HAS MONETARY POLICY REACTED TO ASSET PRICE MOVEMENTS: EVIDENCE FROM THE UK

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Abstract: This paper examines the relationship between monetary policy and asset prices in the context of empirical policy rules. We begin our analysis by establishing the forecasting ability of house and stock price changes with respect to future aggregate demand. We then report estimates of monetary policy reaction functions for the United Kingdom over the period 1992-2003. We find that UK policymakers appear to take into account the effect of asset price inflation when setting interest rates with a higher weight being assigned to property market fluctuations. Asset inflation-augmented rules describe more accurately actual policy, and the results are robust to modelling the effect of the Bank of England independence.

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1. Introduction

In the last decade or so it has been widely recognised that asset prices play an important role in determining business cycles conditions. As Bernanke and Gertler (2001) emphasise, asset market boom and busts have been important factors behind macroeconomic volatility in both industrial and developing countries. Following the financial deregulation in the early 1980s and the increased capital market globalization, industrial economies have witnessed an upward trend in asset prices. Alongside this trend, stock land and property prices have undergone swings around typical business cycle frequencies ranging from three to ten years. For some countries such as Japan and the Scandinavian counties during the late 1980s and the early 1990s, these swings had disruptive effects on domestic financial systems and contributed to prolonged recessions. In the U.K. case of 1990-92, the financial system withstood the asset price collapse but the ensuing recession was anyway severe.

Many recent theoretical and empirical contributions on the transmission mechanism of monetary policy imply that equity and property prices play an important role via wealth effects and balance sheet effects. Mishkin (1999) offers an excellent review of all the related arguments. For example, a rise in stock prices decreases the perceived level of financial distress by households which leads to increased consumption spending. The balance sheet channel implies a positive relationship between the firms' ability to borrow and their net worth which in turn depends on asset valuations. This extra credit can be used to purchase goods and services and thus stimulates economic activity (Kiyotaki and Moore, 1997).

Since banks engage heavily in real estate lending, in which the value of the real estate acts as a collateral, swings in property prices lead to increased financial instability. The monetary and financial stability objectives are intertwined. Sinclair (2000) reminds us that the Central Bank by setting the interest rate controls an important link between the two forms of stability. A low degree of policy smoothness may cause an excess volatility in financial markets, especially when the policy is delayed. Borio and Lowe (2002) point out that while a low and stable inflation environment promotes financial stability, it also raises the likelihood that excess demand pressures will first show up in credit aggregates and asset prices rather than in consumer prices.

Goodhart and Hofmann (2001) find that stock price and house price increases raise future aggregate demand in many major economies. Monetary policy affects asset prices via changes in expected future dividends and/or changes in the discount rate, which consequently affect aggregate demand and future inflation through balance sheet and wealth effects. Some authors have suggested a pro-active view for monetary policy, claiming that macroeconomic performance can be improved by reacting to asset price bubbles, even if conventionally measured inflation appears to be under control (Cecchetti *et al.*, 2000).

Bernanke and Gertler (1999, 2001) show that such results depend crucially on assumptions about the structure of the bubble and argue that asset prices should only be used in inflation forecasts. In their view, an inflation-targeting Central Bank should consider asset prices movements only if they affect inflationary forces through wealth effects on aggregate demand. Goodhart and Hoffman (2000) estimate the potential benefit from including asset prices in a model that predicts inflation. By comparing the inflation forecasting properties of reduced form equations with and without asset prices they show that, particularly at the two-year horizon, house prices are useful indicators for consumer price inflation. Goodhart (1999) has recommended that, since asset prices represent the current price of claims on future as well as current consumption, policymakers should consider a broader price measure that includes housing and stock market indices[1].

The main objective of this paper is to provide some further insight into the interest rate setting behavior of the Bank of England (BoE), by estimating reaction functions that have been augmented to take into account the effect of asset price inflation. We depart from the traditional approach that considers the exchange rate as determinant of aggregate demand and interest rates, and focus on two other important asset prices, equity and house prices. The results indicate that UK monetary policy reacts to asset price fluctuations, especially when they originate from the housing market.

