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MONEY STOCK TARGETING, BASE DRIFT AND PRICE-LEVEL PREDICTABILITY: LESSONS FROM THE U.K. EXPERIENCE

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ABSTRACT

It is controversial whether money stock targeting without base drift (i.e. following a trend-stationary growth path) makes the price level more predictable in the presence of permanent shocks to money demand. Developing a procedure that does not run into the Lucas critique, and applying this procedure to the case of the U.K., the paper finds that the variance of the trend inflation rate in the U.K. would have been reduced by more than one half if the Bank of England had not allowed base drift.

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I. Introduction

During the 1970's, a number of countries attempted to control inflation and promote price stability by targeting monetary aggregates. In setting monetary targets, a general practice has been to calculate next period's target level on the basis of the actual rather than the (previously-announced) target level for the current period. According to this practice (hereafter referred to as base drift), when the money stock diverges from its target growth path, the divergence does not tend to be reversed later and thus results in a permanent change in the money stock. In other words, the money stock series would contain a unit root under base drift.

One popular criticism of the base drift policy is that it would introduce greater uncertainty about the long-run behavior of the money stock, and thus monetary targeting with base drift would not succeed in achieving long-run price stability.[1] Indeed, if the arguments of and shocks to the money demand function are trend stationary (so that the real money stock is trend stationary), the price level series has a unit root if and only if the money stock series has a unit root. In this case, the price level is clearly less predictable over long periods under base drift.

An interesting issue in this context is why a targeting policy whose major goal is long-run price stability would allow base drift under these circumstances. Goodfriend (1987) has suggested the explanation that base drift is induced by a tension arising between the price level smoothing and interest rate smoothing objectives of a central bank. Although base drift makes the price level less predictable over long periods (by introducing a unit root into the price level series), it is allowed because it also helps reduce the variability of interest rates.

An alternative view has been suggested by Walsh (1986, 1987), who has pointed out that if the demand for money is subject to permanent shifts, the price level would not be trend stationary even without base drift. Indeed, base drift would offset the effect of permanent shocks to money demand on the price level, and as Walsh (1986) has shown an optimal policy would involve some (between zero and full) base drift. It is thus possible that full base drift, as compared to no base drift, would make the price level more predictable over long periods. The goal of price stability alone would suffice in this case to explain why central banks follow the base drift policy.

The two views on the effect of base drift on the stochastic behavior of the price level can only be resolved by an examination of the empirical evidence. Such evidence is difficult to obtain because satisfactory structural models of price dynamics are lacking, and therefore, the econometric estimation of price behavior often relies on reduced-form equations. For reasons discussed in the well-known Lucas (1976) critique, reduced-form equations estimated for a regime of base drift cannot be used to predict the behavior of the price level under a counter-factual regime of no base drift.

To avoid this problem, the present paper focuses on the influence of base drift on the behavior of the permanent component of the price level (the trend price level). It can be shown that the forecast variance of the trend price level would dominate the forecast variance of the actual price level over long time horizons and hence provide useful information about long-term uncertainty about the price level. It is possible, moreover, to estimate the hypothetical behavior of the trend price level under no base drift (on the basis of data available from a base drift regime) using a procedure developed in this paper. The procedure exploits the widely accepted proposition that money is neutral in the long run. Empirical implementation of our procedure requires estimation of the long-run components of both the price level and money stock. We estimate these components using an approach that has its roots in the Beveridge and Nelson (1981) method for decomposing univariate series into permanent and transitory components.[2]

Our empirical work focuses on the monetary experience of the United Kingdom since 1976. The Bank of England has allowed full base drift in setting targets for sterling M3, and has implemented its target policy by using interest rate control. Although other countries have also pursued monetary targeting with base drift, the United Kingdom's experience with this policy represents one of the longest periods of targeting without a change in the control procedure. In addition, the United Kingdom has been less successful than most other countries [e.g., the U.S., Canada and Germany] in hitting its targets and has experienced greater variability in its price level than most other countries. Thus the U.K. experience with monetary targeting provides a good case study of the influence of base drift on price stability.

Section II of the paper uses a model with flexible prices and rational expectations to provide a simple example of conditions under which base drift may either increase or decrease the forecast variance of the price level. Section III paves the way for our empirical analysis by introducing a framework in which each variable is decomposed into permanent and transitory components. Using this framework, we then explain our procedure for estimating the behavior of the trend price level in the absence of base drift, making use of the data available from a targeting regime in its presence. Applying this procedure to the case of the United Kingdom in section IV, we examine whether the trend price level in the United Kingdom would have been less predictable if the Bank of England had not allowed any base drift. Our results show that the policy of no base drift would have reduced the forecast variance of the trend price level by slightly more than one half. Interestingly, this substantial reduction in the variance occurs even though the demand for £ M3 in the United Kingdom has exhibited permanent shifts. Thus our evidence is consistent with the view that base drift increases price level uncertainty.

II. Base Drift and Price-Level Variability: A Simple Model

To explore the influence of base drift on the behavior of the price level, we begin with a simple stochastic model that assumes flexible prices and rational expectations. As is the case in the United Kingdom, we assume that the central bank uses an interest rate control procedure to achieve its money stock targets. We also assume that information on the money stock and the price level becomes available (to both the central bank and the public) after a one-period lag. Our set-up is a modified version of the (1981) model that McCallum used to analyze the implications of an interest rate policy rule for price level determinacy. The key differences between our model and McCallum's are that we allow for the possibility of base drift and include a permanent shock in the money demand function. In McCallum's model, the central bank pursues the objective of interest-rate smoothing. Goodfriend (1987) has suggested that this objective induces central banks to incorporate base drift in setting their targets. To keep our theoretical example simple, however, we assume that the central bank is concerned only with keeping the money stock on target (our more general empirical model in section III, however, does allow for other goals such as interest-rate smoothing).

