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Monetary Policy Rules, Asset Prices, and Exchange Rates

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We examine empirically whether asset prices and exchange rates may be admitted into a standard interest rate rule, using data for the United States, the United Kingdom, and Japan since 1979. Asset prices and exchange rates can be employed as information variables for a standard "Taylor-type" rule or as arguments in an augmented interest rate rule. Our empirical evidence, based on measures of the output gap proxied by marginal cost calculations, suggests that monetary policymakers may use asset prices and exchange rates not only as part of their information set for setting interest rates, but also to set interest rates to offset deviations of asset prices or exchange rates from their equilibrium levels. These results are open to several alternative interpretations. [JEL E40, E44, E52, E58]

It has become commonplace to characterize monetary policy as the minimization of inefficient economic fluctuations via the implementation of an interest rate rule (see Taylor, 1993, 1999, 2001). Such an interest rate rule relates the setting of short-term money market rates to the evolution of two key state variables, price inflation and the output gap. A considerable and influential research effort has been directed at establishing to what extent this kind of rule can explain the dynamics of policy rates in major industrialized countries (Clarida, Galí, and Gertler, 1998, 2000, and the references therein). The main virtue of this "Taylor" rule is its simplicity, but therein lies a vice, as there is controversy as to whether such a

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rule can be an adequate representation of a process as complex as that of monetary policy.

Accordingly, an important issue for central banks in recent years, for example, has been what weight to place on asset prices in setting interest rates (Greenspan, 1999; Batini and Nelson, 2000; Goodhart, 2001; Bordo and Jeanne, 2002). In periods of growing optimism, which may be associated with extended periods of economic expansion, asset prices may climb to unsustainable levels even if the path of likely returns (from, say, productivity gains or earnings growth) is unchanged (Shiller, 2000). The possibility of a sharp correction in asset prices may then leave the real economy unduly fragile, and this vulnerability may become an issue for monetary policymakers. Furthermore, macroeconomic imbalances may be exacerbated by the possibility that monetary policymakers may explicitly target asset prices. A useful dichotomy is whether asset price movements are a concern for central banks only because they contain information about future inflation—they are used as information variables—or whether they should be seen as variables to which central banks should react in addition to expected inflation and possibly the output gap. While some authors claim that including stock prices in the central bank's policy rule may be optimal (Cecchetti and others, 2000; Bordo and Jeanne, 2002) and that central banks react significantly to stock market movements by changing the short-term interest rate (Rigobon and Sack, 2003), other studies argue that central banks should not respond directly to asset prices (Bernanke and Gertler, 1999, 2001).[1]

Furthermore, in an open economy, the exchange rate is both a source of extraneous shocks and a mechanism for adjusting to fundamentals' shocks, and thus monetary policy choices are not independent of exchange rate dynamics. Some recent contributions argue that central banks may respond to the exchange rate, and this is particularly relevant in the case of small open economies (Obstfeld and Rogoff, 1995; Taylor, 2001). At the empirical level, some researchers provide evidence that exchange rates are statistically significant in interest rate rules depicting the reaction functions of several major central banks (Clarida, Galí, and Gertler, 1998).

In this paper we extend this analysis to consider jointly the role of asset prices and exchange rates in the interest rate rule to assess whether three major central banks—the U.S. Federal Reserve Bank, the Bank of England, and the Bank of Japan—have responded to asset prices and exchange rates during the last couple of decades or so, or whether they have used asset prices and exchange rates only as information variables that can help predict future inflation or output. In our study we allow asset prices and exchange rates to act both as information variables and as monetary policy targets, which enables us to investigate in an augmented interest rate rule whether and how central banks react to asset price and exchange rate disequilibria.

A subsidiary issue is the extent to which the output gap is correctly measured. In the empirical literature on central banks' reaction functions, the output

[1]For an extensive discussion of the issues related to the role of asset prices in designing macroeconomic policy, see the International Monetary Fund's *World Economic Outlook* (May 2000, Chapter 3, "Asset Prices and the Business Cycle").

gap has generally been measured by detrending output using deterministic trends (Taylor, 1993; Clarida, Galí, and Gertler, 1998, 2000) or smoothed trends representing the equilibrium level of output consistent with perfectly flexible prices (Kozicki, 1999, and the references therein). In practice, however, the efficient level of output may behave quite differently in response to real shocks such that potential output may not follow a smooth trend. As a result the output-gap measure that is relevant for monetary policy purposes may be different from the one captured by detrended output (Woodford, 2001). Recently, several contributions (Galí and Gertler, 1999; Galí, Gertler, and López-Salido, 2001) have shown that, under certain restrictions on technology and labor market structure within a local neighborhood of the steady state, real marginal costs are directly related to the output gap.

In this paper we tie together these different, albeit related, strands of the literature in that we investigate a very general monetary policy rule comprising inflation and the output gap as well as asset prices and exchange rates in a unified framework, while examining the role of alternative measures of real marginal costs and the output gap. Using quarterly data since 1979 for the United States, United Kingdom, and Japan we confirm the main results of the empirical literature on forward-looking monetary policy rules. However, we also show that our baseline estimation including asset prices, exchange rates, and an adjusted labor share as the appropriate proxy for the output gap yields significant parameter estimates for all the arguments in the interest rate rule of the United States and United Kingdom, and for all the arguments except asset prices in the case of Japan. These results are open to several alternative interpretations, which we discuss below.

I. Monetary Policy Rule Specification

In this section we briefly review the standard framework of analysis of forward-looking interest rate (or Taylor) rules, which we then generalize to an interest rate rule that explicitly allows for asset prices and exchange rates to act both as information variables and monetary policy targets. Finally, we also discuss a subsidiary issue relating to the calculation of the proxy of the output gap used in interest rate rules.

Forward-Looking Taylor Rules

Following Clarida, Galí, and Gertler (1998, 1999, 2000), assume that within each operating period the central bank has a target for the nominal short-term interest rate r_t^*, which is a function of the state of the economy represented by the gaps between expected inflation and output from their respective targets:

$$r_t^* = r^* + \beta\left[E\left(\pi_{t+k_\pi} \middle| \Omega_t\right) - \pi^*\right] + \gamma\left[E\left(x_{t+k_y} \middle| \Omega_t\right)\right], \tag{1}$$

where r^* is the desired nominal interest rate when both inflation and output are at their target levels; $E(\pi_{t+k_\pi}|\Omega_t)$ and $E(x_{t+k_y}|\Omega_t)$ denote the expectations of inflation at time $t + k_\pi$ and the output gap at time $t + k_y$; π^* is the level of inflation (implicitly or explicitly) targeted by the central bank; the output gap $x_{t+k_y} \equiv y_{t+k_y} - y^*_{t+k_y}$, where y is output and y^* is the efficient level of output.[2]

Consider the implied target for the ex ante real interest rate.[3] Central banks will pursue monetary policy aimed at stabilizing inflation if the parameter $\beta > 1$, while rules characterized by $\beta \leq 1$ will tend to be destabilizing. Clearly, the parameter γ is consistent with stabilizing output directly only if $\gamma > 0$ (Clarida, Galí, and Gertler, 2000).

Equation (1) may be, however, too restrictive to describe the behavior of short-term rates. As documented in earlier contributions (Goodfriend, 1987), central banks may operate smoothing changes in interest rates for different reasons (e.g., effects on capital markets, loss of reputation, private consensus building). To account for this behavior we can specify equation (1) assuming that the current short-term interest rate r_t adjusts to the target rate r_t^* according to a partial adjustment mechanism of the form $r_t = [1 - \rho(1)]r_t^* + \rho(L)\, r_{t-1} + \upsilon_t$, where $\rho(L)$ is an n – lag polynomial, L is the lag operator, and υ_t is a zero-mean exogenous interest rate shock. Allowing for smoothing in the above fashion in equation (1) results in an interest rate rule where in each period central banks act on the short-term interest rate, reducing the gap between the current target rate and its past value:

$$r_t = \alpha + \lambda\pi_{t+k_\pi} + \vartheta x_{t+k_y} + \rho(L) r_{t-1} + \varepsilon_t, \tag{2}$$

where we have eliminated the unobserved conditional expectations by rewriting equation (1) in terms of realized variables; $\alpha = [1 - \rho(1)](r^* - \beta\pi^*)$; $\lambda = [1 - \rho(1)]\beta$; $\vartheta = [1 - \rho(1)]\gamma$; and the error term $\varepsilon_t = \upsilon_t - [1 - \rho(1)]\{\beta[\pi_{t+k_\pi} - E(\pi_{t+k_\pi} | \Omega_t)] + \gamma[x_{t+k_y} - E(x_{t+k_y} | \Omega_t)]\}$ is the sum of a zero-mean exogenous shock and a linear combination of forecast errors that are orthogonal to the variables considered in the information set Ω_t. Given estimates of the parameters in the forward-looking Taylor rule given in equation (2), one can therefore recover the implied estimates of β and γ.