The remainder of the paper is structured as follows. The next section discusses the properties of the data. In Section 3 we provide some empirical evidence on the magnitude of the wealth effects using a small structural model of the UK economy. Section 4 contains the empirical results from augmented Taylor rules. Section 4.1 accounts for the effect of BoE independence on the inflation-policy rule parameter, and Section 4.2 compares the historical performance of the benchmark forward-looking rule versus the asset augmented rule. Section 5 concludes.

2. Data description

We employ monthly data on short-term interest rates, industrial production, Retail Price Index minus mortgage interest payments (RPIX), stock prices and house prices for the United Kingdom from October 1992 to January 2003. All data is obtained from Datastream, *OECD Historical Statistics* series. Our sample period commences with the establishment of an explicit inflation targeting regime on October 1992. An inflation target of 1-4 % was initially adopted and on June 1995 the

[1] The theoretical foundation of this argument lays in pioneering research on the theory of inflation

target was reset at 2.5% or less. In addition, on May 1997 the Bank of England was awarded operational independence in setting interest rates.

The (annualised) output gap, $\tilde{y}_t$, the difference between actual and potential output, is calculated via quadratic detrending of the industrial production series, as in Clarida, Gali, and Gertler (1998), and Nelson (2000). The 3-month Treasury Bill rate, R_t, is employed as a measure of the stance of the monetary policy[2]. The annual change in stock prices, π_t^{SP}, house prices, π_t^{HP}, and retail prices, π_t, is proxied by the 12th difference of the natural logarithm of the monthly FTSE All Shares stock index, the Halifax house price index, and the (seasonally adjusted) RPIX respectively. We used annual, rather than monthly, changes for retail prices, stock prices and house prices since year on year changes on these variables are much more relevant for monetary policy decisions (Goodhart, 2001).

[FIGURE 1]

As we see in Figure 1(a) stock prices have been far more volatile as compared to house prices. In the aftermath of the burst of the stock market bubble in 2000, the two series start diverging significantly, with house price inflation accelerating, while the stock market collapsed. The output gap in Figure 1(b) indicates a post-bubble weaker UK economy since it is generally declining after peaking in early 2001. Finally, the nominal interest rate is consistently above the strikingly stable (as compared to its 1970's 'rollercoaster' behaviour) inflation rate. Standard Phillips-Perron (PP) tests are employed in order to test for unit roots in our data. The results in Table 1 clearly indicate the rejection of the unit root null hypothesis at the usual levels of significance. Therefore we employ all the aforementioned variables at their levels.

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measurement by Alchian and Klein (1973).

[2] The actual interest rate used by the Bank of England as its instrument has varied over time, and has included Bank Rate (until September 1972), Minimum Lending Rate (1972-81), and the two-week Repo Rate (since 1996). The Treasury Bill rate has historically moved close with these instruments, and is available for the entire sample period.

[TABLE 1]

3. Asset Prices and the Real Economy

As Goodhart and Hofmann (2001) argue, from the early 1990's many countries adopted explicit inflation targeting regimes as a response to the instability of the money demand function, which made monetary growth rates an unreliable proxy of monetary policy and future inflationary pressures. Simplified inflation targeting models include an aggregate supply and an aggregate demand equation and are used to derive optimal monetary policy reaction functions (see e.g. Svensson, 1997). Following Rudebusch and Svensson (1998), we use a model of the UK economy consisting of a aggregate supply, or Phillips curve, describing the dynamics of inflation and an aggregate demand, or IS curve, characterising the behaviour of the output gap.

Due to the inconsistencies between purely forward-looking models and actual inflation and output data, many researchers (see e.g. the references provided in Clarida, Galí and Gertler, 1999) suggest the employment of "hybrid" Phillips- and IS-curves, which include both backward- and forward-looking elements. We specify a hybrid empirical Phillips curve where current inflation depends upon past and expected future inflation and on past demand pressures:

$$\pi_t = \mu_0 + \mu_1 \pi_{t-i} + \mu_2 E_t[\pi_{t+n}] + \mu_3 \tilde{y}_{t-m} + \eta_t \qquad (1)$$

