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ABSTRACT

This paper investigates the performance, in several small-scale models of the Japanese economy, of an operational monetary policy rule related to ones previously considered for the United States. The rule dictates settings of the monetary base that are designed to produce values of nominal GNP close to targets that grow smoothly at a noninflationary rate. Simulations with quarterly data for the period 1972-1992 yield predominantly favorable results. Experiments with an interest rate instrument are also conducted but the simulated performance is less desirable. One section discusses issues concerning monetary base control in Japan.

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I. Introduction

Over the span of years between 1975 and 1990, the Bank of Japan was evidently more successful than most of its counterparts in the United States and Europe in conducting a monetary policy so as to avoid inflation and severe cyclical fluctuations. A recession was encountered in the early 1990s, however, and it is reasonable to speculate that as financial liberalization continues there may be increased difficulties with traditional monetary targets and indicators. Accordingly, this paper reports an application to the Japanese economy of a specific rule for the conduct of monetary policy that has previously been developed and studied in the U.S. context.

The basic objective of the rule in question, which has been analyzed in a series of papers by the author (McCallum 1988, 1990, 1991),[1] is to generate a time path for aggregate nominal spending that grows smoothly at a noninflationary rate. Nominal GNP is the measure used so far, but other nominal spending variables might offer practical advantages without departing from the basic rationale for the rule's design. An important element of this design is that the rule should be operational, i.e., be one that presumes manipulation of an instrument variable that is actually controllable by the central bank and that relies upon information that is actually or potentially available. The presumed instrument in previous studies is the monetary base, a variable that appears on any central bank's own balance sheet and is therefore subject to constant observation and adjustment. In practice, however, the Bank of Japan (BOJ)--like the Federal Reserve in the United States--has never used the monetary base as an operating instrument. Instead, the BOJ has relied primarily on manipulation of short term interest rates as its means of implementing policy. Indeed, it is occasionally suggested that the BOJ could not control the monetary base on a short-term basis even if it desired to do so. Consequently, recognition and discussion of existing procedures and institutions is of particular importance for the present study. In addition, some space will be given to an investigation of the apparent effectiveness of a rule that uses an interest rate instrument in conjunction with the same nominal spending targets as in the primary study.

One substantial modification of the present study, in relation to those conducted previously, is a rather extensive exploration of noninflationary GNP targets expressed in terms of growth rates rather than by a single, preset growth path. The resulting modification of the policy rule offers several attractive features, including reduced volatility of the instrument variable and the associated possibility of stronger feedback responses, as well as (arguably) more appropriate responses to typical, long-lasting shocks. Another modification involves increased attention to open-economy considerations in several of the models utilized.[2]

The outline of the presentation is as follows. In Section II, the rule's design is rationalized and previous applications to the U.S. economy are briefly reviewed. Then in Section III issues relating to the use of the monetary base as an instrument are discussed. The first results for the Japanese economy, based on simulations with very simple atheoretic models of nominal GNP determination, are presented in Section IV. Next, several vector autoregression models are used in Section V. Some slightly more elaborate models, intended to represent simple structural representations of alternative theories of business cycle behavior, are then developed and simulated in Section VI. Some experiments involving an interest rate instrument are described in Section VII and, finally, conclusions are presented in Section VIII. Appendices are devoted to description of the data utilized and presentation of the adjusted monetary base series.

II. Review of Previous Studies

Let us begin by reviewing the policy rule utilized in previous studies, providing an explanation of its rationale, and briefly summarizing results obtained for the U.S. economy. As mentioned above, the rule is one that dictates settings of the monetary base that are designed to keep nominal GNP (or some other measure of nominal spending) close to a target path that grows smoothly and steadily at a noninflationary rate--at the long-term average rate of growth of real output. In my previous work, I have taken this rate to be three percent per year, used GNP as the relevant measure of spending, and worked with quarterly data. With b_t denoting the log of the monetary base (averaged over quarter t), x_t denoting the log of nominal GNP, and x_t^* being the latter's target value, the rule can be written as

$$(1) \quad \Delta b_t = 0.00739 - (1/16)[x_{t-1} - b_{t-1} - x_{t-17} + b_{t-17}] + \lambda(x_{t-1}^* - x_{t-1}).$$

Here the constant term is simply a three percent annual growth rate expressed in quarterly logarithmic units, while the second term subtracts the average growth rate of base velocity over the previous four years. Finally, with λ set to a value such as 0.1 or 0.25, the third term adds a gentle adjustment in response to cyclical departures of nominal GNP from its target path. In most of my previous work this path has been specified as $x_t^* = x_{t-1}^* + 0.00739$ with an initial value equal to actual x_t in the last quarter before the period under study. This gives a single, preset time path that grows at a constant rate of three percent per year. An alternative target specification will be discussed below.