Augmenting Forward-Looking Taylor Rules with Asset Prices and Exchange Rates

Equation (2) does not explicitly consider the role of asset prices and exchange rates as monetary policy targets (or instruments). It is under debate whether and how asset prices and exchange rates should be taken into account in formulating monetary policy (Taylor, 2001; Clarida, 2001). While asset prices may be used as indicator variables, in the last decade or so several major central banks have started taking into consideration the apparent increase in financial instability, of which one important dimension is represented by the volatility of asset prices. In fact, as recorded in several contributions, asset booms and busts have been impor-

[2]This rule can also be obtained in a framework where the central bank faces a quadratic loss function over inflation and output (Svensson, 1997; Bernanke and Woodford, 1997, and the references therein).

[3]The ex ante real interest rate can be obtained as $rr_t^* = r_t - E(\pi_{t+k_\pi} | \Omega_t)$. Substituting this into equation (1), we get that $rr_t^* = rr^* + (\beta - 1)[E(\pi_{t+k_\pi} | \Omega_t) - \pi^*] + \gamma[E(x_{t+k_y} | \Omega_t)]$, where $rr^* = r^* - \pi^*$.

tant factors in macroeconomic fluctuations in both industrial and developing countries (Borio, Kennedy, and Prowse, 1994). Similar arguments apply to the role of exchange rates in central banks' reaction functions (Clarida, Galí, and Gertler, 1998; Taylor, 2001).

Thus, we consider a baseline equation—which may be termed "augmented Taylor rule"—of the form

$$r_t = a + \lambda\pi_{t+k_\pi} + \vartheta x_{t+k_y} + \chi e_{t-1} + \kappa s_{t-1} + \rho(L)r_{t-1} + \varepsilon_t, \tag{3}$$

where e_{t-1} and s_{t-1} denote the lagged real exchange rate and stock price; $a = [1 - \rho(1)](r^* - \beta\pi^* - \phi s^* - \delta e^*)$, where s^* and e^* denote the equilibrium levels of asset prices and exchange rates, respectively; $\lambda = [1 - \rho(1)]\beta$; $\vartheta = [1 - \rho(1)]\gamma$; $\chi = [1 - \rho(1)]\delta$ and $\kappa = [1 - \rho(1)]\phi$, where δ and ϕ are the parameters determining the central bank's response to exchange rates and asset prices disequilibria; and $\varepsilon_t = \upsilon_t - [1 - \rho(1)]\{\beta[\pi_{t+k_\pi} - E(\pi_{t+k_\pi}|\Omega_t)] + \gamma[x_{t+k_y} - E(x_{t+k_y}|\Omega_t)]\}$.

Equation (3) generalizes the standard forward-looking Taylor rule to account for the possibility that both asset prices and exchange rates enter the interest rate rule. Clearly, equation (3) nests not only the standard forward-looking Taylor rule estimated, for example, by Clarida, Galí, and Gertler (2000) but also its variants, which allow only for either exchange rates (Taylor, 2001) or asset prices (Bernanke and Gertler, 1999; Gilchrist and Leahy, 2002), respectively.

Let $\mathbf{z}_t$ denote a vector of instruments comprising the central bank's information set at the time the instrument rate r_t is chosen (i.e., $\mathbf{z}_t \in \Omega_t$). Equation (3) implies a set of orthogonality conditions because the elements of $\mathbf{z}_t$ include lagged variables that help forecast inflation, output as well as contemporaneous variables that are uncorrelated with the interest rate shock υ_t. More formally, since $E[\varepsilon_t|\mathbf{z}_t] = 0$:

$$E\left[r_1 - a - \lambda\pi_{t+k_\pi} - \vartheta x_{t+k_y} - \chi e_{t-1} - \kappa s_{t-1} - \rho(L)r_{t-1}|\mathbf{z}_t\right] = 0. \tag{4}$$

This set of orthogonality conditions is the basis of the estimation of the parameter vector $\{\beta, \gamma, \phi, \delta, \rho\}$, using the Generalized Method of Moments (GMM), with an optimal weighting matrix that accounts for possible serial correlation in ε_t. To the extent that the dimension of the vector $\mathbf{z}_t$ exceeds the number of parameters being estimated, equation (3) implies overidentifying restrictions that can be tested to assess the validity of the set of instruments used (Hansen, 1982). The results of this test can be interpreted as follows: under the null hypothesis the central bank adjusts the interest rate according to equation (3), with the expectations of future inflation and the output gap based on the relevant information available to policymakers at time t. Under general assumptions this implies that there exists a set of parameters $\{\tilde{\beta}, \tilde{\gamma}, \tilde{\phi}, \tilde{\delta}, \tilde{\rho}\}$ such that the residuals obtained from estimating (3) are orthogonal to the information set Ω_t. Under the alternative hypothesis the violation of the orthogonality conditions will lead to the statistical rejection of the model.

In line with some recent empirical literature, we assume that the real exchange rate follows a persistent but stationary process, implying that purchasing power parity (PPP) holds in the long run (Sarno and Taylor, 2002, and the references therein). The real exchange rate equilibrium level can then be captured by a constant included in the intercept term of the reaction function so that effectively the exchange rate variable entering (3) is the lagged deviation from PPP. Indeed, it is often assumed in international macroeconomic models that the instrument rate is affected by past values of exchange rates disequilibria (Obstfeld and Rogoff, 1995; Taylor, 2001; Hunt and Rebucci, 2003). Defining the real exchange rate as the domestic price of foreign currency, the resulting monetary policy rule acts to stabilize the exchange rate if $\delta > 0$. In other words, an appreciation of the real exchange rate will require a cut in short-term interest rates.

Even if equation (3) denotes a forward-looking Taylor rule, it has been assumed, as in Bernanke and Gertler (1999), that only past (once-lagged) stock price disequilibria affect the instrument rate, r_t. This assumption implies that central banks might intervene only when asset prices have been observed to deviate from equilibrium, rather than anticipating potential misalignments. Given our formulation, to stabilize asset prices we must have $\phi > 0$. This implies that whenever asset prices positively (negatively) deviate from their equilibrium level—for example, because of poor regulatory practice or “irrational exuberance” (Shiller, 2000)—central bankers will increase (decrease) the instrument rate to offset the anomalous price dynamics.

A Subsidiary Issue: The Role of Output-Gap Measures

In the previous subsections we have discussed how central banks may act by setting a target for the nominal short-term interest rate as a function of the state of the economy, which includes, inter alia, the output gap. In the empirical literature on central banks’ reaction functions, the output gap has generally been measured by detrending output using quadratic deterministic trends or smoothed trends representing the equilibrium level of output that would be obtained under perfectly flexible prices (Hodrick and Prescott, 1997; Baxter and King, 1999, and the references therein). In practice, the efficient level of output is likely to behave differently in response to different real shocks (e.g., shocks related to changes in technology, government expenditure plans, productivity, or consumers’ preferences). As a result the output-gap measure that is relevant for monetary policy purposes may be different from the one captured by detrended output.

A series of recent contributions (Galí and Gertler, 1999; Galí, Gertler, and López-Salido, 2001) emphasize this point in the context of a structural derivation of the Phillips curve. They show that under certain restrictions on technology and labor market structure (Rotemberg and Woodford, 1999) within a local neighborhood of the steady state, real marginal costs are proportionately related to the output gap. A measure of real marginal costs might therefore be considered as an appropriate proxy for the output gap (Woodford, 2001).