The backward-looking element in Equation 1 reflects inertia in inflation that is justified not only empirically, but also theoretically on the assumption that a fixed proportion of firms has backward-looking price setting behavior (see Galí and Gertler, 1999). The forward-looking element derives from the rational expectations staggered-contracting models of Taylor (1980), and Calvo (1983). The GMM estimates[3] of the

[3] GMM estimation with MA(12) autocorrelation correction has been used. The instrument set includes six lags of inflation, agricultural commodity prices, and the output gap. The *J*-statistic indicates that the overidentifying restrictions are not rejected.

hybrid Phillips curve in 1.1 over the period 1992:10-2003:01 imply that the backward-looking component is stronger in UK data, as $\mu_1 > \mu_2$. The estimated magnitude of μ_1 is close to the value of 0.8 which has been used as central value by Batini and Haldane (1999) in simulations for the UK. The output gap coefficient, μ_3, is positive and highly significant indicating that current demand pressures feed into higher future inflation.

$$\underset{}{\pi_t} = \underset{(.001)}{-0.001} + \underset{(.04)}{0.81}\pi_{t-3} + \underset{(.05)}{0.15}\pi_{t+3} + \underset{(.005)}{0.09}\tilde{y}_{t-12} + \hat{\eta}_t \tag{1.1}$$

SE = 0.003, J-stat = 14.25 [0.38]

The demand-side of the economy is modelled as a hybrid IS, that is consistent with dynamic optimising behaviour by the agents (micro-foundations) and also allows for some persistence in output. Thus, equation 2 allows the output gap to be a function of past and expected future output, lagged real interest rate, and lagged values of a vector, **X**, of additional explanatory variables.

$$\tilde{y}_t = \lambda_0 + \lambda_1 \tilde{y}_{t-i} + \lambda_2 E_t[\tilde{y}_{t+n}] + \lambda_3 \bar{i}_{t-s} + \mathbf{\Omega}'\mathbf{X}_{t-k} + v_t \tag{2}$$

where, $\bar{i}_t = (\frac{1}{12})\sum_{i=0}^{11}(R_{t-i} - \pi_{t-i})$ is the twelve-month average *ex post* real interest rate. As we already pointed out there are many channels via which, changes in equity prices and house prices affect consumer wealth. Direct effects include the change in consumption plans as a response to swings in asset prices (Modigliani, 1971), while indirect effects operate mainly via households and firms' balance sheets. There is a growing consensus that, apart from the conventional explanatory variables, output is also determined by changes in consumption and investment demand induced by changes in the level of asset prices. Therefore, the aggregate demand function given by Equation 2 is estimated with GMM including house price inflation (Equation 2.1) and stock price inflation (Equation 2.2).

$$\tilde{y}_t = 0.01 + 0.03\tilde{y}_{t-3} + 0.91\tilde{y}_{t+3} - 0.30\overline{i}_{t-12} + 0.08\pi^{HP}_{t-12} + \hat{v}_{1t}$$

$$\begin{array}{lllll} (.002) & (.03) & (.03) & (.05) & (.01) \end{array} \qquad (2.1)$$

SE = 0.008, J-stat = 18.91 [0.48]

The results indicate that, in contrast to the aggregate supply, aggregate demand is more affected by its forward-looking component, since λ_2 is positive and significant while λ_1 turns out to be insignificant. The one-year lagged real interest rate is negative and strongly significant, as expected. Finally, the coefficient of lagged house price changes shows that a 10 % increase in house prices boosts current aggregate demand by a factor of 0.8 %, indicating significant wealth effects[4].

$$\tilde{y}_t = 0.01 + 0.007\tilde{y}_{t-3} + 0.78\tilde{y}_{t+3} - 0.55\overline{i}_{t-12} + 0.04\pi^{SP}_{t-12} + \hat{v}_{2t}$$

$$\begin{array}{lllll} (.002) & (.04) & (.03) & (.06) & (.01) \end{array} \qquad (2.2)$$

SE = 0.009, J-stat = 16.12 [0.64]

The estimated coefficients in 2.2 reveal similar patterns. Future output is strongly significant, real interest rate affects output negatively with a lag, and stock price inflation exerts a positive impact on aggregate demand. The magnitude of the wealth effect depends among other factors on the share of the respective asset in private sector wealth, with housing constituting the most significant asset in the households' portfolio in most countries. Indeed, we find that the magnitude of the wealth effect due to stock price increases is much smaller than the effect due to house price increases. The coefficient of stock price inflation is half the coefficient of stock price inflation (0.04 as compared to 0.08) a result that is in line with previous evidence by Goodhart and Hofmann (2001) for the UK. Using a panel of 14 developed economies Case, Quigley, and Shiller (2001) also find that the housing market appears to be more important than the stock market in affecting the real economy.

