analysis reveals results that are consistent with Taylor (2007). The size of the interest rate response is reduced even further over 2003–07 and, most importantly, becomes statistically insignificant (table 7).[11]

In his analysis, Taylor (2007) considers the policy response to inflation and the output gap. Our analysis identifies that at the same time there is also a lower response to asset prices, at least compared with the recent past. However, let us stress that our result does not necessarily imply that monetary policy was too expansionary over that period. As Taylor (2007) recognizes, at that time there were good reasons for a prolonged period of low interest rates—in particular, to avoid a deflationary spiral similar to that of Japan in the 1990s. More research is definitely needed on these issues.

All in all, our analysis shows that the policy response estimated using identification through heteroskedasticity declines over time and is not statistically different from zero during the "bubble period." Moreover, we show that sizable responses identified in previous studies rely heavily on the stock market crash of 1987. Therefore, our analysis offers sound evidence to suggest that the monetary policy response to stock prices in normal times has been substantially lower than that found in previous studies.

It is reassuring that some very recent papers find evidence supportive of our results. First, Ravn (2010) applies the RS procedure

[11]In the working paper version of this article (Furlanetto 2008) we show that the lower response to stock prices was not compensated by a higher response to a proxy for real estate prices given by the REIT index. In fact, the response to real estate prices becomes also statistically insignificant over the "bubble period." Interestingly, however, the response to real estate prices is increasing over time up to 2003, unlike for stock prices.

for the United States over the period 1998–2008, allowing for asymmetric responses to stock market increases and decreases. Consistent with our results, he finds that the interest rate response has declined over time. Interestingly, he shows that the response to stock market decreases is larger than the response to stock market increases. Second, Milani (2010) estimates a DSGE model with learning and meaningful asset price dynamics over the period 1960–2007 with monthly data. Interestingly, he reestimates the model for the post-1984 sample and identifies a significant reduction in the response of monetary policy to stock market fluctuations. Our evidence reinforces the Milani (2010) result, indicating that the response has declined even further when we consider the "bubble period" in isolation. Finally, Bjørnland and Leitemo (2009) estimate a monthly VAR identified through a combination of short-run and long-run restrictions. According to their estimates, a 10 percent increase in stock prices would lead to an impact increase in the interest rate of 40 basis points over the period 1983–2002. This value is somewhat larger than our estimate, although the comparison is only illustrative given that the sample is different. Moreover, given that stock market volatility (on an annualized basis) is higher at daily frequency, it is not surprising that we find a lower response using higher-frequency data. However, interestingly, Bjørnland and Leitemo (2009, online appendix) also find a sizable decline in the impact response when they use a shorter sample (starting in January 1987).

4. Identification through Heteroskedasticity: Discussion

Our paper questions the robustness and the stability of the main result in RS (2003), i.e., the positive and sizeable response of monetary policy to stock prices. However, we believe that our study confirms that the RS procedure is extremely powerful. Using identification through heteroskedasticity, it is possible to estimate the model in short samples (only a few years), whereas other econometric techniques need much larger samples. The use of daily data enables the RS approach to discover interesting subsample dynamics, as shown in the preceding sections. Using Taylor rules or other VAR, we could not detect the decline in reaction observed during the "bubble period."

Table 8. Estimates for the United States, 1988–2003, No Common Shock

	Regimes 1,2	Regimes 1,3	Regimes 1,4	Regimes All
Mean of Distribution	0.0083	−0.0031	0.0025	0.0005
Std. Dev. of Distribution	0.0018	0.0034	0.0021	0.0019
Mass Below Zero	0%	84.5%	10.6%	39%

A potential weakness in our approach is that the three-month interest rate, which is our measure of the monetary policy stance, can react for reasons that are unrelated to monetary policy. This is a relevant objection that can, however, be countered by the following considerations: The presence of the unobservable common shock z_t, in our opinion, takes into account, at least to some extent, the factors that induce variations in interest rates and stock prices that are unrelated to policy expectations (like shifts in investor risk preferences). Identification through heteroskedasticity relies heavily on the common shock and, in fact, when the common shock is excluded, the results worsen significantly. As an example, in table 8 we report the estimates for the United States over the period 1988–2003 excluding the common shock (to be compared with table 6).[12]

Moreover, RS (2003) use other measures for the interest rate variable and find that the results are confirmed. Furthermore, the fact that our estimates for the period 1985–99 are in the same ballpark as the ones in studies that use the policy rate is also reassuring (see Chadha, Sarno, and Valente 2004 and Bjørnland and Leitemo 2009).

Another potential weakness of our analysis is the use of daily data. One could argue that the response of monetary policy is extremely gradual and is not influenced by day-to-day movements in financial markets. On the other hand, the use of daily data is also the strength of our approach because it is at daily frequency that heteroskedasticity is more present and therefore more helpful in obtaining accurate estimates. Importantly, when we estimate our model for the United States with weekly and monthly data, we find similar results (table 9). As expected, identification through

[12] Only two regimes are enough to identify β in the absence of a common shock.

Table 9. Estimates for the United States, 1988–2003, Based on All Regimes

	Daily Data	Weekly Data	Monthly Data
Mean of Distribution	0.0080	0.0163	0.0146
Std. Dev. of Distribution	0.0016	0.0064	0.0108
Mass Below Zero	0%	0.2%	6.3%

heteroskedasticty achieves less-precise estimates using low-frequency data. However, estimates are still positive and significant.

A relevant question that is always present in the literature on monetary policy and asset prices is whether the identified response represents a direct or indirect reaction to asset prices. RS (2003) argue that the magnitude of their estimate is compatible with a central bank that is concerned only with the impact of the stock market on output and inflation. Fuhrer and Tootell (2008) and Milani (2010) provide strong supporting empirical evidence. According to their results, central banks respond to equity price returns not directly (the central bank is not targeting asset prices) but only indirectly, since variations in stock prices can have important implications for the macroeconomic outlook. The main result from our cross-country and cross-period analysis is that it is difficult to find even an indirect response in recent years. We believe that our results show that central banks have been extremely cautious in their response to asset price movements, in keeping with arguments in Bernanke and Gertler (2001).

5. Conclusion

In this paper, we show that previous estimates of the monetary policy response to stock market fluctuations in the United States are not confirmed when we apply the analysis to other countries and when the sample period for the United States is extended until the beginning of the recent financial crisis.

For the other countries, we find a significant, albeit small, response only in Australia. This highlights the peculiarity of the U.S. case in the international context.

For the United States, we find that the response declines over time and becomes statistically insignificant during the bubble period (2003–07). Moreover, we show that the inclusion (or not) of the stock market crash of 1987 in the sample period significantly affects the results.

As far as we know, this is the first paper that provides evidence of a significant decline in the response of monetary policy to stock prices. However, supportive evidence for our claim can be found in papers that are focused on broader issues (Bjørnland and Leitemo 2009, Milani 2010, and Ravn 2010).

We believe that our results need further investigation. An interesting extension of this paper would be the estimation of the interest rate response to the stock market in a model with time-varying parameters. We plan to work on this project in the future.

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