The most common measure of marginal cost used in the literature is represented by the cost of increasing labor input. If output is a differentiable function of the labor input and firms are wage takers, then the (real) marginal cost is equal to the (real) wage divided by the marginal product of labor. More formally, assume for simplicity that the aggregate production function is given by a Cobb-Douglas specification $Y_t = A_t N_t^{1-\alpha}$, where Y_t is aggregate output, N_t is employment at time t, A_t is a common technology factor, and $1 > \alpha > 0$. The real marginal cost (RMC) can be defined as

$$RMC_1 = \frac{W_t/P_t}{(1-\alpha)(Y_t/N_1)}, \tag{5}$$

where W_t is the nominal wage, and P_t is the aggregate price level. Real marginal costs move over time according to observed fluctuations in economic activity and employment. Indeed, there is a large literature supporting the view that marginal costs increase more than prices during expansions, inducing real marginal costs to increase. This is motivated on the basis of two arguments. First, holding constant other determinants of labor supply, real wages must rise during expansions to induce more people to work. Second, if one makes the standard assumption that the production function is concave, for a given level of technology the marginal product of labor is a decreasing function of labor input (Rotemberg and Woodford, 1999). This behavior is therefore consistent with the procyclical dynamics of real marginal costs.

However, variations in real unit labor costs have often been found to be negatively (instead of positively) correlated with detrended GDP (Rotemberg and Woodford, 1999; Woodford, 2001; Galí, Gertler, and López-Salido, 2001; Sbordone, 2002). To shed light on this countercyclical behavior, it is instructive to decompose movements in real unit labor costs in order to isolate the factors driving this variable. Following Galí, Gertler, and López-Salido (2001), real marginal costs are equal to the measure of distance of output from its efficient level:

$$rmc_1 = \log \mu_t^w + [(c_t - \varphi n_1) - (y_t - n_t)], \tag{6}$$

where c_t is consumption, μ_t^w is the gross wage markup, and lowercase variables are defined in logarithms. The expression in brackets $[(c_t - \varphi n_t) - (y_t - n_t)]$ is the (log-linearized) measure of distance of output from its efficiency level, or, using the Galí, Gertler, and López-Salido (2001) terminology, the "inefficiency wedge." In equation (6), $(c_t - \varphi n_t)$ is the marginal cost of labor supply, and φ is the inverse of the elasticity of labor supply. If we assume for simplicity an economy with just consumption goods ($c_t = y_t$), then we can derive an expression that explicitly links the "inefficiency wedge" to the output gap:

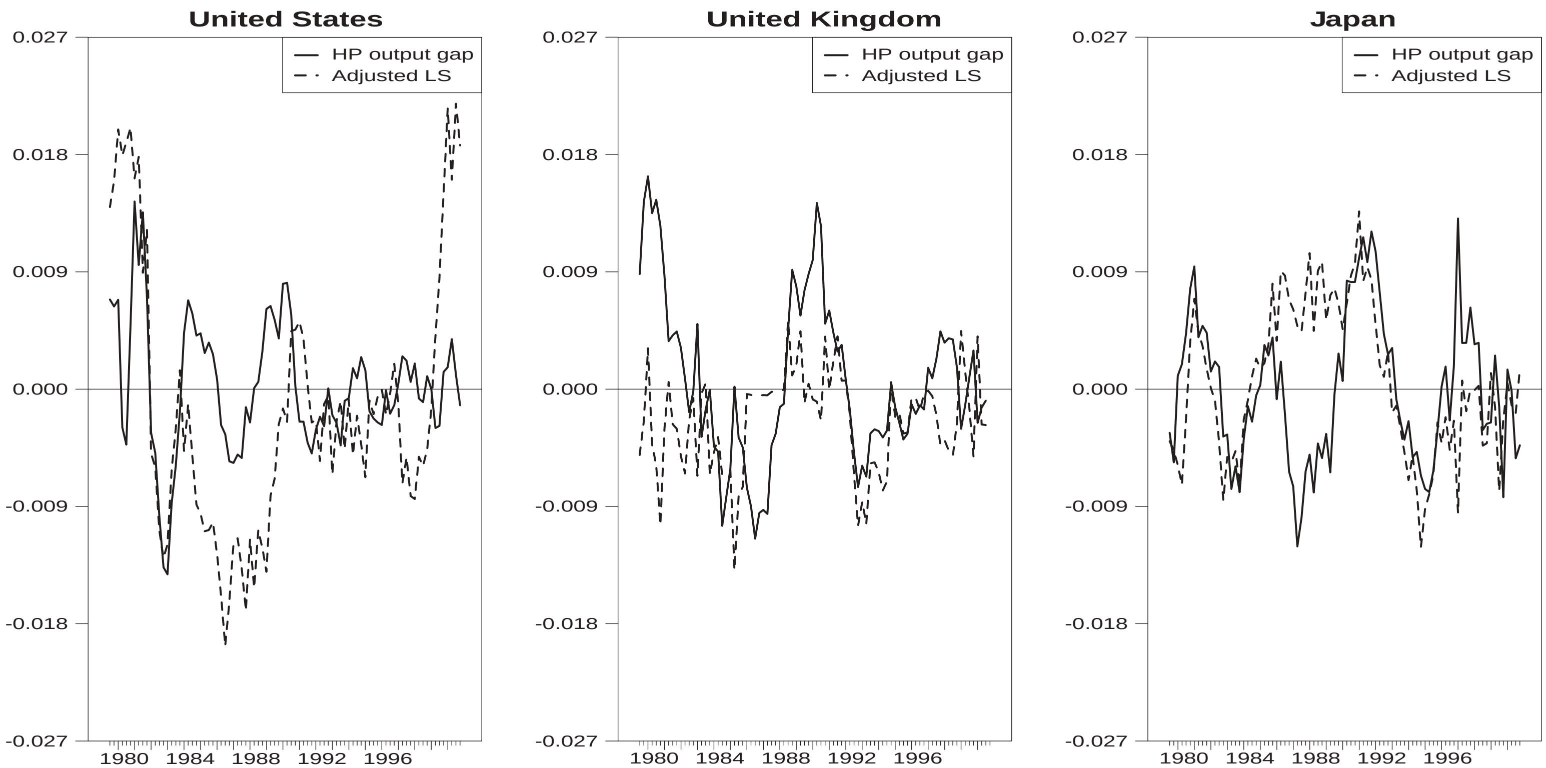
$$[(c_t - \varphi n_1) - (y_t - n_t)] = -\Theta + \omega\left(y_t - y_t^*\right), \tag{7}$$

where Θ is an indicator of steady state distortions as a result of the existence of market power in labor and goods markets, and y_t^* is the level of output that would be obtained under perfectly flexible prices and wages.

The countercyclical behavior of real marginal costs can be explained by using equations (6)–(7) to obtain

$$rmc_t = \log \mu_t^\omega - \Theta + \omega\left(y_t - y_t^*\right). \tag{8}$$

Figure 1. Output Gaps and Labor Shares



Equation (8) tells us that the presence of frictions in the labor market may drive the dynamics of real marginal costs. In fact, as documented in Woodford (2001) and Galí, Gertler, and López-Salido (2001) and as suggested by equation (8), while the "inefficiency wedge" moves closely with the business cycle over time, the wage markup should behave in a countercyclical fashion, suggesting the likelihood of temporary wage rigidities. By removing this markup component from real marginal costs, it should be possible to get a representative measure of the output gap, which might be used as an indicator of stabilization.

Figure 1 shows our time series for real unit labor costs after adjusting for wage markup and detrended output gap for the United States, the United Kingdom, and Japan. The calculation of the adjusted labor costs follows exactly our description of the markup adjustment given above, assuming, as is standard in this literature, the labor supply elasticity $\varphi = 1$ (Galí, Gertler, and López-Salido, 2001, p. 1263). The graphs in Figure 1 suggest that the prediction of the simple decomposition outlined above is confirmed, as evidenced by the fact that the adjusted real marginal cost measure appears to be positively correlated with the business cycle.