Having demonstrated that movements in asset prices affect future aggregate demand and consequently also future inflation, in the next section we shall focus on the key empirical question of this paper, that is, what has been the response of the monetary policy instrument to asset price inflation.

4. An Alternative Taylor Rule Specification

It is generally assumed that the monetary policy interest rate instrument responds with fixed, positive weights to deviations of inflation from a pre-specified target, and deviations of output from its potential level (Taylor, 1993). The past decade has seen a vast amount of empirical and theoretical work considering monetary policy reaction functions. Clarida, Gali and Gertler (1998) present econometric estimates of the Taylor rule coefficients for the United States. Nelson (2000) provides empirical evidence for the United Kingdom under alternative monetary policy regimes, over the period 1972-97, prior to the Bank of England (BoE) receiving operational independence. Focusing on the inflation-targeting period 1992-1997, Nelson's results suggest that a forward-looking looking Taylor rule outperforms a backward looking specification. His results contradict those of Kuttner and Posen (1999).

In contrast with Nelson, who used quarterly data, Kuttner and Posen employed monthly data over the period October 1992-December 1997 and found a coefficient of zero on inflation. Other important differences with the Nelson study include the use of the unemployment rate, instead of the real GDP for the construction of the output gap proxy, and the employment of the annualised month-to-month rather than annual inflation rate. As Nelson argues though, the use of annual inflation in the estimated rule is crucial for the results since the BoE's inflation target has always been expressed in terms of the annual (year-ended), rather than the monthly inflation rate.

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[4] Throughout the last 3 years of house price increases, the average annual house price inflation has been 12% in nominal terms, or about 10 % in real terms.

Following Clarida *et al.* (1998) we assume that the Central Bank has an operating target for the nominal short term interest rate that is based upon the state of the economy. In the benchmark model the state of the economy is characterised by the evolution of the output gap and expected inflation. This forward looking behaviour is consistent with a central bank that operates in the context of an inflation targeting regime (Kent and Lowe, 1997). In each period, the actual interest rate partially adjusts towards the target value. Svensson (1997) justifies the partial adjustment mechanism by including the change in interest rates in the Central Bank's loss function. Combining the target rule with the partial adjustment mechanism we obtain the empirical form of the monetary policy reaction function:

$$R_t = \left(1 - \sum_{i=1}^{l} \varphi_i \right)\left\{a + \beta(E_t[\pi_{t+n}] - \pi^*) + \gamma E_{t-1}[\tilde{y}_t]\right\} + \sum_{i=1}^{l} \varphi_i R_{t-i} + u_t \qquad (3)$$

where $\sum_{i=1}^{l} \varphi_i \in [0,1]$ measuring the degree of interest rate smoothing, π^* is the inflation target (2.5 %), and $\alpha = r^* - \beta\pi^*$, with r^* denoting the long-run equilibrium nominal interest rate. Due to the fact that monetary policymakers cannot observe $\tilde{y}_t$ when setting R_t, we replace the actual value of the output gap with its expected level, $E_{t-1}[\tilde{y}_t]$; see McCallum and Nelson, 1999, and Orphanides *et al.*, 2000 for a further discussion of the uncertainties faced by the policymaker with respect to output. The error term, u_t, represents a white noise monetary policy shock. We consider an inflation forecast horizon of one year, therefore we set n equal to 12 in our monthly sample.