Our model is described by the following equations:

$$m_t = p_t + \alpha_t - \beta r_t + e_t , \tag{1}$$

$$\alpha_t = \alpha_{t-1} + a_t , \tag{2}$$

$$p_t = E_{t-1}p_{t+1} - r_t + \delta_t , \tag{3}$$

$$\delta_t = \delta_{t-1} + d_t , \tag{4}$$

$$\mu_t = (1-\theta)m_{t-1} , \tag{5}$$

$$r_t = -(1/\beta)[\mu_t - E_{t-1}(p_t + \alpha_t + e_t)] , \tag{6}$$

where m_t and p_t represent the logarithms of the money stock and the price level, r_t is the nominal rate of interest, e_t, a_t and d_t are white noise disturbances, and the operator E_{t-1} denotes the expectation of the indicated variable conditional on information on all variables in the model up to period t-1.

Equations (1) and (2) represent the money market. The demand for money is assumed to depend on a permanent shock α_t as well as a temporary shock e_t. For simplicity, we assume that the permanent shock is a random walk while the temporary shock is serially uncorrelated. Equations (3) and (4) summarize the goods market. The variable δ_t in these equations is the real rate of interest, and like α_t, is assumed to follow a random walk. Following McCallum (1981), we assume that an "island" model underlies the determination of δ_t.[3] As δ_t in such a model would represent an average of "local" real rates (that utilize global information on the nominal interest rate but only local information on prices), we use $E_{t-1}p_{t+1}$ rather than E_tp_{t+1} to express the expected value of the next period's price level in (3). Note that shocks to the real interest rate (d_t) may be correlated with those to the demand for money (a_t and/or e_t) but we cannot be sure about the signs of these correlations without further specification of the underlying model.[4] Finally, (5) and (6) provide a simple characterization of money-stock targeting with interest rate control. In these equations, μ_t represents the target level of m_t. To simplify the discussion, we assume that the target rate of money growth equals zero. The setting of the target stock in (5) allows for full base drift when $\theta = 0$ and no base drift when $\theta = 1$. As our concern here is to highlight the difference between two widely-discussed policy alternatives of zero or full base drift, we do not attempt to derive an optimal policy rule. Assuming that the central bank pursues no other goal, the interest rate is set in (6) such that the expected stock of money equals the target stock. Note that given the one-period information lag, m_t can diverge from μ_t because of unanticipated changes in money demand and the price level.

The above model is easily solved by the method of undetermined coefficients to yield the following solutions for m_t, p_t and r_t:

$$m_t = (1-\theta)m_{t-1} + a_t + e_t + d_t \,, \tag{7}$$

$$p_t = [(1-\theta)/(1+\beta\theta)]m_{t-1} - \alpha_{t-1} + \beta\delta_{t-1} + d_t \,, \tag{8}$$

$$r_t = -[\theta(1-\theta)/(1+\beta\theta)]m_{t-1} + \delta_{t-1} \,. \tag{9}$$

With no base drift, $\theta = 1$, and the coefficient of m_{t-1} equals zero in (7) as well as (8). In this case m_t is a stationary series but p_t is still non-stationary because of permanent shocks to both the demand for money and the real rate of interest. In contrast, if there is full base drift, $\theta = 0$, and the coefficient of m_{t-1} equals 1 in both (7) and (8). Now m_t also becomes a non-stationary series. Interestingly, with either $\theta = 1$ or $\theta = 0$, the coefficient of m_{t-1} equals zero in (9). In both of these cases r_t equals δ_{t-1} (the expected value of δ_t conditional on t-1 information), and thus the behavior of the interest rate remains the same in the two regimes.

Letting $p_t(1)$ and $p_t(0)$ represent the behavior of p_t for $\theta = 1$ and $\theta = 0$, respectively, and using (2), (4) (7) and (8), we have

$$p_t(1) = -\alpha_{t-2} - a_{t-1} + \beta\delta_{t-1} + d_t \,, \tag{10}$$

$$p_t(0) = m_{t-2} - \alpha_{t-2} + e_{t-1} + \beta\delta_{t-1} + d_{t-1} + d_t \,. \tag{11}$$

Comparing the behavior of prices in (10) and (11), it is clear that while the policy of full base drift, as compared to no drift, eliminates the shock a_{t-1} from the price equation, at the same time it introduces new shocks e_{t-1} and d_{t-1}.

To explore the influence of base drift on price variability, we let $FV_t^k(\theta) \equiv E_t[p_{t+k}(\theta) - E_t p_{t+k}(\theta)]^2$, $\theta = 1, 0$, denote the k-period forecast variance of p, and use (2), (4) (10) and (11) to obtain

$$FV_t^k(1) = (k-1)\beta^2\sigma_d^2 + \sigma_d^2 + (k-1)\sigma_a^2 - 2(k-1)\beta\sigma_{ad}, \qquad (12)$$

$$FV_t^k(0) = (k-1)(1+\beta)^2\sigma_d^2 + \sigma_d^2 + (k-1)\sigma_e^2 + 2(k-1)(1+\beta)\sigma_{ed}, \qquad (13)$$

where σ_a^2 and σ_e^2 are the variances of a and e while σ_{ad} and σ_{ed} are the covariances between a and d, and e and d. According to (12) and (13), the one-period forecast variances are the same (and equal to σ_d^2) with or without base drift. However, for longer forecast intervals, the two regimes imply different variances. For k > 2, it follows from (12) and (13) that

$$FV_t^k(0) - FV_t^k(1) = (k-1)[\sigma_e^2 + (1+2\beta)\sigma_d^2 - \sigma_a^2 + 2(\beta\sigma_{ad} + (1+\beta)\sigma_{ed})]. \qquad (14)$$

Thus, when k > 2 the difference between the full-drift and no-drift k-period forecast variances increases with variances of shocks e and d but decreases with that of shock a. This difference is also influenced by the covariances between shocks a and d, and e and d, but the direction of the influence would depend on whether these covariances are positive or negative.

As the above simple analysis illustrates, which targeting regime would lead to a lower variability of the price level is essentially an empirical question. In the next section, we develop a procedure that can be used to estimate forecast variances of the trend price level for the two regimes, using data generated under the base-drift regime.