We now turn to a description of our data set.

II. Data

Our data set comprises quarterly time series spanning from September 1979 to December 2000 for the United States, the United Kingdom, and Japan.[4] We use as instrument interest rates the federal funds rate for the United States, the base rate for the Bank of England, and the call money rate for the Bank of Japan. The baseline inflation measure is, for all three countries examined, the (annualized) rate of change in the consumer price index (CPI), although we also report results using the GDP deflator. These time series are obtained from the International Monetary Fund's International Financial Statistics database.

The baseline output-gap measure is the adjusted labor share calculated as in Galí, Gertler, and López-Salido (2001, p. 1263). The data on consumption per household, employment per household, real wage, and labor productivity have been calculated using data from the *OECD Quarterly National Accounts* and *OECD Quarterly Labor Force Statistics*.[5] Results are also reported for alternative measures of the output gap, taken as the quadratically detrended and Hodrick-Prescott (HP) detrended GDP, respectively.

The measure of stock price disequilibrium in the baseline estimation is the dividend-price ratio calculated using *Datastream* composite stock price indices for each country examined.[6] The measure of real exchange rate is the log-real

[4]We decided to use this sample period because, as documented in previous contributions, there is evidence of structural instability in monetary policy reaction functions when using pre-1979 data (Clarida, Galí, and Gertler, 1998).

[5]For further details, see the relevant discussion in Galí, Gertler, and López-Salido (2001), which we followed closely in the construction of our time series.

[6]A robustness exercise was carried out using measures based on the S&P 500 index for the United States, the FTSE100 index for the United Kingdom, and the NIKKEI 225 index for Japan (see "Robustness" in Section III).

effective exchange rate, obtained for each country from the IMF's International Financial Statistics database.

The instrument set includes lags of the instrument rate, inflation, output gap, as well as the same number of lags of a world commodity price index and, except for Japan, the spreads between the respective long-term bond rate and the short-term bill rate. These time series are also obtained from the IMF's International Financial Statistics database.

III. Empirical Results

In this section we report our main empirical results from estimating both standard forward-looking Taylor rules as well as augmented forward-looking Taylor rules, which also allow for exchange rates and asset prices to enter the central bank's reaction function. We then report a battery of robustness checks and provide a discussion of the alternative potential explanations of our findings.

Forward-Looking Taylor Rules

Table 1 reports GMM estimates of the interest rate rule parameters $\{\tilde{\beta}, \tilde{\gamma}, \tilde{\rho}\}$ in a standard forward-looking Taylor rule for the United States, the United Kingdom, and Japan. We begin with an interest rate rule—equation (2)—where only expected inflation and expected output gap are considered as explanatory variables, also experimenting with various measures of these two arguments. Following previous research (Clarida, Galí, and Gertler, 1998, 2000), we consider a specification of the interest rate rule with two lags of the interest rate for the United States and only one lag for the United Kingdom and Japan. We allow for a maximum of four lags of the variables in the instrument set. The first four specifications we estimate consider two measures of inflation, namely the CPI and the GDP deflator, and two conventional measures of the output gap, namely quadratically and HP-detrended GDP.

The estimation results reported in Table 1 (specifications 1 to 4) yield parameter values that are consistent with the results recorded in the literature (Clarida, Galí, and Gertler, 1998, 2000). In particular, for the United States, for each specification considered, the estimate of β is always correctly signed, strongly statistically significant, and greater than unity, while the estimate of γ is not statistically different from zero at conventional levels of significance (see Panel A of Table 1). In turn, this implies that the Federal Reserve has responded only to deviations of expected inflation from the implicit inflation target, not to the expected output gap, over the sample examined. For the United Kingdom (see Panel B of Table 1), the estimate of β is also correctly signed and strongly statistically significant for each of specifications 1 to 4; the estimated β is only slightly greater than unity and generally lower than for the U.S. reaction function. Estimates of γ obtained using the HP-detrended GDP series are not statistically significant at conventional significance levels (Kozicki, 1999). For Japan (see Panel C of Table 1), the estimate of β is again correctly signed and significant, and its size is somewhere in between the estimates for the United Kingdom and the ones of the U.S. reaction function. Estimates of γ are statistically

Table 1. Standard Forward-Looking Taylor Rule: The Role of the Output Gap

Panel A: United States	$\tilde{\beta}$	$\tilde{\gamma}$	$\tilde{\rho}$	J-test
1. $\pi_t = \pi_t^{CPI}$, $x_t = x_t^{Q}$	1.688	0.021	0.827	0.253
	(0.591)	(0.328)	(0.267)	
2. $\pi_t = \pi_t^{CPI}$, $x_t = x_t^{HP}$	1.764	0.624	0.816	0.192
	(0.456)	(1.887)	(0.285)	
3. $\pi_t = \pi_t^{GDP}$, $x_t = x_t^{Q}$	1.720	–0.021	0.746	0.394
	(0.426)	(0.211)	(0.257)	
4. $\pi_t = \pi_t^{GDP}$, $x_t = x_t^{HP}$	1.582	–1.169	0.663	0.998
	(0.324)	(2.943)	(0.215)	
5. $\pi_t = \pi_t^{CPI}$, $x_t = LS_t$	1.925	–0.855	0.851	0.085
	(0.285)	(0.332)	(0.08)	
6. $\pi_t = \pi_t^{GDP}$, $x_t = LS_t$	2.036	–0.723	–0.830	0.258
	(0.125)	(0.143)	(0.05)	
7. $\pi_t = \pi_t^{CPI}$, $x_t = LS_t$	2.205	0.869	0.852	0.471
	(0.267)	(0.387)	(0.126)	
8. $\pi_t = \pi_t^{GDP}$, $x_t = LS_t$	2.392	1.073	0.844	0.714
	(0.389)	(0.492)	(0.228)	

Panel B: United Kingdom	$\tilde{\beta}$	$\tilde{\gamma}$	$\tilde{\rho}$	J-test
1. $\pi_t = \pi_t^{CPI}$, $x_t = x_t^{Q}$	1.039	0.241	0.786	0.067
	(0.203)	(0.185)	(0.038)	
2. $\pi_t = \pi_t^{CPI}$, $x_t = x_t^{HP}$	0.984	0.232	0.906	0.412
	(0.155)	(1.653)	(0.017)	
3. $\pi_t = \pi_t^{GDP}$, $x_t = x_t^{Q}$	1.054	0.315	0.738	0.079
	(0.228)	(0.208)	(0.097)	
4. $\pi_t = \pi_t^{GDP}$, $x_t = x_t^{HP}$	1.019	0.360	0.814	0.354
	(0.253)	(1.518)	(0.071)	
5. $\pi_t = \pi_t^{CPI}$, $x_t = LS_t$	1.457	–0.586	0.790	0.117
	(0.459)	(0.299)	(0.026)	
6. $\pi_t = \pi_t^{GDP}$, $x_t = LS_t$	1.271	–0.613	0.846	0.163
	(0.124)	(0.281)	(0.024)	
7. $\pi_t = \pi_t^{CPI}$, $x_t = LS_t$	1.086	0.448	0.894	0.999
	(0.100)	(0.174)	(0.007)	
8. $\pi_t = \pi_t^{GDP}$, $x_t = LS_t$	1.208	0.596	0.866	0.999
	(0.080)	(0.213)	(0.012)	

Panel C: Japan	$\tilde{\beta}$	$\tilde{\gamma}$	$\tilde{\rho}$	J-test
1. $\pi_t = \pi_t^{CPI}$, $x_t = x_t^{Q}$	1.737	0.050	0.972	0.991
	(0.259)	(0.020)	(0.003)	
2. $\pi_t = \pi_t^{CPI}$, $x_t = x_t^{HP}$	1.504	0.025	0.970	0.986
	(0.182)	(0.008)	(0.003)	
3. $\pi_t = \pi_t^{GDP}$, $x_t = x_t^{Q}$	1.328	0.005	0.892	0.853
	(0.174)	(0.001)	(0.028)	

(continued)