In order to estimate the model, unknown expected future variables are replaced with their ex post realised values. This leads us to Equation 4[5]:

[5] The disturbance term in Equation 2 is a linear combination of the inflation and output gap forecast errors and the exogenous monetary policy shock u_t, $\omega_t = -(1 - \sum_{i=1}^{l} \varphi_i)\left\{\beta\left(\pi_{t+n} - E_t[\pi_{t+n}]\right) + \gamma\left(\tilde{y}_t - E_{t-1}[\tilde{y}_t]\right)\right\} + u_t$

$$R_t = \left(1 - \sum_{i=1}^{l} \varphi_i\right)\left\{a + \beta(\pi_{t+n} - \pi^*) + \gamma \tilde{y}_t\right\} + \sum_{i=1}^{l} \varphi_i R_{t-i} + \omega_t \tag{4}$$

The set of orthogonality conditions implied by Equation (4) is:

$$E_t\left[R_t - \left(1 - \sum_{i=1}^{l} \varphi_i\right)\left\{a + \beta(\pi_{t+n} - \pi^*) + \gamma \tilde{y}_t\right\} + \sum_{i=1}^{l} \varphi_i R_{t-i} \middle| Z_t\right] = 0 \tag{5}$$

where Z_t represents all the variables in the Central Bank's information set available at time *t* when the interest rate is chosen. Z_t is a vector of variables that are orthogonal to ω_t. These instruments are lagged variables that help forecasting inflation and output, and contemporaneous variables that are uncorrelated with the exogenous monetary policy shock, u_t. The benchmark reaction function given by Equation (4) is estimated using the Generalised Method of Moments (GMM). The instruments employed in the estimation include a constant and six lags of the nominal short-term interest rate, inflation, output gap, and a world commodity price index (agricultural raw materials). Since the number of instruments is greater than the number of elements of the parameter vector [φ_i, α, β, γ], we test for the validity of the over-identifying restrictions using Hansen's *J*-statistic. As pointed out by Clarida *et al.* (1998), failure to reject orthogonality implies that the Central Bank considers lagged variables in its reaction function, only to the extent that they forecast future inflation or output.

The GMM estimation results in Table 2, column 2, indicate that the benchmark specification satisfies the dynamic stability criterion since the estimated inflation coefficient, β, is greater than one (1.02). If β was smaller than the stability threshold of one, then this would imply a positively sloped aggregated demand, with output decreasing in response to an inflation shock (Taylor, 1999). The output gap coefficient, γ, is positive and statistically significant at the 1 % level, although quite modest in magnitude (0.03). Its estimate implies that, holding expected inflation constant, one-percent increase in the level of output gap induces the BoE to raise interest rates by 3 basis points. This result is consistent with those reported by Martin and Milas (2001) who employed quarterly UK data. Therefore, during the inflation-targeting period that we consider, U.K. monetary policy has put more weight on price stability than output stabilisation. The sum of the interest rate smoothing parameters is close to one (0.92) indicating a high level of persistence in short term interest rates. Finally, the *J*-statistic indicates that the over-identifying restrictions of the benchmark model are not rejected.

[TABLE 2]

As pointed out in the previous section, asset prices contain important information about future aggregate demand and consequently inflation pressures. Also, there are theoretical arguments in favour of including asset price inflation in the reaction function of the Central Bank. Cecchetti *et al*. (2000) find that, on the basis of simulations, it would be desirable to include asset inflation in the Taylor rule. Therefore, we proceed by considering alternatives to our benchmark specification, by allowing asset prices to enter in the Taylor rule. The augmented reaction functions we consider are of the form:

$$R_t = \left(1 - \sum_{i=1}^{l} \varphi_i\right)\left\{a + \beta(E_t[\pi_{t+n}] - \pi^*) + \gamma E_{t-1}[\tilde{y}_t] + \mathbf{\Theta}'\mathbf{X}_t\right\} + \sum_{i=1}^{l} \varphi_i R_{t-i} + \varepsilon_t \qquad (6)$$

where $\mathbf{X}_t = [x_{1t} \quad \ldots \quad x_{jt}]'$, and $\mathbf{\Theta} = [\theta_1 \quad \ldots \quad \theta_j]'$ denote the vector of *j*-additional explanatory variables, and the relevant coefficient vector respectively. In the cases that we will examine, X_t contains contemporaneous house price and/or stock price inflation. We use contemporaneous, and not expected, asset price inflation due to the well known difficulties involved in forecasting asset price movements. Also, weak form efficiency implies that the current asset price reflects all past history, thus there is no need to incorporate lags.