One obvious feature of the foregoing rule is that it uses nominal spending as the monetary authority's principal target variable, rather than a monetary aggregate such as M1 or M2. There are several reasons for preferring a nominal spending target to the monetary aggregates that have traditionally received more attention in both theoretical and practical literature. First, the average rate of nominal spending growth necessary to yield a desired average inflation rate over extended periods of time can be more accurately determined. Thus, for example, it is highly probable that the average growth of real output over the next 20 years in a nation such as the United States (or Japan) will be within one percentage point of 2.5 percent (or 4 percent) per annum, so achievement of that rate of growth of nominal GNP will result in approximately zero inflation. Considerable uncertainty exists, by contrast, as to the average growth rate of M1 or M2 velocity, and therefore to the average growth rate of M1 or M2 that would yield zero inflation.

Second, the maintenance of a steady growth rate for nominal income has better automatic stabilization properties in response to some types of shocks.[3] If these shocks are predominant, better cyclical behavior of the economy should result from an arrangement that stabilizes nominal income rather than money around a smooth path. And, third, regulatory change and technological innovation in the payments industry require occasional revisions in operational measures of monetary aggregates. It is possible, consequently, that any chosen measure will be less reliably related to instrument values than would nominal spending.

Another feature of the rule is the specification of a constant growth target for nominal GNP, rather than a target rate that varies over the cycle. The reason for this feature is that in practice a central bank cannot control, or predict with any accuracy, how nominal GNP growth will be split on a quarter-by-quarter basis between real growth and inflation. And academic economists can--as mentioned in footnote 3--do no better in that regard. Thus it seems best to avoid attempts at fine tuning, instead being satisfied to smooth out fluctuations in nominal spending in the hope that such an achievement would reduce fluctuations in real magnitudes. It would at least eliminate policy surprises as a source of undesirable fluctuations.

Still another feature of rule (1) is that it specifies use of the monetary base as the policy instrument (or "operating variable," in the language of Suzuki (1986)). In that regard the rule is desirably operational, in the sense that it specifies settings for an instrument variable that the central bank is capable of controlling with accuracy.[4] It could also do so with an interest rate instrument, but my initial presumption is that a quantity variable would be preferable because of the well-known ambiguity of interest rates as indicators of policy stance.[5] But interest rates are preferred by most actual central banks, so a preliminary investigation of an interest-rate rule analogous to (1) will be undertaken below in Section VII.

A fundamental issue is whether nominal income targeting is actually feasible, i.e., whether nominal GNP targets can be accurately achieved by control of the base or some other instrument available to the central bank.[6] To investigate that issue in the U.S. context was the main purpose of a fairly extensive study (McCallum, 1988), which yielded highly encouraging results. The next few paragraphs will briefly review the design and outcome of that study.

To determine whether policy rule (1) would in fact keep nominal GNP close to its steady growth path, given the existence of stochastic shocks of various types, the researcher[7] needs to conduct simulations incorporating such shocks in a system that includes the rule and an econometric model that describes the response of x_t to the generated values of b_t. The fundamental problem in this regard is that there is no agreed-upon model. As mentioned in footnote 3, the macroeconomics profession does not possess a satisfactory model of the short-run, dynamic behavior of aggregate supply that governs the response of real variables to monetary policy actions--not even at the qualitative level. In light of this problem, my preferred method of investigation has been to determine whether policy rule (1) will perform reasonably well in a variety of different models. Thus in my 1988 study I conducted simulations with two single-equation atheoretic specifications, several vector autoregression (VAR) systems, and finally three models that were intended to be structural (i.e., policy invariant). These latter models are quite small in scale but were designed to represent leading alternative theories of business cycle dynamics--specifically, the "real business cycle" (RBC) theory of Kydland and Prescott (1982), the monetary-misperceptions theory of Lucas and Barro, and a more Keynesian theory (PC) patterned on the Phillips-curve and price-adjustment specifications of the Federal Reserve's quarterly MPS model.