Table 1. *(concluded)*

Panel C: Japan	$\tilde{\beta}$	$\tilde{\gamma}$	$\tilde{\rho}$	J-test
4. $\pi_t = \pi_t^{GDP}$, $x_t = x_t^{HP}$	1.102	0.014	0.963	0.993
	(0.082)	(0.005)	(0.003)	
5. $\pi_t = \pi_t^{CPI}$, $x_t = LS_t$	1.473	–0.023	0.903	0.998
	(0.549)	(0.006)	(0.035)	
6. $\pi_t = \pi_t^{GDP}$, $x_t = LS_t$	1.259	–0.009	0.850	0.769
	(0.196)	(0.003)	(0.029)	
7. $\pi_t = \pi_t^{CPI}$, $x_t = LS_t$	2.366	0.034	0.946	0.828
	(0.389)	(0.015)	(0.019)	
8. $\pi_t = \pi_t^{GDP}$, $x_t = LS_t$	1.854	0.018	0.877	0.921
	(0.186)	(0.008)	(0.037)	

Notes: The estimated parameters refer to equation (2). The interest rate targets used are the U.S. federal funds rate, the Bank of England base rate, and the call money rate for Japan, respectively. The sample period is September 1979 to December 2000. The instruments set includes a constant, lagged output gap, inflation, the interest rate, and the log-difference of a world commodity price index. Estimates are obtained by GMM with a correction for moving average autocorrelation obtained by two-step nonlinear two-stage least square (Hansen, 1982). The optimal weighting matrix is obtained from the first step of two-stage least square parameter estimates. π_t^{CPI} and π_t^{GDP} are inflation rates measured using the CPI and GDP deflator, respectively. x_t^Q and x_t^{HP} are measures of the output gap (calculated over the period 1960–2000) obtained by means of quadratic-trend regression and Hodrick-Prescott filter (smoothing parameter set to 1,600), respectively. LS_t and LS_t are the labor share (calculated as in the text) and the labor share adjusted for the wage markup, respectively. J-test is the test for overidentifying restrictions (Hansen, 1982), which is distributed as χ^2 under the null. For this test only *p*-values are reported. $\tilde{\rho}$ denotes the sum of the autoregressive parameters associated with the lagged interest rate instrument. Standard errors are reported in parentheses.

significant but very small in magnitude. Also, for all of these specifications and for each country, the overidentifying restrictions cannot be rejected, and hence the Hansen test supports the validity of the instrument set used.

The last four specifications in Panels A to C of Table 1 show the results from estimating the forward-looking Taylor rule with arguments in expected inflation and output gap where the relevant measure of the output gap is given by a proxy of real marginal costs. Woodford (2001) argues that this output-gap measure may be more appropriate in this context. Specifications 5 and 6 are estimated using the real marginal cost as a proxy for the output gap, while specifications 7 and 8 consider the real marginal cost adjusted for the wage markup, following the procedure described in “A Subsidiary Issue: The Role of Output-Gap Measures” in Section I. The parameters estimates yield, also in this case, fairly clear results. Whenever we consider real unit labor cost into the interest rate rule, the estimated parameter γ is found to be statistically significant for each country examined, but with the wrong sign. This finding can be heuristically explained by the documented evidence of negative correlation between general measures of output gap and real marginal costs (Galí, Gertler, and López-Salido, 2001; Woodford, 2001).

Specifications 7 and 8 are estimated using real marginal costs adjusted for wage markup. As discussed in Section I's "Role of Output-Gap Measures," this correction circumvents the problem of the negative correlation between real marginal costs and the business cycle. Interestingly, using this measure of real marginal costs, the estimated parameter γ is, for each country, correctly signed (positive) and statistically significant, providing evidence in favor of the adjusted real marginal costs measure (Woodford, 2001).

An Augmented Forward-Looking Taylor Rule

Next, we estimate the interest rate rule parameters $\{\tilde{\beta}, \tilde{\gamma}, \tilde{\phi}, \tilde{\delta}, \tilde{\rho}\}$ for the United States and the United Kingdom in an augmented forward-looking Taylor rule of the form (3) using CPI inflation and the adjusted real marginal costs as proxies for inflation and output gap, respectively, and using the dividend-price ratio and the real effective exchange rate as proxies for stock market and exchange rate disequilibria, respectively. The target horizon for the forward-looking variables is assumed to be one quarter for both inflation and the output gap (i.e., $k_\pi = k_y = 1$ in equation (3)).

The results of the GMM estimation of equation (3), reported in Table 2, produce estimates of the parameters that are *all* statistically significant and exhibit the expected sign for both the United States and the United Kingdom. For Japan, the only insignificant parameter is the one associated with asset prices. More precisely, the estimated parameters ϕ and δ, associated with asset price and exchange rate disequilibria, respectively, are positive for each country examined, although

Table 2. Augmented Forward-Looking Taylor Rule: Baseline Estimates

	$\tilde{\beta}$	$\tilde{\gamma}$	$\tilde{\delta}$	$\tilde{\phi}$	$\tilde{\rho}$	J-test
United States	1.252	0.719	0.073	0.015	0.767	0.935
	(0.195)	(0.229)	(0.009)	(0.005)	(0.040)	
	[0.246]	[0.192]	[0.221]	[0.188]	[0.147]	
United Kingdom	1.034	0.542	0.056	0.007	0.889	0.999
	(0.026)	(0.037)	(0.006)	(0.003)	(0.001)	
	[0.160]	[0.104]	[0.175]	[0.173]	[0.136]	
Japan	1.115	0.013	0.002	0.0002	0.895	0.745
	(0.440)	(0.003)	(0.0008)	(0.0005)	(0.034)	
	[0.122]	[0.386]	[0.160]	[0.179]	[0.332]	

Notes: See notes to Table 1. The estimated parameters refer to equation (3). The sample period is September 1979 to December 2000. The baseline specification was estimated using the CPI to measure inflation and the adjusted labor share to measure the output gap. The forward-looking horizon for inflation and the output gap is one quarter ($k_\pi = 1$, $k_x = 1$). The instruments set includes a constant plus four lags of output gap, inflation, the interest rate, log difference of a world commodity price index, real effective exchange rate, and dividend-price ratio. Values in brackets are *p*-values from executing the Hansen (1992) test statistic for the null hypothesis of parameter stability.

they are small in absolute value. Interestingly, δ is always statistically significant for each country, and ϕ is statistically significant for the United States and the United Kingdom but insignificantly different from zero for Japan.[7] Further the large *p*-values of the test for overidentifying restrictions suggest that the instrument set used is valid.

The results in Table 2 also confirm the presence of interest rate inertia (detected in all of our previous specifications and well documented in the relevant literature), as captured by the lagged values of interest rates in equation (3). These findings support the view that the Federal Reserve, the Bank of England, and the Bank of Japan smooth the adjustment of interest rates toward their target values (Clarida, Galí, and Gertler, 2000; Eijffinger and Huizinga, 1999).

Some recent contributions point out that monetary policy rules estimated over long sample periods may be affected by parameter instability (Oliner, Rudebusch, and Sichel, 1996; Rudebusch, 1998, 2001). To test for parameter instability, we carried out a Hansen (1992) test for the null hypothesis of parameter stability, applied individually to each of the estimated parameters reported in Table 2. The results, reported in square brackets in Table 2, show that for each country examined the null hypothesis of parameter stability cannot be rejected at conventional statistical significance levels.

Before discussing and further interpreting our empirical results, we report some robustness checks carried out on our baseline estimation results.

Robustness

To investigate the robustness of the results in Table 2, we reestimated equation (3) using alternative measures of inflation, asset price disequilibria, and alternative target horizons for each of the forward-looking variables. As discussed below, in each country and over different specifications, this robustness analysis confirms, at least in a qualitative sense, the findings of our baseline estimates reported in Table 2. Using a different measure of inflation based on the GDP deflator (see specification 1 of Table 3), the estimated parameters are not qualitatively different from the estimates recorded in Table 2.