First, we allow annual house price inflation to enter the reaction function. The results are presented in Table 2, column 3. The house inflation coefficient, θ_1, is positive and highly significant. Monetary policy tightens in response to increases in house prices: a one percent rise in house prices increases interest rates by 15 basis points. The response to expected inflation is stronger than in the benchmark case, with a smaller standard error. The estimated inflation coefficient is 1.6 close to the theoretical value of 1.5, as suggested by Taylor (1993), thus ensuring that real rates increase in response to inflationary pressures. Second, we add stock price changes in the benchmark model. The estimated coefficient, θ_2, (Table 2, column 3) is still positive and statistically significant but its value (0.06) is much smaller as compared to house price coefficient. When both asset returns are included (Table 2, column 4), the magnitude of the coefficients confirms that house prices enter more significantly the monetary policy reaction function, since $\theta_1 > \theta_2$. As in the benchmark specification, the *J*-statistic cannot reject the overidentifying restrictions.

4.1. Accounting for Independence

There is a wide consensus among academics and practitioners that central bank independence produces lower average inflation (Cukierman, 1992). Spiegel (1998) finds that the BoE independence on May 1997 had a significant negative impact on agents' inflationary expectations. In order to account for the change in the underlying regime and preferences we allow the expected inflation coefficient to be different post-independence. We therefore introduce a multiplicative dummy variable, D_t, in the reaction function, where $D_t = 0$ prior to independence and 1 onwards:

$$R_t = \left(1 - \sum_{i=1}^{l} \varphi_i\right) \left\{ a + (\beta + \mu D_t)(E_t[\pi_{t+n}] - \pi^*) + \gamma E_{t-1}[\tilde{y}_t] + \mathbf{\Theta}'\mathbf{X}_t \right\} + \sum_{i=1}^{l} \varphi_i R_{t-i} + v_t \qquad (7)$$

We would expect the dummy coefficient, μ, to be positive and significant indicating that the BoE becomes more inflation-averse, which leads to lower inflation expectations.

[TABLE 3]

The results in Table 3 confirm our predictions. In the benchmark model, μ is equal to 0.41 suggesting that post-independence the BoE reacts to one percent increase in expected inflation by raising interest rates by an additional 41 basis points. Even when we control for the effect of asset prices, the dummy remains positive and significant, reaffirming a higher degree of inflation aversion as a result of independence. We notice that the magnitude of the estimated asset inflation coefficients does not change. The monetary policy response to house prices is always stronger than the one with respect to stock prices. For instance, in the case of both assets entering the reaction function, θ_1 is equal to 0.13 and significant at the 1% level, while θ_2 is equal to 0.03 and significant at the 10%.The coefficient of house price inflation in particular is almost double the coefficient of stock price inflation. This supports the findings in the previous section where we reported a strong link between house price inflation and aggregate demand which induces policy makers to track more closely developments in the housing market.

4.2. Historical Performance

In this section we examine the historical performance of our estimated Taylor rules against the actual policy setting of the BoE. In Figure 2, we plot the implied target rates from the dummy augmented model versus the actual short term interest rate. As Clarida *et al.* (1998) point out, employment of the target rate as opposed to the fitted rate, that includes the lagged interest rate, allows for a better comparison of the alternative specifications. Figure 2(a) uses the target interest rate implied by the benchmark model, while Figure 2(b) uses the target rate implied by the asset price augmented policy rule[6].

[FIGURE 2]

It is clear that the benchmark target rate underperforms in capturing actual Central Bank behaviour. With the exception of two short periods at the beginning and the end of our sample, actual interest rates were consistently higher than the rule predicted value. Thus, the BoE was far tighter than the simple benchmark forward looking model would predict. When asset prices are allowed to be one of the state variables monitored by the Central Bank, the picture becomes clearer. As we notice in Figure 2(b), the target rate tracks the general trend of the actual interest rate for most of the period under investigation. Indeed, summary statistics presented in Table 4 indicate that when asset prices are not considered, the target interest rate is on average more than 1 % lower than the actual rate. Thus, the benchmark Taylor rule doesn't seem to explain well interest rate setting in the UK. The alternative specification produces a target rate which is much closer to the actual behaviour in terms of both mean and variability.