III. A Methodology for Estimating the Influence of Base Drift on the Trend Price Level

Reduced-form equations [(7) - (9)] for m_t, p_t and r_t in the previous section were derived from a particular model. In this section, we consider reduced-form equations for these variables of a very general form that would be consistent with a broad class of models of assets and goods markets and of monetary policy rules. We only require that the underlying structure obey certain minimal restrictions on the behavior of the permanent components of these variables. To incorporate these restrictions into the analysis, each variable is decomposed into a trend or a permanent component and a cyclical or a transitory component as follows:

$$y_t = \bar{y}_t + u_t , \tag{15}$$

where y_t is a 3 x 1 vector of variables m_t, p_t, r_t; $\bar{y}_t = [\bar{m}_t \; \bar{p}_t \; \bar{r}_t]$ a vector of permanent components of these variables; and $u_t = [u_{1t} \; u_{2t} \; u_{3t}]$ a vector of transitory components. The permanent components are assumed to evolve as

$$\bar{y}_t = \gamma + \bar{y}_{t-1} + v_t , \tag{16}$$

where $\gamma = [\gamma_1 \gamma_2 \gamma_3]$ is a vector of constants and $v_t = [v_{1t} \; v_{2t} \; v_{3t}]$ a vector of shocks to the permanent components. Both u_t and v_t are covariance stationary.

We assume that money is neutral in the long run in the sense that the distribution of permanent components of real variables is independent of the behavior of the money stock. In this context, permanent components could be viewed as representing values in the long run or the "natural" equilibrium. The concept of the long-run equilibrium, however, is often not well defined and its meaning differs from one model to another. For example, deviations of variables from their long-run values are explained in terms of price stickiness by one approach, and information lags by another. The length of time required to reach the long-run equilibrium could thus depend on what type of short-run friction is assumed.

To avoid identifying the long run with a specific period of time, it is appealing to use the concept of the permanent component suggested by Beveridge and Nelson (1981). According to this concept

$$\bar{y}_t = \lim_{k \to \infty} (E_t y_{t+k} - k\gamma) \quad . \tag{17}$$

The permanent components defined by (17) represent forecasts (after adjusting for deterministic trends) of corresponding variables in the future far enough to eliminate the influence of all types of short-run frictions. These forecasts, moreover, utilize all current information available in the model. As discussed below, an immediate implication of (17) is that the shocks to permanent components (v_t) are white noise disturbances. Also note that one or more components of v_t can be set equal to zero. In such special cases, the corresponding permanent component(s) would simply follow a deterministic trend line over time.

Having defined our measure of permanent components, we next discuss two relations that link the permanent components of the two real variables in our model, the real stock of money and the real rate of interest. First, we assume a "long-run" demand for money of the form:

$$\bar{m}_t - \bar{p}_t = \bar{\alpha}_t - \beta \bar{r}_t \ , \tag{18}$$

where $\bar{\alpha}_t$ is a random variable that represents shifts in the long-run demand. Second, we express the permanent component of the nominal interest rate as

$$\bar{r}_t = \bar{\delta}_t + E_t(\bar{p}_{t+1} - \bar{p}_t) \quad , \tag{19}$$

where $\bar{\delta}_t$ is the permanent component of the real interest rate (i.e., the natural rate of interest). In conformity with (17), we use E_t as the expectation operator but our analysis would not be affected if (as in our model in the previous section) we used E_{t-1} instead. Note that since v_{2t} is a white noise, (16) implies that the second term in (19) is a constant. Although our analysis in this section makes use of only the long-run demand for money it is interesting to note that (15) and (18) imply the following "short-run" money demand:

$$m_t - p_t = \bar{\alpha}_t - \beta r_t + e_t', \tag{20}$$

where $e_t' = u_{1t} - u_{2t} + \beta u_{3t}$, and is thus a stationary random variable.

Using (16) and (18), moreover, $\bar{\alpha}_t = \bar{\alpha}_{t-1} + \gamma_1 + v_{1t} - \gamma_2 - v_{2t} + \beta(\gamma_3 + v_{3t})$, and thus it represents either a random walk or a deterministic trend (in the special case where $v_t = 0$). We can, therefore, consider $\bar{\alpha}_t$ as representing permanent shocks and e_t' temporary shocks to money demand in (20).

Now our assumption that money is neutral in the long run can be restated as that $\bar{m}_t$ does not affect the distributions of both $\bar{\alpha}_t$ and $\bar{\delta}_t$. An important implication of this assumption that we exploit below is that the behavior of both $\bar{\alpha}_t$ and $\bar{\delta}_t$ will be the same under regimes of full and no base drift. Note that we need not assume superneutrality as the deterministic trend rates of money growth and inflation (γ_1 and γ_2) are assumed below to be the same in both regimes.[5] The behavior of $\bar{m}_t$ would differ between the two regimes, however, as this series would be a trend stationary process under no base drift but non-stationary under base drift.

To examine the behavior of $\bar{m}_t$ under different regimes, let the money stock target be set as follows:

$$\mu_t = \gamma_1 + (1-\theta)m_{t-1} + \theta\mu_{t-1} . \tag{21}$$

The setting of the money stock target in (21) is more general than (5) in that it allows for a deterministic trend rate of money growth equal to γ_1. We also now consider the possibility that the central bank might be concerned with goals (such as interest-rate smoothing) other than targeting of the money stock. In the theoretical model of the previous section, the money stock deviated from the target path because of the one-period information lag. The deviation, moreover, was a white noise disturbance. The pursuit of other goals would now provide another reason why the money stock would deviate from the target path. The deviations of m_t from μ_t for this reason could be serially correlated. The central bank policy is assumed, however, to ensure that the money stock reverts to its targeted path in the long run. The difference $(m_t - \mu_t)$, therefore, would represent a stationary series.

In the discussion below, we use the notation $x_t(\theta)$, $\theta = 0, 1$, to represent the value of a variable x_t under regimes of full and no base drift, respectively. The behavior of the money stock under the two regimes can be derived from (21) as

$$m_t(1) = \mu_0 + \gamma_1 t + z_t(1) , \qquad (22)$$

$$m_t(0) = \gamma_1 + m_{t-1}(0) + z_t(0) , \qquad (23)$$

where t is a time trend, μ_0 is the value of μ_t for t = 0, and $z_t = m_t - \mu_t$. Since z_t is a stationary series, $m_t(1)$ is generated by a trend-stationary process while $m_t(0)$ is generated by a difference-stationary process.