The results for the U.S. economy, in counterfactual simulations pertaining to the period 1954.1-1985.4, are summarized in Table 1.[8] The entries in this table are root-mean-square errors (RMSE)--i.e., deviations from the target path--in simulations with systems including rule (1) and the five models indicated. In each case the simulation begins with initial conditions prevailing at the start of 1954 and continues with shocks fed into the system in each quarter, these shocks being estimated as residuals from the relations estimated in the respective models. It will be seen that for λ values in the range of 0.1 to above 0.25, the RMSE values are about 0.02 (i.e., two percent) with all five models.[9] Thus performance is satisfactory in all of these cases, and distinctly superior to that with no feedback (i.e., with $\lambda = 0$). Higher values of λ give rise to the possibility of dynamic instability--i.e., explosive oscillations--which occurs with $\lambda = 0.5$ in the VAR system (and with $\lambda = 1.0$ in the other systems). But with moderate

Table 1

Basic Results for U.S. Economy, 1954-1985

RMSE Values with Five Models

Model	Value of λ in Rule (1) 0.00	0.10	0.25	0.50
Single Equation	0.0488	0.0249	0.0197	0.0162
4-Variable VAR	0.0479	0.0216	0.0220	0.1656
Real business cycle	0.0281	0.0200	0.0160	0.0132
Monetary misperceptions	0.0238	0.0194	0.0161	0.0137
Phillips curve	0.0311	0.0236	0.0191	0.0174

Table 2

Additional Results for U.S. Economy, 1954-1985

Results with x_t^{**} Target Value and $\lambda = 0.25$

Model	RMSE relative to x_t^{**}	RMSE relative to x_t^{*}	Standard deviation of Δb_t	Standard deviation of Δb_t using x_t^{*} Target
Single Equation	.0102	.0400	.0036	.0063
4-Variable VAR	.0102	.0394	.0036	.0069
Real business cycle	.0107	.0229	.0040	.0054
Monetary misperception	.0113	.0196	.0037	.0051
Phillips curve	.0100	.0259	.0042	.0066

values of λ, the rule succeeds in generating paths of x_t that are noninflationary and, in addition, somewhat smoother than those that have obtained historically. A plot of x_t and the target path x_t^* for the VAR model and $\lambda = 0.25$ is shown in Figure 1.

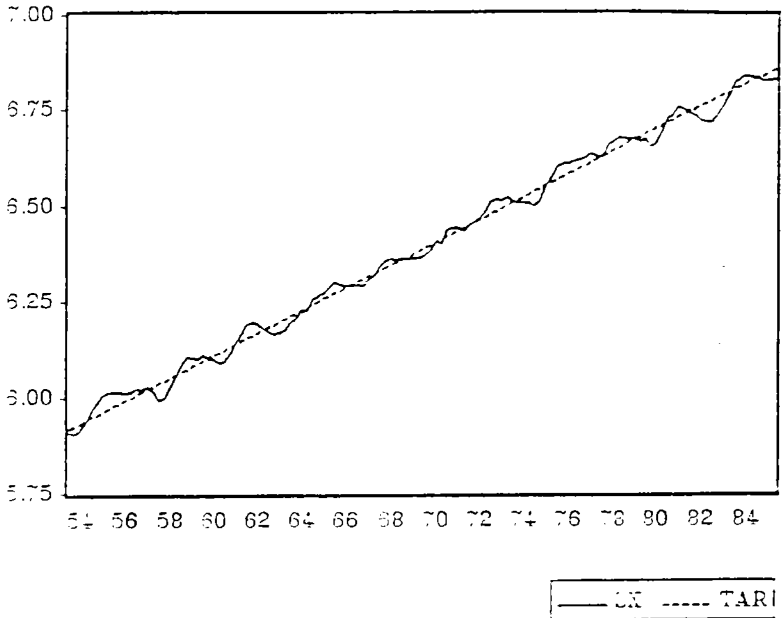
The foregoing results were developed in my 1988 paper; additional findings are reported in McCallum (1990a) and (1991). In the first of these it was found that substitution of an explicit price level target, rather than nominal GNP, is somewhat less satisfactory since this change increases the likelihood of dynamic instability. Also, a few experiments with an interest rate instrument were attempted; these will be built upon in Section VII. In the 1991 paper, by contrast, the purpose was to determine whether adherence to rule (1) would have prevented the Great Depression of the 1930s. Counterfactual historical simulations for 1923-1941 were conducted with a small model of GNP determination, estimated with quarterly U.S. data for 1922-1941. The simulation results indicate that nominal GNP would have been kept reasonably close to a steady three percent growth path over 1923-1941 if the rule had been in effect, in which case it seems highly unlikely that real output and employment would have collapsed, as they did in fact.