[TABLE 4]

During the period of strong stock market performance prior to the Asian financial crisis of 1997, actual monetary policy was looser than the model would predict; and then it was somewhat tighter during the first years of independence. One possible interpretation of these findings is that the BoE was trying to signal its commitment to keep inflation on track, even though some would argue that the falling stock market required lower interest rates. A notable divergence of the target from the actual occurs after 2001, when the large increases in house prices call for much tighter policy than the one followed.

[6] The target interest rate in Figure 1(b) has been constructed considering both house price and stock

5. Conclusions

This paper examines the empirical reaction of monetary policy to asset prices using forward-looking Taylor rule models of interest rates. The intuition for monetary policy to consider asset prices lays on the fact that consumption wealth effects and investment balance sheet effects may destabilise aggregate demand and inflation, the two main variables of interest for the CB. Changes in real property prices and stock prices have significant impact on households' consumption and firms' investment. Using UK data over the period October 1992 to January 2003 we show that movements in asset prices, especially house prices, have a significant positive impact on aggregate demand. Demand pressures feed into higher future inflation, thus there is scope for an inflation targeting CB to consider asset inflation in its forward-looking reaction function.

The main contribution of our paper is to find that policymakers in the UK appear to take into account both stock price inflation and house price inflation when setting interest rates, with the results suggesting that they are more concerned about developments in the property market. When the standard forward-looking Taylor rule is augmented by house prices and stock prices, the estimated coefficient of house price changes is always greater. The benchmark Taylor rule conditions short term interest rates upon expected inflation and the output gap and fails to provide an accurate characterisation of actual policy. When asset prices are included in the reaction function, the implied target rate describes the general trends of the actual interest rate much better. In addition, we model the effect of Central Bank independence on the policy preferences towards expected inflation. We find that the relationship between asset price inflation and interest rates is robust, and that inflation aversion increased post-independence.

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price inflation.

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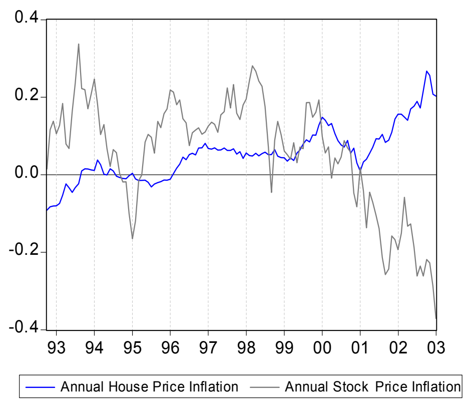
FIGURES

Figure 1: Annual House Price and Stock Price inflation, 1992:10-2003:01.

(a)



(b)

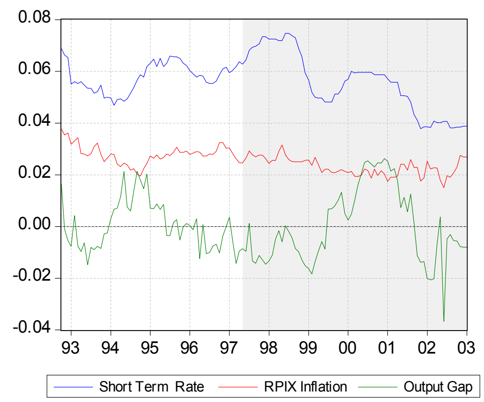
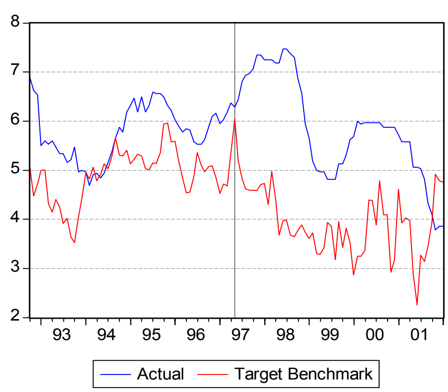
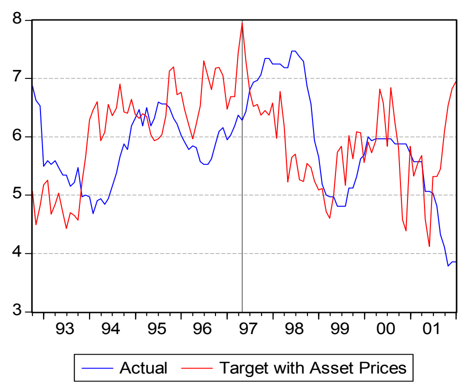


Figure 2: Actual and Target interest rate, 1992:10-2002:01.