Now suppose that for a certain period, the central bank of a country follows a targeting policy with full base drift. We could use the data for this regime to estimate $\bar{p}_t(0)$. We discuss below our econometric procedure for estimating permanent components. Here, we first explain how the data for the base-drift regime can also be used to estimate $\bar{p}_t(1)$. Our assumption of long-run neutrality of money implies that $\bar{\alpha}_t(1) = \bar{\alpha}_t(0)$ and $\bar{\delta}_t(1) = \bar{\delta}_t(0)$. Our model in this section also implies that the expected rate of trend inflation , $E_t(\bar{p}_{t+1} - \bar{p}_t)$, equals γ_2 (the drift term in the equation for $\bar{p}_t$). Given that γ_2 is the same under the two regimes, $\bar{r}_t(1) = \bar{r}_t(0)$ according to (19). In view of (18), it follows that the real stock of money would be the same under the two regimes, and hence

$$\bar{p}_t(1) = \bar{m}_t(1) - \bar{m}_t(0) + \bar{p}_t(0) . \tag{24}$$

As $\bar{m}_t(1) = \mu_0 + \gamma_1 t$, according to (22), (24) shows that we need estimates of only $\bar{m}_t(0)$ and $\bar{p}_t(0)$ in order to obtain estimates of $\bar{p}_t(1)$.

To estimate the permanent component of the money stock and the price level under base drift, we use a measure that is based on a multivariate version of the Beveridge-Nelson (1981) methodology. To explain this measure, we first note that since Δy_t is covariance stationary, it will have the following Wold representation:

$$\Delta y_t = \gamma + C(L)\varepsilon_t , \tag{25}$$

where $C(L)$ ($\equiv C_0 + C_1 L + C_2 L^2 + \ldots.$) is a 3 x 3 matrix of polynomials in the lag operator L, and $\varepsilon_t = [\varepsilon_{1t}\ \varepsilon_{2t}\ \varepsilon_{3t}]$ a vector of innovations in m_t, p_t and r_t. Using (17) and (25), it can be shown that

$$\bar{y}_t = \gamma + \bar{y}_{t-1} + D\varepsilon_t , \tag{26}$$

where $D = \sum_{i=0}^{\infty} C_i$ is the matrix of long-run multipliers. The measure of $\bar{y}_t$ given by (26) always exists. Also comparing (26) with (16), it is clear that since each element of v_t is a linear function of the components of ε_t, it represents a white noise process.

It is important to emphasize that while the long-run neutrality of money may have implications about how structural disturbances affect permanent components, it does not imply any restrictions on the effects of reduced-form shocks ε_t on $\bar{y}_t$ (i.e., on the elements of matrix D). For example, given a constant expected trend inflation rate, a nominal shock would not affect $\bar{r}_t$ or $\bar{m}_t - \bar{p}_t$ according to the neutrality proposition. However, as the shock to the reduced-form equation for $m_t(\varepsilon_{1t})$ would in general be a combination of both nominal and real structural disturbances, its effects on $\bar{r}_t$ or $\bar{m}_t - \bar{p}_t$ are not restricted.

In the above discussed Beveridge-Nelson decomposition, innovations in both the permanent and transitory components are the same. Other models of the decomposition of a variable into permanent and transitory components allow innovations in the two components to be different and introduce a priori restrictions on correlations between the two innovations. As Cochrane (1988) has demonstrated (in terms of a univariate process), however, the innovation variance of the permanent component is the same regardless of what decomposition is used. Thus, although we use (26) to estimate the forecast variances of $\bar{p}_t(0)$ and $\bar{p}_t(1)$, our estimates of these variances would not change if another model of the permanent-transitory decomposition were chosen.

As discussed in the next section, innovation variances of $\bar{p}_t$ under the two regimes can be calculated using estimates of D and of the covariance matrix of ε derived from the base drift regime. To obtain these estimates, our strategy is to identify and estimate an appropriate VAR system for the United Kingdom. The VAR system is then used to obtain estimates of D and the covariance matrix of ε.

IV. Empirical Results for the United Kingdom

Before presenting evidence on the effect of base drift in the United Kingdom, we briefly describe the targeting policy followed by the Bank of England. The Bank began announcing monetary growth targets for dates from July 1976. The decision to adopt explicit monetary targets was apparently taken in recognition of the need for monetary control to limit inflation and sterling depreciation. The Bank chose to target the broad aggregate sterling M3 (£M3)[6] rather than the narrower M1 aggregate followed by the U.S. and Canada for several reasons. First, econometric studies undertaken in the early 1970's indicated that the demand for £M3 was more stable than that for M1. Second, £M3 corresponds more closely to items on the asset side of the consolidated Banking sector balance sheet which the Bank believes it can directly influence through its policy [see Goodhart (1983)] . Since 1982, the Bank has also started declaring targets for other monetary aggregates. But as the experience with these other targets is not very long, this paper focuses on the Bank's targeting of £M3.

Each year since 1976, the Bank has announced a target range for the rate of growth of £M3. The target range normally applies to a 12-month period and the rate of growth is calculated using the actual (rather than the previously-announced) stock of £M3 in a specific month of the year as the base.[7] This procedure thus allows a base drift to occur every year. Figure 1 shows the behavior of both the actual and the target levels of £M3. The target levels are calculated using the mid-points of the announced target ranges for the rates of £M3 growth. As the figure shows actual £M3 rose significantly above the target path during 1980-82. Given the base drift policy, however, the target levels were adjusted upwards in this period. This adjustment made it possible for £M3 to stay close to the target path from the middle of 1982 to the beginning of 1985. In Figure 1 we also show the hypothetical target path that would have obtained if the Bank allowed no base drift and used a rate of money growth equal to the average of announced rates. It is clear from the figure that without base drift the target levels would have been much lower since 1980 and would have induced a very different monetary policy than actually followed.