More recently, however, I have come to believe that a strong case can be made for expressing the nominal GNP target in terms of growth rates rather than levels corresponding to a single predetermined growth path. The main reason is that, because real shocks that affect the economy's natural-rate output level are highly persistent, it may be undesirable to quickly drive x_t values back to the predetermined x_t^* path after shocks have occurred. Instead, it would seem to be preferable to treat past shocks as bygones, which could be accomplished by adopting $x_t^{**} = x_{t-1} + 0.00739$, rather than $x_t^* = x_{t-1}^* + 0.00739$, as the target value for period t. This sort of growth rate target has recently been favored by several economists, including Feldstein

Figure 1

Simulation for United States with Rule (1),

VAR Model, and $\lambda = 0.25$: 1954-1985



LX: log of nominal GNP, simulated

TAR: target path

and Stock (1993). Such a change would admittedly result in a nominal GNP path that has a unit-root component--indeed, that is close to a random walk with drift. But if the drift magnitude were 0.00739--or whatever is the average rate of output growth--then expected inflation over any horizon would be zero. Furthermore, price level variability over practical planning horizons would not be excessive if the extent of single-period variability--i.e., the variability of $x_t - x_t^{**}$ -- were small.

This suggestion for using growth rate targets merits consideration for two additional reasons besides the one just mentioned. First, it seems likely that instrument variability--variability of Δb_t values in the case at hand--should be reduced for any given values for the feedback coefficient λ. That is a significant reason because one of the main objections to rule (1) that has been expressed by central-bank economists is that it is likely to call for more variability of instrument settings than has prevailed historically. Second, it should accordingly be possible to use larger λ values, implying stronger feedback, without inducing instrument instability.

Consequently, in an unpublished paper dated October 1990, I conducted some preliminary studies using a modified version of rule (1) that substitutes x_t^{**} for x_t^{*}. The results, summarized in Table 2 for the VAR case with $\lambda = 0.25$, are highly encouraging. In particular, the RSME values relative to the target x_t^{**} are only about 0.01 while the variability of the Δb_t instrument is reduced considerably relative to its magnitude with the x_t^{*} target. Accordingly, a prominent role will be given in Sections IV and V below to results for the Japanese economy obtained with the modified version of rule (1) that utilizes the x_t^{**} target.

III. The Base as a Potential Instrument

We now turn to the issue of the instrument variable. It is well known, as mentioned above, that the BOJ has never used the monetary base (or any other closely-related narrow aggregate) as its instrument variable. Indeed, among the central banks of the industrialized world, the Swiss National Bank is perhaps the only one currently or recently to use a base-type instrument. That fact is not necessarily a first-order problem for the present study, however, for its basic purpose is to estimate how the evolution of nominal GNP would have differed from the historical record _if_ policy had been conducted _differently_--specifically, if it had conformed to rule (1).[10]

Such an exercise would be of limited interest, however, if there were no logical possibility of conducting policy with a base instrument. Thus there is a need to respond to the suggestion, mentioned by Okina (1991) and Ueda (1991), that the BOJ could not control the monetary base on a short-term basis even if it tried to do so. In this regard it will readily be admitted that any attempt to tightly control base values on a (say) weekly basis would lead to some increase in weekly variability of short-term interest rates. But the suggestion at issue is evidently more substantial than that. What is emphasized in the relevant literature is that legal reserve requirements in Japan are of the lagged reserve accounting type[11] and that excess reserve holdings are minuscule.[12] Accordingly, reserve demand at the end of each reserve maintenance period is virtually predetermined. If the BOJ were to fail to provide the stipulated reserves, therefore, some banks would necessarily violate their legal requirements. So the BOJ cannot use the base as its instrument, according to this point of view.

The basic reply, of course, is that reserve requirements could be changed so as to be of the contemporaneous type. And, indeed, such a change would probably be warranted if the BOJ were to adopt a base instrument.[13] But even with a continuation of lagged reserve requirements, it is conceivable that the system could adjust to a regime featuring stringent base control. One adjustment that would occur naturally, i.e., via the self-interested behavior of privately motivated banks, is that higher levels of excess reserves would be held as a matter of course and managed in an interest-elastic fashion.[14] With excess reserves of 2-3% of required reserves--still only a tiny fraction of deposits--banks would be able to avoid violation of the legal requirements except in highly unusual circumstances.[15] Indeed, a major reason why excess reserve holdings are currently so small is that the BOJ has routinely supplied or removed reserves at the end of each maintenance period so as to smooth interest rates--i.e., to keep them from rising or falling sharply.[16]