(a)



(b)



TABLES

Table 1: Unit Root tests

Variables	Phillips Perron t-statistic
R	-3.620234 **
y^{gap}	-4.529641 ***
π	-4.001144 **
π^{SP}	-3.803221 **
π^{HP}	-3.748362 **

Note:

1. In order to correct for serial correlation the Phillips Perron (PP) test uses a non-parametric estimator of the variance-covariance matrix. A truncation lag of twelve was employed. An intercept and a linear trend term were included in the PP regressions.
2. The reported t-statistics test the null hypothesis that the variable contains a unit root. *, **, *** indicate rejection of the unit root - null at the 0.10, 0.05, and 0.01 significance level respectively.

Table 2: GMM Estimates of Forward Looking Taylor Rule, 1992:10-2003:1

	Benchmark Model	$X_t = [\pi_t^{HP}]'$	$X_t = [\pi_t^{SP}]'$	$X_t = [\pi_t^{HP}, \pi_t^{SP}]'$
a	2.59 *	4.00 ***	2.51 ***	4.56 ***
β	1.02 *	1.60 ***	1.01 ***	1.82 ***
γ	0.03 ***	0.04 ***	0.08 ***	0.02 ***
$\sum_{i=1}^{l} \varphi_i$	0.92 ***	0.91 ***	0.91 ***	0.83 ***
θ_1	-	0.15 ***	-	0.12 ***
θ_2	-	-	0.06 ***	0.07 ***
S.E. of Reg.	0.0023	0.0022	0.023	0.0030
J- Stat.	14.10	17.38	21.19	15.51

Table 3: GMM Estimates of Forward Looking Taylor Rule adjusted for the Effect of Bank of England Independence, 1992:10-2003:1

	Benchmark Model	$X_t = [\pi_t^{HP}]'$	$X_t = [\pi_t^{SP}]'$	$X_t = [\pi_t^{HP}, \pi_t^{SP}]'$
a	4.53 **	5.85 ***	3.27 **	5.62 ***
β	1.81 **	2.35 ***	1.30 ***	2.25 ***
γ	0.03 ***	0.04 ***	0.02 ***	0.03 ***
μ	0.41 *	0.47 **	0.61 ***	0.40 *
$\sum_{i=1}^{l} \varphi_i$	0.92 ***	0.88 ***	0.89 ***	0.87 ***
θ_1	-	0.12 ***	-	0.13 ***
θ_2	-	-	0.04 ***	0.03*
S.E. of Reg.	0.0022	0.0024	0.0022	0.0025
J- Stat.	13.30	15.76	23.18	15.78

Note:

1. Estimates are obtained by GMM estimation with correction for MA(12) autocorrelation. Two-stage least squares estimation is employed to obtain the initial estimates of the optimal weighting matrix.
2. In the benchmark model the instruments used are a constant and lags 1 to 6 of the nominal short term interest rate, inflation, output gap, and a world commodity price index (agricultural raw materials). In the models that include asset price inflation, lags 1 to 6 of the relevant asset price inflation variable are also included.
3. *J*-stat denotes the test statistic for overidentifying restrictions.
4. *, **, *** indicate level of significance of 10%, 5%, and 1% respectively.

Table 4: Descriptive Statistics of Actual and Taylor Rule Target Interest Rate, 1992:10-2002:1.

	Actual	Target Benchmark	Target with Asset Prices
Mean	5.83	4.42	5.92
Median	5.87	4.57	6.03
Maximum	7.46	6.04	8.24
Minimum	3.78	2.25	4.11
Std. Dev.	0.82	0.77	0.85

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