To hit its monetary targets, the Bank uses an interest rate control procedure. For several years after the Competition and Credit Control Act of 1971, the Bank followed a procedure similar to that of the Federal Reserve before 1979 and the Bank of Canada before 1982, that is, interest rates were set to make the demand for £M3 equal to what the Bank wished to supply. Disillusioned with the margin of error surrounding the money demand function in 1972-73, the Bank switched to a policy focused directly on the asset counterparts to £M3. According to an accounting identity based on consolidating the balance sheets of the Bank of England's banking department and the commercial banks, the Bank links changes in £M3 to asset counterparts including as principal components: changes in bank lending to the private sector; the Public Sector Borrowing Requirement (PSBR) less private lending to the government i.e. the sale of government securities (gilts) to the public. Based on this identity and information about the PSBR and forecasts of bank lending, market interest rates (such as the rate on 3-month Treasury bills) are set to sell the required amounts of gilts necessary to hit the £M3 target.[8]

We next discuss the data used to estimate an empirical model for the United Kingdom based on the methodology of section III. We considered both monthly and quarterly data. The variable r was measured by the 3-month Treasury bill rate expressed as a fraction, and m by the logarithm of £M3.[9] Two different price indexes, logarithms of the Consumer Price Index (known as the retail price index in the U.K.) and the GDP deflator (available only on a quarterly basis), were used to measure p on a monthly and quarterly basis.[10] A month (as compared to a quarter) appears to be a more appropriate unit of time for the purpose of representing the Bank of England's policy of interest-rate control. However, since the quarterly data includes a more satisfactory price index, this paper focuses on estimates based on quarterly data. The results derived from the monthly model are not reported but are similar.

To facilitate the selection of an appropriate form of the VAR system, Table 1 tests the three series, m, p and r, for stationarity and cointegration for the period 1977:2 to 1985:4.[11] Panel A of this table presents two types of tests of the unit-root hypothesis for a univariate series: one based on Dickey and Fuller (1979) and the other on Stock and Watson (forthcoming). According to both tests, the results do not reject the hypothesis that the series in levels of m, p and r all contain a unit root. The indication of a unit root in the m series, moreover, is fully consistent with our interpretation that the Bank's targeting procedure involves full base drift. In the case of first-differenced series, the unit-root hypothesis is rejected at the conventional levels for Δr according to both tests and for Δm according to the Stock-Watson test. The case for rejecting the hypothesis for Δp is less strong. The Stock-Watson statistic rejects the hypothesis that a unit root is present in Δp only at the 17% level. The Dickey-Fuller statistic also does not reject the hypothesis at the conventional levels but the standard error of ρ (the coefficient of the lagged dependent variable) for Δp is large and the power of the test is not high in this case.[12] Thus we do not consider this evidence to provide strong indication of non-stationarity in Δp.

Panel B of Table 1 provides Stock-Watson tests of common trends in m, p and r. As the results show, the hypothesis that these series have three distinct unit roots is clearly not rejected against the alternatives of one or two unit roots. The absence of common trends in m, p and r is consistent with the view that the demand for money is subject to permanent shocks (the shift variable $\bar{\alpha}_t$ contains a unit root).[13] In view of the above evidence, we assume that m, p and r are first-difference stationary and are not cointegrated with each other. We thus estimate a VAR system where each of the three variables is entered in the first difference form.

Before describing our results further, we note that as emphasized recently by Cochrane (1988), tests of unit roots have a low power in the sense that it is difficult to distinguish a stationary series from a stationary series plus a small random walk. It is thus instructive to examine how big the random walk component is in the series. Cochrane (1988) suggests that the variance of the shock to the random walk component relative to the variance of the first difference of the series provides a good measure of the size of the random walk component. He uses the variance of the long difference of the series (i.e., the difference between values over long periods) to estimate the variance of the shock to the random walk component. The targeting regime in the UK is not long enough to provide a satisfactory estimate of this statistic. However, as discussed below we do estimate the variance of $\Delta\bar{p}(\Delta\bar{m})$ from the VAR model and this variance is large in relation to the variance of $\Delta p(\Delta m)$. Thus, at least on the basis of the VAR estimates, the random walk components do not appear to be small in these series.

The VAR model is estimated over the period 1976:2 to 1985:4 and includes four lags for each variable (the first observation for the dependent variable is thus 1977:2). The estimation period was not long enough to explore additional lags. We did consider models with two or three lags but these were rejected against the alternative of a model with four lags. We also introduced a time trend in (each equation of) the system but as this variable was found to be insignificant, it was dropped from the model.

The Thatcher administration which began in 1979 introduced a number of programs including the Medium Term Financial Strategy, in which announced money growth targets were to be reduced over a sequence of years. This strategy was intended to give financial markets some indication of the government's objectives. One issue is whether the monetary policy regime actually changed after Thatcher took office. Again, we did not have sufficient degrees of freedom to examine whether VAR coefficients were significantly different before and after Thatcher. As a crude attempt to explore the influence of Thatcher, however, we did try a dummy variable (equal to one after 1979:2, zero otherwise) in each equation but this variable also turned out to be insignificant.

As the trend price level is a random walk, the one-period forecast variance of $\bar{p}$ is the same as the variance of $\Delta\bar{p}$.[14] This variance is estimated from the VAR system as follows: Letting $V\Delta\bar{p}(0)$ and $V\Delta\bar{p}(1)$

denote variances of $\Delta\bar{p}$ under full and no base drift, noting that $\Delta\bar{p}_t(1) = \gamma_1 + \Delta\bar{p}_t(0) - \Delta\bar{m}_t(0)$ according to (24), and using (26) to estimate $\Delta\bar{p}(0)$ and $\Delta\bar{m}(0)$, we obtain

$$V\Delta\bar{p}(0) = D_2\Sigma_\varepsilon D_2', \tag{27}$$

$$V\Delta\bar{p}(1) = (D_2-D_1)\Sigma_\varepsilon(D_2-D_1)', \tag{28}$$

where D_1 and D_2 are the first two rows of matrix D under base drift, and Σ_ε is the covariance matrix of shocks ε under the same regime. As discussed below, estimates of D_1, D_2 and Σ_ε are readily obtained from a VAR system.