In principle, then, it would be possible for the BOJ to use the monetary base as its operating variable. And accurate attainment of the base values stipulated by rule (1) could be combined with some interest rate smoothing on a daily basis. One possibility is that an interest rate could be adopted as the variable manipulated on a day-to-day basis but with target values set according to another rule designed to yield quarterly-average values of the base that conform to rule (1). But increased interest rate variability at the monthly or quarterly interval would then be probable. The question ultimately at issue is what combination of target and instrument variables is optimal. This question cannot be answered solely on the basis of macroeconomic performance, since a central bank has responsibilities for the stability of the financial system as well as macroeconomic performance (i.e., avoidance of inflation and cyclical fluctuations). But it also cannot be answered without attention to macroeconomic performance, which is the topic of the present paper.

From the purely macroeconomic perspective, it seems likely a priori that a monetary base instrument would be better than an interest rate instrument because of the ambiguity of nominal interest rates as indicators of monetary tightness or ease. High interest rates, that is, are associated with tight monetary policy from a short-term perspective but are associated with easy monetary policy over longer horizons. This implies that the interest rate effects of a monetary tightening--e.g., an open-market sale of securities--are in opposite directions from the short-term and long-term perspectives. Accordingly, the design of a policy rule for the control of nominal spending would appear to be more delicate and difficult if the instrument variable is an interest rate than if it is a quantity variable. And the base is the most natural quantity variable to select as a instrument because it provides a summary of the impact of the central bank's monetary operations. Furthermore, the base is controllable with a fairly high degree of accuracy since it is the sum of items that appear on the central bank's own balance sheet. Its magnitude can therefore be monitored daily and adjusted with open-market sales or purchases if the intended value does not prevail.

In Japan, as elsewhere, the quantity of currency in circulation is demand-determined in the sense that deposits can be redeemed in currency at the wish of the deposit holder. Some readers have argued that this implies that the base--the sum of currency and reserves--would be inferior to reserves alone as an instrument variable. But reflection suggests the contrary. Aggregate spending depends positively on both components, since reserves and deposits are closely related. So spending is apt to be more strongly related to the sum of the two components than to either of them alone, in which case better control of spending would be provided via manipulation of the base than of reserves alone, even though currency is being left to the public's choice. Still better control might be afforded by some other linear combination of reserves and currency, rather than their sum, but the criterion of simplicity indicates that initial studies should concentrate on the base.

But despite this a priori case for the base as an instrument variable, the present study will provide some less extensive evidence relating to the suitability of a short term interest rate. Specifically, Section VII will report results of simulations in which a rule analogous to (1), but dictating settings for the three-month bill rate, is utilized. Such evidence should be helpful in reaching a conclusion as to the relative merits of base and interest rate instruments from the macroeconomic perspective.

In empirical work with U.S. data, it is standard practice to use a measure of the monetary base that has been adjusted to take account of changes in the schedule of reserve requirements.[17] No such series is published in Japan, evidently, but it is quite important that adjustments be made for changes in the BOJ's Reserve Requirement Ratios, despite the fact that these ratios are kept rather low. The reason is that quite a few changes have been made, over the period to be studied, that are large in relative terms. In October of 1991, for example, the ratio for "other deposits" (in banks with deposits of over ¥ 2.5 trillion) was reduced from 2.5 percent to 1.3 percent--reduced to just over half of the earlier figure. Accordingly, the quantity of reserves held by the deposit banks fell sharply between August and November. But since the BOJ was smoothing interest rates, this change did not represent a sharp tightening of policy and would not be reflected in a well-designed policy measure. Thus the design of the measure to be used below is as follows. Since the magnitude of deposits is approximately equal to the product of the (reserve requirements) ratio and the volume of reserves, an appropriate measure of adjusted reserves would be the raw value of reserves multiplied by an adjustment factor that is inversely proportional to the current value of the ratio. In the work that follows, this factor was scaled to equal 1.0 during the long period between April 1981 and October 1991 when no changes were made. The values used for the ratio at each point in time were those pertaining to "Other Deposits" at the largest-sized banks (the size categories changed over time, of course, as the economy grew). The adjusted reserves value was calculated for the end of each month and then averaged over the three months of each quarter to obtain a quarterly series. That series, finally, was seasonally adjusted in the same manner as with other series that are reported only in a seasonally unadjusted fashion. That procedure will be described in the following section. The adjusted base series is reported, together with several constituent series, in Appendix B.