In the baseline results discussed in the previous subsection, we assumed that the central banks examined have a one-quarter horizon for each of the forward-looking variables (inflation and output gap). We examine the robustness of our results to this particular assumption. In particular, we consider in turn the cases in which central banks have a target horizon of up to one year for inflation and up to two quarters for the output gap. This is the horizon often referred to as appropriate in describing the lag with which monetary policy affects the target variables (Bernanke and Mihov, 1998; Clarida, Galí, and Gertler, 2000). The results from estimating equation (3) using these different horizons for the forward-looking variables are reported as specifications 2–5 in Table 3. These results suggest that, while the estimates of γ, δ, ϕ are

[7] As discussed in our “Robustness” section below, these findings hold up when using different measures of asset prices disequilibria.

Table 3. Augmented Forward-Looking Taylor Rule: Robustness Analysis

Panel A: United States	$\tilde{\beta}$	$\tilde{\gamma}$	$\tilde{\delta}$	$\tilde{\phi}$	$\tilde{\rho}$	J-test
1. $\pi_t = \pi_t^{GDP}$, $k_\pi = 1$, $k_x = 1$	1.226	0.817	0.065	0.011	0.721	0.961
	(0.209)	(0.135)	(0.010)	(0.005)	(0.069)	
2. $\pi_t = \pi_t^{CPI}$, $k_\pi = 4$, $k_x = 1$	1.289	0.981	0.075	0.017	0.785	0.946
	(0.215)	(0.337)	(0.013)	(0.007)	(0.064)	
3. $\pi_t = \pi_t^{GDP}$, $k_\pi = 4$, $k_x = 1$	1.562	0.824	0.064	0.008	0.719	0.935
	(0.267)	(0.385)	(0.014)	(0.004)	(0.078)	
4. $\pi_t = \pi_t^{CPI}$, $k_\pi = 4$, $k_x = 2$	1.048	0.907	0.049	0.022	0.788	0.995
	(0.210)	(0.278)	(0.017)	(0.006)	(0.069)	
5. $\pi_t = \pi_t^{GDP}$, $k_\pi = 4$, $k_x = 2$	1.752	0.890	0.046	0.008	0.779	0.999
	(0.116)	(0.109)	(0.005)	(0.004)	(0.042)	
6. $\pi_t = \pi_t^{CPI}$, $k_\pi = 1$, $k_x = 1$, $s_t = s_t^C$	2.035	0.796	0.076	0.037	0.765	0.994
	(0.049)	(0.106)	(0.008)	(0.015)	(0.052)	
7. $\pi_t = \pi_t^{CPI}$, $k_\pi = 1$, $k_x = 1$, $s_t = s_t^{BG}$	1.561	0.593	0.081	0.036	0.822	0.999
	(0.099)	(0.139)	(0.009)	(0.019)	(0.037)	

Panel B: United Kingdom	$\tilde{\beta}$	$\tilde{\gamma}$	$\tilde{\delta}$	$\tilde{\phi}$	$\tilde{\rho}$	J-test
1. $\pi_t = \pi_t^{GDP}$, $k_\pi = 1$, $k_x = 1$	0.925	0.778	0.075	0.008	0.886	0.998
	(0.013)	(0.040)	(0.006)	(0.002)	(0.002)	
2. $\pi_t = \pi_t^{CPI}$, $k_\pi = 4$, $k_x = 1$	1.439	0.423	0.098	0.007	0.892	0.867
	(0.211)	(0.09)	(0.013)	(0.003)	(0.004)	
3. $\pi_t = \pi_t^{GDP}$, $k_\pi = 4$, $k_x = 1$	1.553	0.462	0.034	0.006	0.813	0.913
	(0.388)	(0.014)	(0.008)	(0.003)	(0.003)	
4. $\pi_t = \pi_t^{CPI}$, $k_\pi = 4$, $k_x = 2$	1.462	0.284	0.036	0.009	0.863	0.999
	(0.343)	(0.046)	(0.006)	(0.005)	(0.002)	
5. $\pi_t = \pi_t^{GDP}$, $k_\pi = 4$, $k_x = 2$	1.531	0.269	0.032	0.011	0.816	0.905
	(0.355)	(0.017)	(0.008)	(0.003)	(0.003)	
6. $\pi_t = \pi_t^{CPI}$, $k_\pi = 1$, $k_x = 1$, $s_t = s_t^C$	1.041	0.535	0.064	0.023	0.891	0.999
	(0.020)	(0.034)	(0.004)	(0.001)	(0.001)	
7. $\pi_t = \pi_t^{CPI}$, $k_\pi = 1$, $k_x = 1$, $s_t = s_t^{BG}$	1.079	0.713	0.028	0.018	0.903	0.998
	(0.044)	(0.093)	(0.011)	(0.002)	(0.003)	

Panel C: Japan	$\tilde{\beta}$	$\tilde{\gamma}$	$\tilde{\delta}$	$\tilde{\phi}$	$\tilde{\rho}$	J-test
1. $\pi_t = \pi_t^{GDP}$, $k_\pi = 1$, $k_x = 1$	1.094	0.014	0.0012	0.0001	0.901	0.811
	(0.434)	(0.001)	(0.0005)	(0.0006)	(0.051)	
2. $\pi_t = \pi_t^{CPI}$, $k_\pi = 4$, $k_x = 1$	1.378	0.012	0.0015	0.0001	0.940	0.751
	(0.541)	(0.002)	(0.0006)	(0.0005)	(0.049)	
3. $\pi_t = \pi_t^{GDP}$, $k_\pi = 4$, $k_x = 1$	1.548	0.011	0.003	0.0001	0.859	0.765
	(0.543)	(0.002)	(0.0012)	(0.0004)	(0.058)	
4. $\pi_t = \pi_t^{CPI}$, $k_\pi = 4$, $k_x = 2$	1.277	0.010	0.007	0.0002	0.926	0.825
	(0.374)	(0.002)	(0.0017)	(0.0008)	(0.051)	
5. $\pi_t = \pi_t^{GDP}$, $k_\pi = 4$, $k_x = 2$	1.670	0.010	0.007	0.0001	0.894	0.788
	(0.318)	(0.001)	(0.0019)	(0.0006)	(0.032)	

(*continued*)

Table 3. *(concluded)*

Panel C: Japan	$\tilde{\beta}$	$\tilde{\gamma}$	$\tilde{\delta}$	$\tilde{\phi}$	$\tilde{\rho}$	J-test
6. $\pi_t = \pi_t^{CPI}$, $k_\pi = 1$, $k_x = 1$, $s_t = s_t^C$	1.564	0.012	0.0012	0.0003	0.928	0.825
	(0.227)	(0.001)	(0.0005)	(0.0014)	(0.038)	
7. $\pi_t = \pi_t^{CPI}$, $k_\pi = 1$, $k_x = 1$, $s_t = s_t^{BG}$	1.343	0.011	0.009	0.0003	0.967	0.827
	(0.441)	(0.001)	(0.0035)	(0.0018)	(0.029)	

Notes: See notes to Table 1. The estimated parameters refer to equation (3). k_π and k_x denote the forward-looking horizon for inflation and the output gap, respectively. Standard errors are reported in parentheses. s_t^C denotes the transitory component calculated as in Cochrane (1994) using *Datastream* stock price indices. s_t^{BG} denotes the log-differenced changes in the *Datastream* stock price indices (Bernanke and Gertler, 1999); under s_t^{BG}, six lags of the log-differences in the stock prices have been used.

virtually unaffected, the magnitude of β changes. However, the differences across specifications are not marked since the new estimates are always located in the confidence intervals implied by the standard errors of the parameters estimated in previous specifications.