One general problem associated with the use of a VAR model is that impulse response functions generated from VAR residuals do not generally provide meaningful information about the effect of structural disturbances. For our present purpose, however, it is easy to show that the variance of $\Delta\bar{p}$ remains unchanged regardless of whether it is estimated in terms of VAR residuals or structural disturbances. For instance, let the structural model be

$$A\Delta y_t = \pi + B(L)\Delta y_{t-1} + \eta_t, \tag{29}$$

where π is a vector of constants, B(L) is a matrix of polynomials in the lag operator L, and η_t is a vector of structural disturbances. Premultiplying both sides of (29) with A^{-1}, we obtain the following VAR form (that we estimate in this section):

$$\Delta y_t = \phi + F(L)\Delta y_{t-1} + \varepsilon_t, \tag{30}$$

where $\varepsilon_t = A^{-1}\eta_t$, $F(L) = A^{-1}B(L)$ and $\phi = A^{-1}\pi$. Given that (30) can be inverted to obtain the moving average process (25), (29) will imply a moving average representation in terms of the structural disturbances as follows:

$$\Delta y_t = \gamma + C(L)A^{-1}\eta_t. \tag{31}$$

Using (31) it is straightforward to show that the variance of $\Delta\bar{p}$ in terms of structural disturbances η_t is exactly the same as that in terms of VAR residuals ε_t. It can similarly be shown that any Choleski orthogonalization of shocks to the VAR system would not make any difference to the variance of $\Delta\bar{p}$.

In Table 2 we show certain results from the VAR model that are needed to estimate the variance of $\Delta\bar{p}$. Panel A of this table shows the correlation/covariance matrix of the residuals in the three equations. Panel B displays the accumulated responses (over 10, 20, 30 and 40 quarters) of both Δm and Δp to a unit shock to each of the three equations. The accumulated responses do not tend to change much beyond 20 quarters. We use the sums of responses over 40 quarters to approximate the long-run multipliers that correspond to the elements of D_1 and D_2.

Using the above estimates, rows 1 and 2 of Table 3 show the variances of the trend inflation rate with and without base drift [i.e., $V\Delta\bar{p}(0)$ and $V\Delta\bar{p}(1)$] calculated according to (27) and (28). As can be seen from the ratio of the two variances in row 3 of this table, the variance of $\Delta\bar{p}$ under no base drift is less than one-half of that under base drift. Our estimates thus imply that a targeting policy without base drift would have brought about a large reduction in the variability of the trend inflation rate. To illustrate this result, we construct the series $\Delta\bar{p}_t(0)$ and $\Delta\bar{p}_t(1)$ according to (24) and (26), using estimates of ε_{1t} and ε_{2t} available from the VAR model. These series are exhibited in Figure 2. As the figure clearly demonstrates, the variability of the trend inflation rate would have been much smaller under a policy of no base drift.

As our empirical work is concerned with estimating the effect of base drift only on the behavior of the permanent component of the price level, it is interesting to examine how big the permanent or the random walk component is in this series. As discussed above, Cochrane (1988) has suggested that one way to answer this question is to compare the variance of $\Delta\bar{p}$ with that of Δp. The variance of Δp estimated for the period 1977:2 to 1988:4 is shown in row 4 of Table 3. Comparing this variance with our estimate of the variance of $\Delta\bar{p}$ under base drift, row 5 shows that the latter is about seven times as large as the former.[15] The size of the permanent component thus seems to be very prominent in the case of the UK price level.[16]

V. Conclusions

If the real stock of money includes a random walk component, the price level would not be trend stationary even if targeting policy allows no base drift. In this case, it is not clear whether the presence of base drift would make the price level more or less predictable. Our theoretical analysis suggests that the answer to this question depends on the relative strength of different types of shocks. According to our analysis in section II the price level would be more predictable with than without base drift if permanent shocks to money demand dominate. The model yields the opposite result, however, if shocks to the real rate of interest (and/or temporary shocks to money demand) dominate.

To obtain empirical evidence on this issue, the paper develops a procedure for estimating the effect of base drift on the forecast variance of the trend price level or equivalently the variance of the trend inflation rate (which is an indicator of the predictability of the actual price level over long periods). We find that the practice of base drift in the U.K. was responsible for lower predictability of the trend price level. The case in favor of base drift made by Walsh (1986) relies on the argument that money demand is subject to permanent shifts. For the U.K., the error term in money demand is indeed non-stationary but despite this evidence of permanent shifts in the U.K.'s money demand we estimate that a policy of allowing no base drift would have decreased the forecast variance of the trend price level in the U.K. by more than one half.

This paper does not explore the issue of why the Bank of England permitted base drift to reduce the predictability of the price level over long periods. The reason may well lie in Goodfriend's (1987) explanation that base drift is induced by the objective of smoothing interest rates. Such a goal may have been followed by the Bank to ensure orderly financial markets, to aid the government in meeting its fiscal objectives and to stabilize the exchange rate.

The empirical analysis in this paper is based on a general model that restricts the underlying structure essentially in requiring that money be neutral in the long-run -- that is, the behavior of permanent components of real variables be the same under different monetary regimes. Although this paper focuses on the influence of base drift, the empirical methodology can clearly be used to examine the effect of other types of changes in monetary regimes on the long-run predictability of the price level.

Footnotes

1. For this and other criticisms of the base drift policy, see, for instance, Poole (1976), Friedman (1982), and Broaddus and Goodfriend (1984). Also see the (1985) report of the Shadow Open Market Committee and the (1985) Economic Report of the President.

2. For an extention of the Beveridge-Nelson methodology to multivariate models, see, for example, Huizinga (1987) and King, Plosser, Stock and Watson (1987).

3. An equation similar to (3) is derived by McCallum using a model where the IS function depends on the real interest rate and output is constant. Such an equation is also implied by Barro's (1981, Chapter 2) model in which both the demand and supply of output in each local market are a stochastic function of a locally perceived real rate of interest.

4. For example, if a Barro (1981, Chapter 2) type model underlies the determination of δ_t, and α_t(and/or e_t) depends positively on output, then a positive economy-wide shock to the supply of output would decrease δ_t but increase α_t(and/or e_t) via its effect on output. Both of these variables would increase, however, in the case of a positive economy-wide shock to the demand for output.

5. It may be argued that a change in the average inflation rate may cause financial innovations which would alter the behavior of $\bar{\alpha}_t$. However, our assumption below that γ_2 is the same in the two regimes implies that the average inflation rates would also tend to be the same over long periods (e.g., see Figure 2).