IV. Preliminary Results

Now let us turn to our first set of results pertaining to the Japanese economy. For application to Japan, the policy rule given in (1) needs to be modified in one respect. In particular, the rule's constant term needs to be somewhat larger, reflecting a higher target growth rate for nominal GNP. One basic reason is that the long-term average growth rate of <u>real</u> output has been--and is expected to remain--higher in Japan than in the U.S., around 4 percent per year rather than 2 1/2 - 3 percent. But there is also a second reason for making the target path of nominal GNP steeper, which is that the BOJ may consider its long term inflation target to be somewhat above zero when expressed in terms of the GNP deflator. Accordingly, in the simulations that follow, the constant term has been set at 0.01468, which amounts to a 6 percent annual growth rate expressed in quarterly logarithmic units.[18] The policy rule to be used, then, is

$$(2) \quad \Delta b_t = 0.01468 - (1/16)[x_{t-1} - b_{t-1} - x_{t-17} + b_{t-17}] + \lambda(x^*_{t-1} - x_{t-1})$$

Here, as in Section II, b_t and x_t denote logarithms of the monetary base and nominal GNP.[19] For the moment, the target variable x^*_t is defined as $x^*_t = x^*_{t-1} + 0.01468$, reflecting a single preset path.

The other ingredient needed for each simulation is a model of Japanese GNP determination. As in my work with U.S. data, we begin with a set of preliminary results based on two versions of an atheoretic regression model that simply relates nominal GNP growth, Δx_t, to base money growth, Δb_t, and to some lagged values of these variables. The first version for the United States included only Δx_{t-1} and Δb_t as regressors--see McCallum (1988, pp. 178-9)--but in the case of Japan two lagged values of each variable are important in explaining movements in Δx_t. Thus the first model, which was estimated with seasonally-adjusted quarterly data for 1964.2-1992.4,[20] is as follows:

$$\text{(3)} \quad \Delta x_t = \underset{(.002)}{0.002} + \underset{(.086)}{0.040}\Delta x_t + \underset{(.087)}{0.228}\Delta x_{t-2}$$

$$+ \underset{(.063)}{0.223}\Delta b_t + \underset{(.066)}{0.135}\Delta b_{t-1} + \underset{(.068)}{0.207}\Delta b_{t-2} + e_{3t}$$

$$R^2 = 0.480 \qquad SE = 0.0117 \qquad DW = 2.12$$

Here, and in regressions reported below, the figures in parentheses are estimated standard errors, SE is the estimated standard deviation of the disturbance term, and DW is the Durbin-Watson statistic.[21] Also, e_{3t} denotes the estimated disturbance or shock realization for period t.

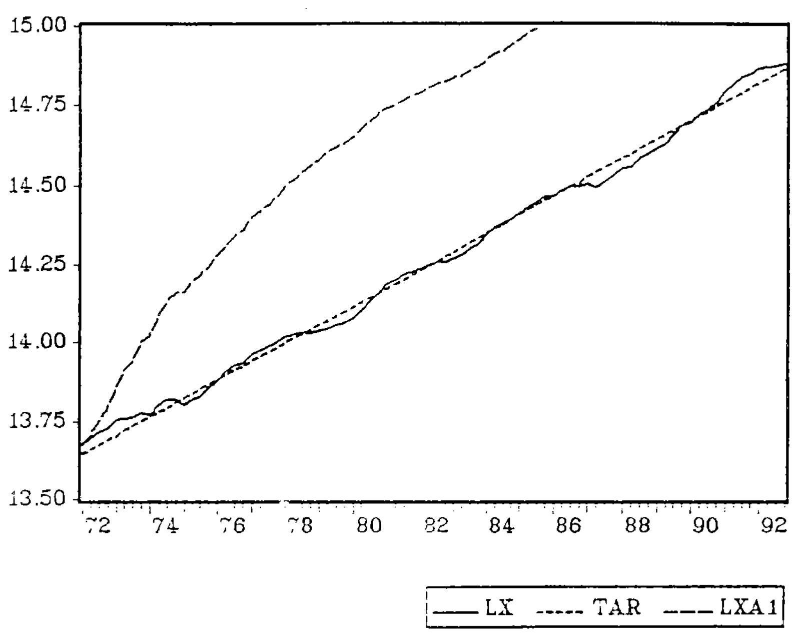
To carry out a counterfactual historical simulation we use equations (2) and (3) to generate 84 values for b_t and x_t, with initial values corresponding to those that actually obtaining in 1972.1 and with e_{3t} values fed into the system as estimates of the shocks that hit the economy over 1972-1992.[22] The result of one such simulation exercise with a λ value of 0.25 is shown in Figure 2, where TAR denotes the target path, LX is the simulated path for x_t, and LXA1 is the actual historical path of x_t. Algebraically, root-mean-square error (RMSE) magnitudes, (analogous to those in Table 1) are reported for a few values of λ in the first row of Table 3.