Next we check for robustness the statistical significance of the parameter in the monetary policy rule associated with asset prices, using two different measures of asset price disequilibria. First we employ a permanent-transitory decomposition as in Cochrane (1994) to retrieve transitory deviations from the equilibrium asset price level. Second, following Bernanke and Gertler (1999), we employ lags of the log-differences in asset prices. As we can see in Table 3 (specifications 6 and 7), the estimates are qualitatively and quantitatively similar to our baseline estimates, confirming that for each country the coefficient in the monetary policy rules associated with asset price disequilibria are small in absolute value but statistically significant at conventional statistical levels.[8]

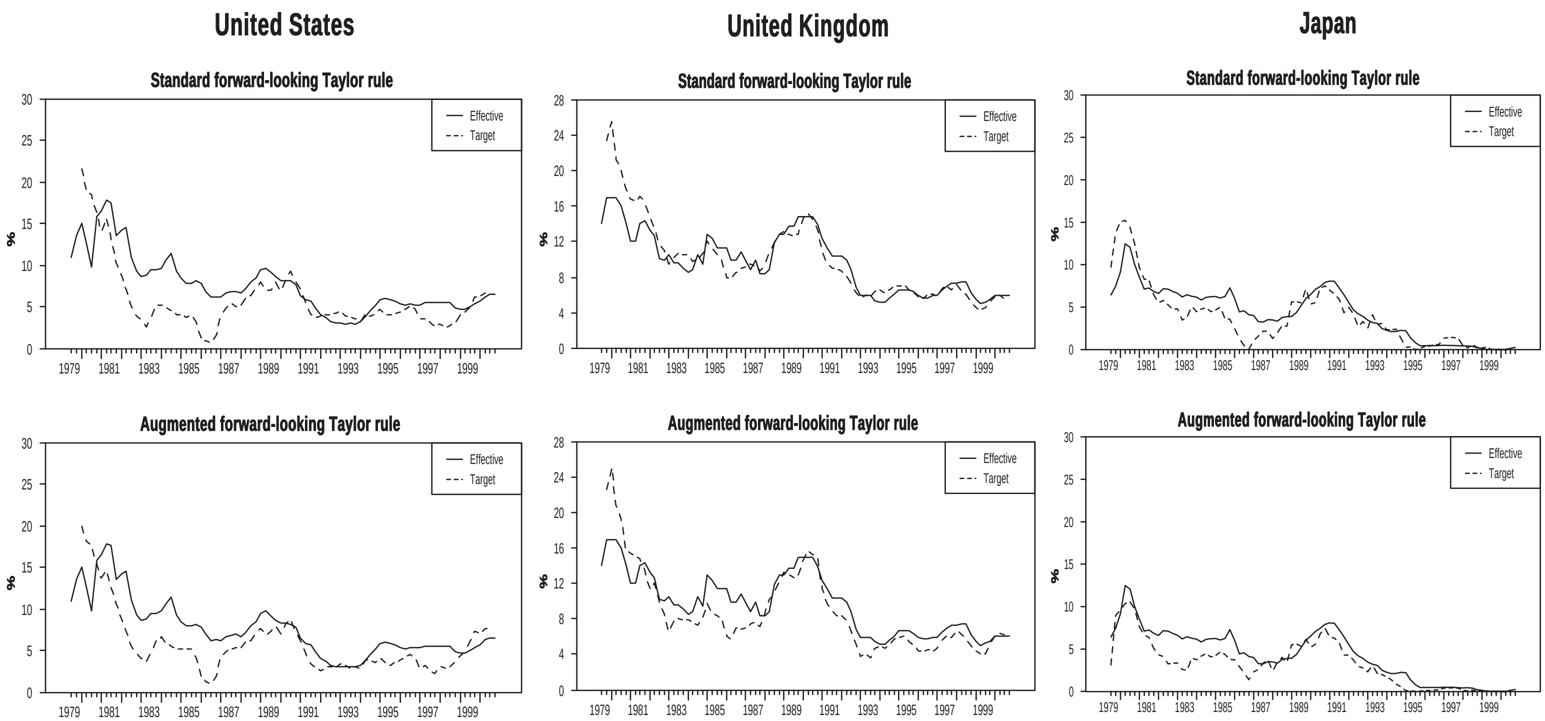
Overall, the robustness checks reported here suggest that our baseline estimates given in Table 2 appear to be fairly robust to changes in the proxy for inflation, asset price disequilibria, and to the choice of the horizon over which the central bank forms expectations of inflation and the output gap for the purpose of implementing monetary policy.

Discussion

The results of our baseline estimation reported in Table 2 deserve further discussion. To aid our discussion, in Figure 2 we report, for each country examined, the actual value and the estimated target value of the instrument rates implied by

[8]We also executed the same robustness exercise using the S&P 500, FTSE100, and NIKKEI 225 indices for the United States, the United Kingdom, and Japan, respectively, over a shorter sample period (because of data availability). In addition, we checked whether having asset prices and exchange rates enter the augmented policy rule contemporaneously made a difference. The results, not reported, are qualitatively and quantitatively similar to the ones reported in Tables 2 and 3.

Figure 2. Interest Rates: Actual and Target



specification 1 in Table 1 and by our baseline estimation reported in Table 2. The target value differs from the fitted value in that the latter incorporates the interest rate smoothing parameters and the former implicitly sets these parameters to zero (as in equation (1)).

The interest rates implied by the estimated rules characterize well the behavior of the instrument rates. Upon examination of Figure 2, we find that the two different specifications (i.e., the standard forward-looking Taylor rule and the augmented forward-looking Taylor rule, which allows for both exchange rates and asset prices) are both satisfactory in tracing the dynamics of the instrument rates. Indeed, this simple visual analysis of the models' target predictions suggests that the contribution of the two additional variables allowed for—asset price and exchange rate disequilibria—is not sufficient to distinguish clearly between the two specifications.

However, looking at Figure 3, where the contribution in percentage terms of each explanatory variable to the target value is shown, provides more insightful information. The graphs in Figure 3 suggest that, even if the level of the implied instrument rate target is very similar to the one implied by the standard forward-looking Taylor rule, the contribution of the different explanatory variables to its determination is different. In each country, central banks set up their targets mainly according to inflation dynamics. However, the other explanatory variables in the augmented policy rule also make some contribution to the instrument rate target determination.

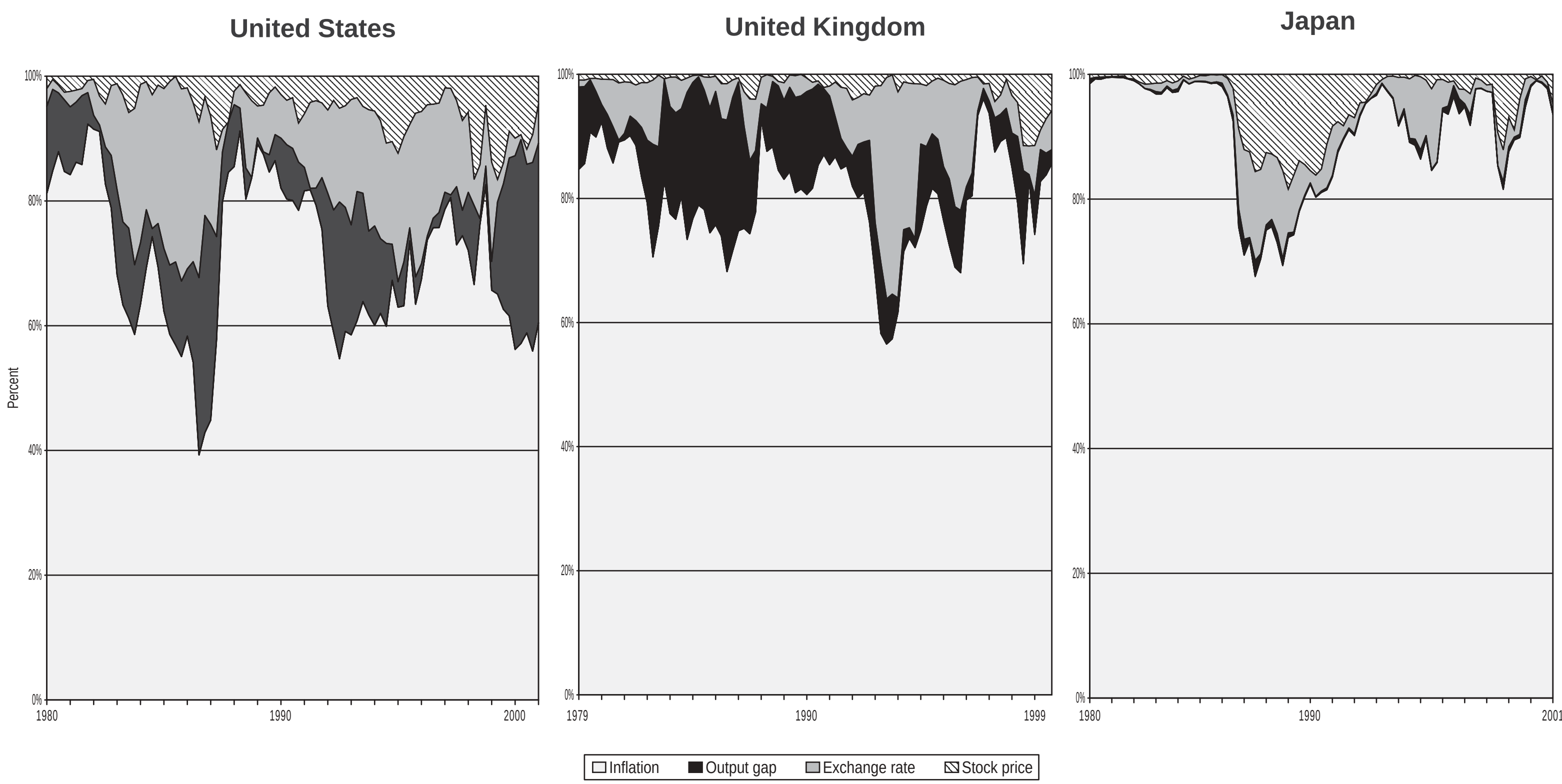
At face value these results may suggest that the Federal Reserve and the Bank of England use asset prices and exchange rates as targets and the Bank of Japan targets exchange rates, implying that they attempt to stabilize these variables in much the same way as they stabilize inflation or output. However, central bankers have often been clear that they do not target asset prices, either because asset prices are not necessarily related to the objectives of monetary policy or because they are only important insofar as they provide information about expected inflation. Alternatively, it could be that targeting asset prices is very difficult in practice given the difficulty of establishing their true equilibrium values (Gertler and others, 1998, and the references therein).