6. The Bank first announced a monetary target for M3 in 1976 and then shifted in 1977 to a target for sterling M3 that excludes non-sterling balances from M3.

7. The base month was April for each year from 1976 through 1978, June for 1979 and February for subsequent years. The announced target ranges were as follows:

Year	Target	Year	Target	Year	Target
1976	9.0-13.0	1977	9.0-13.0	1978	8.0-12.0
1979	8.0-12.0	1980	7.0-11.0	1981	6.0-10.0
1982	8.0-12.0	1983	7.0-11.0	1984	6.0-10.0
1985	5.0- 9.0				

8. To affect the Treasury bill rate, the Bank has used its short-term interest rate (originally the Bank Rate, subsequently the Minimum Lending Rate and recently the clearing rate for bills [see Walters (1986), p. 115]).

9. The source of both series is Bank of England, Quarterly Bulletin. The series on£M3 is seasonally adjusted by the Bank. Quarterly data represent averages of monthly data.

10. The source of both series is OECD, Main Economic Indicators. The series on the GDP deflator is seasonally adjusted.

11. With four lags used in these tests as well as the VAR model estimated below, the first observation for the data is 1976:2 which represents roughly the starting date for announced targets in the U.K.

12. For Δp, the standard error of ρ equals .216. Thus, for example, the t-value for the hypothesis that ρ = .3 would be 1.64. Also note that since the standard error of ρ is high in the case of Δm as well, the test of stationarity also does not have much power for this series.

13. Since $\bar{\alpha}_t + e'_t = m_t - p_t + \beta r_t$, according to (20), stationarity of $\bar{\alpha}_t$ would imply that m, p and r are cointegrated.

14. Moreover the n-period forecast variance simply equals n times the one-period forecast variance in this case.

15. Note that according to (15), the ratio of the variance of $\Delta \bar{p}_t$ to that of Δp_t will lie between zero and one if the covariance between $\Delta \bar{p}_t$ and Δu_{2t} is zero. This ratio can exceed one if, as in the case of our model, the covariance between $\Delta \bar{p}_t$ and Δu_{2t} is negative, and two times the absolute value of the covariance is greater than the variance of Δu_{2t}.

16. It may also be of interest to examine the size of the permanent component in £M3. Using the estimate of the variance of $\Delta \bar{m}$ derived from our VAR model, we find that this variance is 1.2 times the sample variance of Δm.

Table 1

Tests of Stationarity and Cointegration

A. Univariate Series

Series	ρ [$\tau(\rho)$]	q(1,0) [p-value(%)]
m	.865 [-1.498]	-5.418 [79.50]
p	.948 [-1.631]	-2.825 [94.50]
r	.778 [-2.211]	-6.086 [73.75]
Δm	.186 [-2.978]	-24.195 [2.75]
Δp	.654 [-1.605]	-15.377 [17.00]
Δr	-.077 [-4.107]	-28.385 [1.00]

B. Multivariate Series

	q(3,2) [p-value(%)]	q(3,1) [p-value(%)]
m, p and r	-12.159 [91.00]	-1.929 [99.75]

NOTE: ρ is the coefficient on the lagged value of the dependent variable in a regression that also includes 3 lagged first differences of the dependent variable, a constant and a time trend. $\tau(\rho)$ is the Dickey-Fuller (1979) Statistic that tests the null hypothesis that $\rho = 1$. According to the distribution tabulated by Fuller (1976), the critical value (corresponding to approximately the same degrees of freedom as in our test) for $\tau(\rho)$ is -3.24 at the 10%, and -3.60 at the 5% level. q(1,0) is the Stock-Watson Statistic (forthcoming) that tests the null hypothesis of one unit root against the alternative of no unit root; q(3,2) and q(3,1) are the Stock-Watson Statistics testing for 3 unit roots against the alternatives of 2 and 1 unit roots. All of these statistics use linear detrending and four lags.

Table 2

Selected Results from the VAR Model

A. Correlation/Covariance Matrix of VAR Residuals

	ε_1	ε_2	ε_3
ε_1	$.149 * 10^{-3}$	-.069	-.336
ε_2	$-.634 * 10^{-5}$	$.573 * 10^{-4}$	.411
ε_3	$-.459 * 10^{-4}$	$.348 * 10^{-4}$	$.125 * 10^{-3}$

B. Selected Long-Run Multipliers

		The Response to a Unit Shock Summed Over			
Variable	Innovation	10 Quarters	20 Quarters	30 Quarters	40 Quarters
Δm	ε_1	.709	.494	.486	.489
	ε_2	1.189	1.749	1.749	1.743
	ε_3	.679	.780	.792	.790
Δp	ε_1	-.857	-1.026	-1.022	-1.020
	ε_2	3.097	3.434	3.421	3.418
	ε_3	.949	1.077	1.074	1.073

NOTE: Innovations, ε_1, ε_2 and ε_3 represent, respectively, residuals in the equations explaining Δm, Δp and Δr. In panel A, the values above the diagonal represent correlation coefficients while those on and below the diagonal represent variances and covariances. In panel B, values represent responses to shocks equal to 1.0 in the case of each innovation.

Table 3

Estimates of Selected Variances

		Value
1.	The variance of $\Delta\bar{p}$ with base drift	$.13695 * 10^{-2}$
2.	The variance of $\Delta\bar{p}$ without base drift	$.06143 * 10^{-2}$
3.	Row 2 divided by Row 1	.44852
4.	The variance of Δp with base drift	$.01906 * 10^{-2}$
5.	Row 1 divided by Row 4	7.1856

FIGURE 1

The Behavior of £M3 (in logarithms) Compared to Target Levels With and Without Base Drift