In terms of these RMSE values, the performance of rule (2) is slightly inferior to that obtained in the U.S. context, as can be seen by comparing Tables 1 and 3. The absolute level of performance is quite satisfactory, however, since the rule yields RMSE values far below the actual historical value of 0.4909 and well below the historical value relative to a fitted time trend of 0.0922. The latter value pertains only to the smoothness of the historical series and therefore does not reflect any penalty for an average

Figure 2

Simulation vs. Actual for Japan, 1972-1992

Model (3), Rule (2) with $\lambda = 0.25$



LX: log of nominal GNP, simulated

TAR: target path

LXA1: log of nominal GNP, actual

inflation rate in excess of the 2 percent inflation target value.[23] Graphical plots of the simulated x_t series for λ values of 0.0 and 0.5 are given in Figure 3, which shows that stronger feedback--i.e., higher values of λ--tends to produce smaller target misses. That effect cannot be taken to the extreme, however, since excessive values of λ will generate explosive oscillations reflecting so-called "instrument instability." Figure 4 shows that this type of instability obtained in the model at hand if λ is raised to the value 2.5.

The second single-equation atheoretic model differs from (3) in that the current-period value of Δb_t is deleted, the reason being that it seems likely that some of the response reflected in the 0.223 coefficient in (3) is due to simultaneous equation bias--i.e., to policy reactions within the quarter to macroeconomic conditions. With that change, the estimated relation becomes:

$$(4) \quad \Delta x_t = \underset{(.002)}{0.004} + \underset{(.091)}{0.074}\Delta x_{t-1} + \underset{(.089)}{0.302}\Delta x_{t-2} + \underset{(.069)}{0.146}\Delta b_{t-1} + \underset{(.070)}{0.254}\Delta b_{t-2} + e_{4t}$$

$$R^2 = 0.420 \qquad SE = 0.0123 \qquad DW = 2.19$$

Simulations for 1972.1 - 1992.4 based on model (4), conducted in the same fashion as described above, yield results as reported in the second row of Table 3 and shown for $\lambda = 0.25$ in Figure 5. The performance of the rule is not terrible, but is noticeably poorer than with model (3). The reason, of course, is that (4) implies that there is no current-period response of nominal income to base movements--i.e., that the average response lag is longer. That also increases the possibility of instrument instability: explosive oscillations occur in this second model with $\lambda = 1.0$.

The foregoing results all pertain, however, to cases with targets given by the single predetermined path for x^*_t, rather than unchanging growth rate

Table 3

Initial Results for Japan, 1972-1992

RMSE Values with Simplest Models

Model	Value of λ in Rule (2) 0.00	0.10	0.25	0.50
Equation (3)	0.0623	0.0316	0.0245	0.0199
Equation (4)	0.0789	0.0420	0.0324	0.0312

Table 4

Additional Results for Japan, 1972-1992

RMSE Values Relative to Target x_t^{**}

Model	Value of λ 0.00	0.25	0.50	1.00
Equation (3)	0.0099	0.0098	0.0098	0.0100
Equation (4)	0.0105	0.0104	0.0102	0.0103