Nevertheless, several heuristic explanations can be used to justify our findings, which we present below. First, central banks may lie, because there may be gains from secrecy. They might target asset prices and exchange rates, but they do not want to admit it—for example, the case of foreign exchange intervention operations, which are often secret even though standard theory suggests that official intervention in the foreign exchange market is (more) effective if it is announced (Sarno and Taylor, 2001).[9]

Second, asset prices that have explanatory power in characterizing interest rate changes are only a proxy for that part of expected inflation and output that is not explained by the instruments used in the GMM estimation. Asset prices play a role in helping to get closer to the central banks' forecasting information set. The main implication of this conjecture is that central banks do not target asset prices per se, but they use them as good information variables (Clarida, 2001).

[9]See also the related arguments put forth by Carare and Stone (2003).

Figure 3. Interest Rates Target Decomposition



Third, central banks stabilize asset prices because they provide some information for aggregate price and output determination outside of their impact on firms' cost schedules. In other words, if asset prices only reflect the expected profit stream that accrues to firms that employ capital and labor at their rental prices and maximize profits subject to marginal costs, then asset prices are irrelevant. However, if asset prices diverge from this path and send incorrect signals to the corporate sector, there may be a prima facie case for offsetting this signal.

Finally, under a different perspective, our findings can be justified by the fact that what we are depicting in our estimation as a systematic rule may be, in fact, driven by occasional discretionary policy aiming, for example, at curbing bubbles or large misalignments in asset prices and exchange rates. Figure 3 supports this interpretation. This figure suggests that asset price and exchange rate disequilibria are particularly important in determining the interest rate target at times that are easily recognizable as large misalignments. For example, they are clearly recognizable for the United States, during the period following the 1987 crash of the stock price bubble, the dot-com bubble in the mid- to late 1990s, and the appreciation of the dollar in the mid-1980s. For the United Kingdom, one could claim that asset price disequilibria have hardly been economically important, while exchange rate disequilibria have been particularly important in the period around the sterling devaluation of 1992 that characterized the United Kingdom's exit from the Exchange Rate Mechanism and the protracted sterling appreciation of the mid- to late 1990s. For Japan, exchange rates were largely unimportant until 1986 or so, when they became important for some time, to then become very important in the mid-1990s, consistent with the results reported by Jinushi, Kuroki, and Miyao (2000) and McKinnon and Schnabl (2003).

Overall, it seems fair to conclude that, while both asset prices and exchange rate disequilibria are statistically significant in forward-looking Taylor rules, the evidence provided in this section suggests that, although these variables may not be key for *systematic* monetary policy, they induce a response from central banks when misalignments are relatively large.[10]

IV. Conclusions

We have examined the relationship between short-term interest rates, macroeconomic fundamentals, and asset prices and exchange rates to estimate monetary policy rules using quarterly data for the 1979–2000 period for the United States, the United

[10]Our interpretation that the reaction of central banks to asset price and exchange rate disequilibria may depend on the size of the disequilibria themselves suggests that a nonlinear characterization of interest rate rules may be a logical extension of our research. The possibility of a nonlinear reaction function can also be rationalized on the theoretical work of Bordo and Jeanne (2002) and is indeed suggested by the earlier work on escape clauses of Flood and Isard (1989). However, a nonlinear generalization is not straightforward in this context since the reaction function would be partly linear—in expected inflation and output gap—and partly nonlinear—in asset prices and exchange rates. To the best of the authors' knowledge, no GMM or instrumental variables estimator exists to date for models of this kind (or indeed for any multivariate threshold model), which makes estimation and statistical inference especially cumbersome in this context.

Kingdom, and Japan. In particular, we have not only limited our analysis to the effects of inflation and output disequilibria on the setting of short-term interest rates, but we have explicitly considered the effects of asset price and exchange rate disequilibria. Further, a subsidiary issue investigated is the extent to which the output gap is correctly measured in this context. Following the argument that the output-gap measure relevant for monetary policy purposes may be different from the one represented by detrended output (Woodford, 2001), we employed in our analysis real marginal costs as a proxy for the output gap.

By estimating monetary policy rules of the United States, the United Kingdom, and Japan, we find results that provide some evidence in favor of the conjecture that central banks respond to shifts in real marginal costs, rather than to any shock causing deviations of output from its trend. Further, in contrast with previous literature, we find that the parameters in the monetary rules associated with asset price and exchange rate disequilibria are statistically significant, yet small, for the United States and the United Kingdom, and exchange rates enter significantly the monetary policy reaction function for Japan. However, our results seem to suggest that asset prices and exchange rates may not be key for systematic monetary policy, albeit representing an important aspect of monetary policy design. In other words, we view our empirical results as suggesting that, in consideration of the risks that large misalignments in asset markets pose to macroeconomic stability and to the soundness of the financial sector, while targeting asset prices and exchange rates has not been a policy goal pursued systematically by the Federal Reserve, the Bank of England, and the Bank of Japan, these major central banks have reacted to these variables on a few occasions during the sample. More precisely, our findings are consistent with the view that, while committed to running monetary policy to keep inflationary pressures under control and bring output growth in line with potential, major central banks do act in response to exchange rates or asset prices on occasions when there is a need to prevent an abrupt correction in asset markets that could be destabilizing for the economy.

Although our baseline results have been shown to be robust to different specifications and proxy variables, several caveats are in order. Throughout the paper, the estimation of monetary policy rules has been carried out by means of the Generalized Method of Moments. The clear nonrejection of the null hypothesis of a valid instrument set in all our estimations, however, does not rule out the possibility that monetary rules may be misspecified. In particular, it might be the case that the weights assigned to different targets may be linked in a nonlinear fashion to the arguments in the reaction function, which would be in contrast with the implicit assumption of linearity maintained in conventional forward-looking monetary policy rules. Indeed, if our interpretation that asset prices and exchange rates are economically important in central banks' reaction functions when they are substantially misaligned from their fundamental equilibrium values, then a logical extension of this research involves using nonlinear reaction functions that may explicitly capture this behavior (see Bordo and Jeanne, 2002, for a theoretical rationale of nonlinearity in this context). Investigation of these issues remains on the agenda for future research.

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