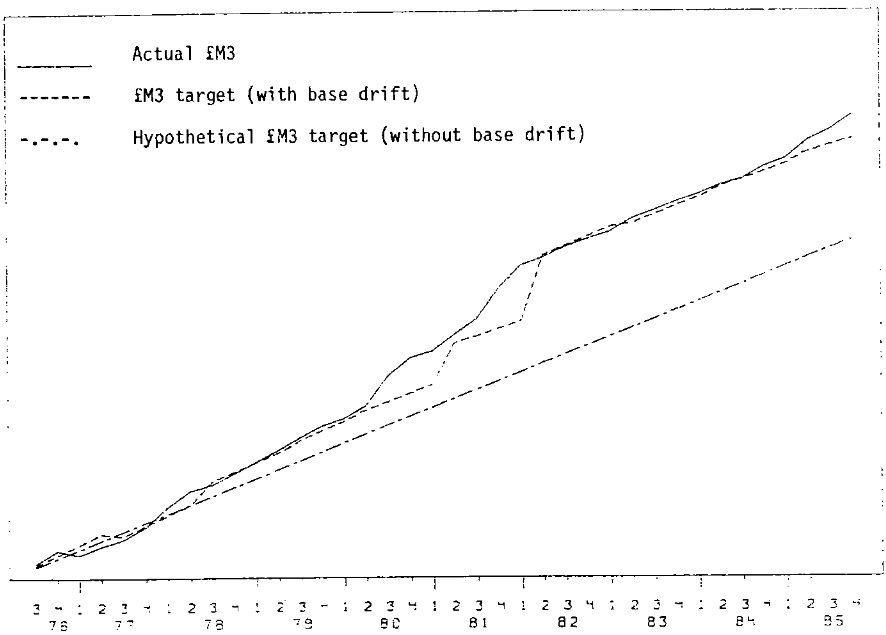
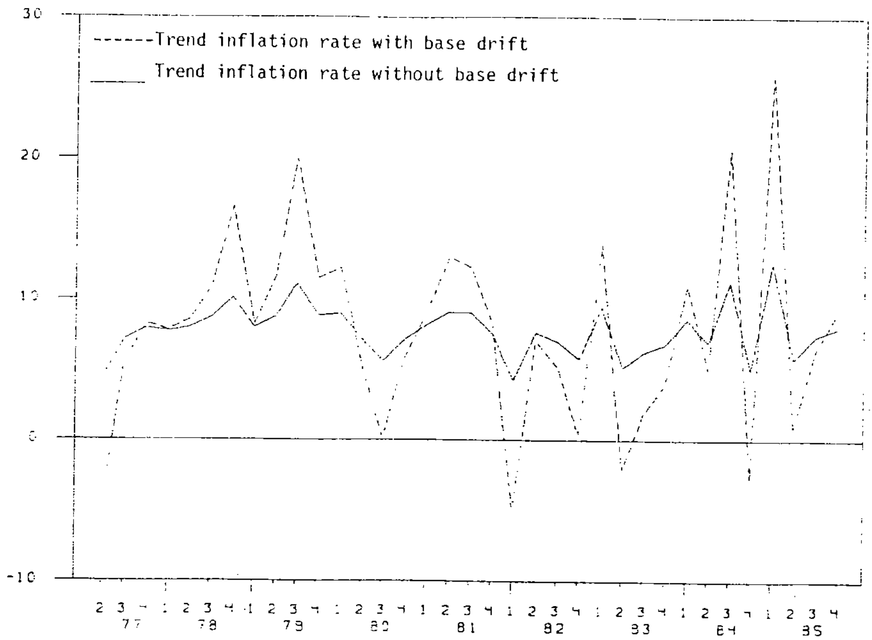


FIGURE 2

Trend Inflation Rates (Percent Per Year) With and Without Base Drift



References

Barro, Robert J., Money, Expectations and Business Cycles: Essays in Macroeconomics (New York: Academic Press), 1981.

Beveridge, Stephen, and Charles R. Nelson, "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to Measurement of the 'Business Cycle'", Journal of Monetary Economics 7 (1981), 151-74.

Broaddus, Alfred, and Marvin Goodfriend, "Base Drift and the Longer Run Growth of M1: Experience from a Decade of Monetary Targeting", Federal Reserve Bank of Richmond, Economic Review 70 (November/December 1984), 3-14.

Cochrane, John H., "How Big is the Random Walk in GNP?", Journal of Political Economy 96 (October 1988), 893-920.

Dickey, D.A., and W.A. Fuller, "Distribution of the Estimators for Auto Regressive Time Series with a Unit Root", Journal of the American Statistical Association 74 (1979), 427-31.

Friedman, Milton, "Monetary Policy: Theory and Practice", Journal of Money, Credit and Banking 85 (February 1982), 191-205.

Fuller, Wayne A., Introduction to Statistical Time Series (New York: Wiley), 1976.

Goodfriend, Marvin, "Interest Rate Smoothing and Price Level Trend-Stationarity", Journal of Monetary Economics 19 (May 1987), 335-348.

Goodhart, Charles A.E., Monetary Theory and Practice (London), 1983.

Huizinga, John, "An Empirical Investigation of the Long-Run Behavior of Real Exchange Rates", Carnegie Rochester Conference Series on Public Policy 27 (Autumn 1987).

King, Robert, Charles Plosser, James Stock and Mark Watson, "Stochastic Trends and Economic Fluctuations", NBER Working Paper No. 2229 (April 1987).

Lucas, Robert E., "Econometric Policy Evaluation: A Critique", Carnegie Rochester Conference Series on Public Policy 1 (1976), 19-46.

McCallum, Bennett T., "Price Level Determinacy with an Interest Rate Policy Rule and Rational Expectations", Journal of Monetary Economics 8 (1981), 319-29.

Poole, William, "Interpreting the Fed's Monetary Targets", Brookings Papers on Economic Activity (1976 No. 1) 247-59.

Shadow Open Market Committee, "Policy Statement", Center for Research in Government Policy and Business, Graduate School of Management, University of Rochester, March 1985.

Stock, James, and Mark Watson, "Testing for Common Trends", Journal of the American Statistical Association (forthcoming).

Walsh, Carl E., "In Defence of Base Drift", American Economic Review" 76 (September 1986), 692-700.

Walsh, Carl E., "The Impact of Monetary Targeting in the United States: 1976-84", NBER Working Paper no. 2384, (September 1987).

Walters, Alan A., Britain's Economic Renaissance: Margaret Thatcher's Reforms 1979-1984 (Oxford: Oxford University Press), 1986.

U.S. Council of Economic Advisers, Economic Report of the President (Washington), 1985.