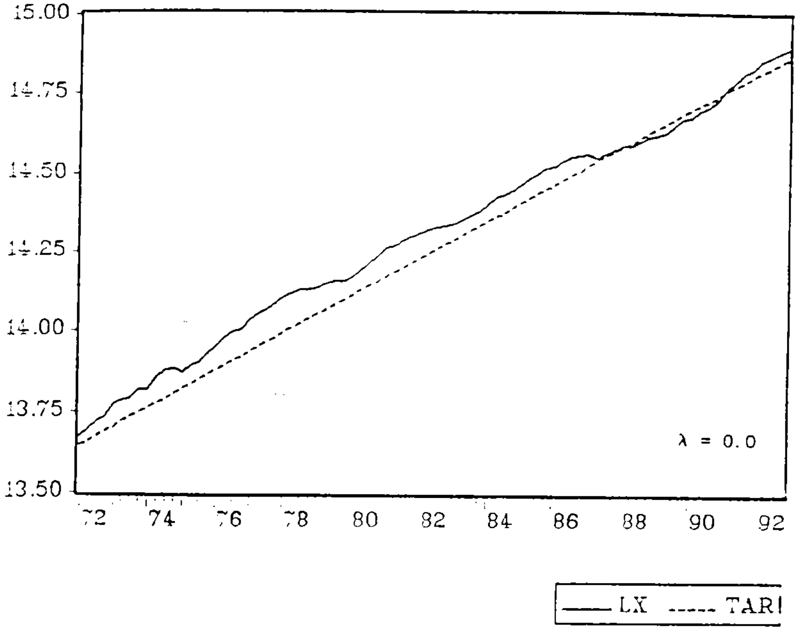
Table 5

Other Results for Japan, 1972-1992, with Model (4)

RMSE Values Using Target x_t^{*a}

Relative to:	Value of λ 0.00	0.25	0.50	1.00
x_t^{*a}	.0184	.0133	.0118	.0108
x_t^{**}	.0105	.0106	.0106	.0107
x_t^{*}	.0789	.0471	.0350	.0250

Figure 3
Results with Alternative values of λ
Model (3), Rule (2), 1972-1992



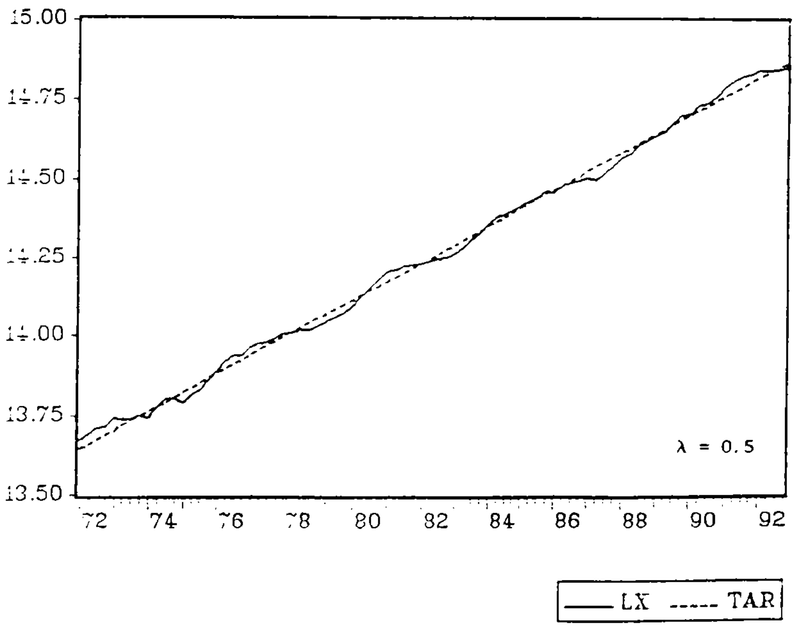


Figure 4

Results with Excessive Value of λ

Model (3), Rule (2), 1972-1992

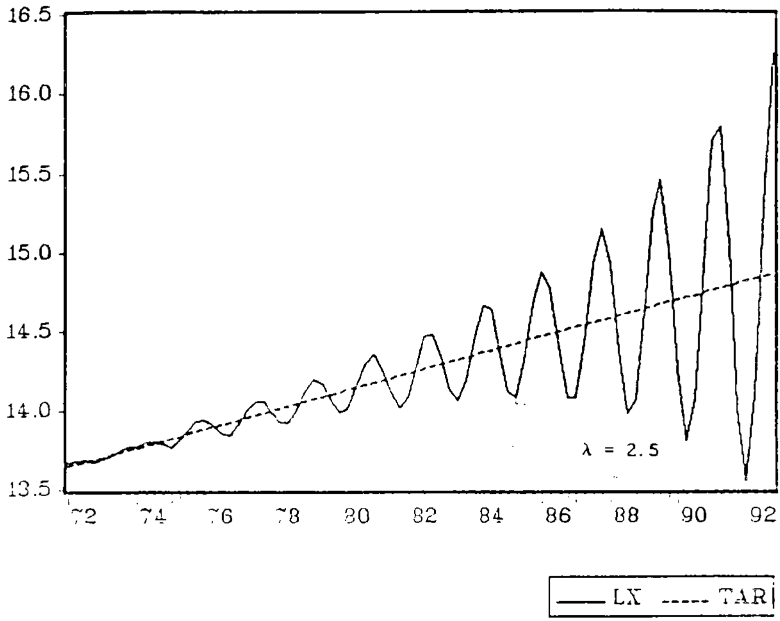
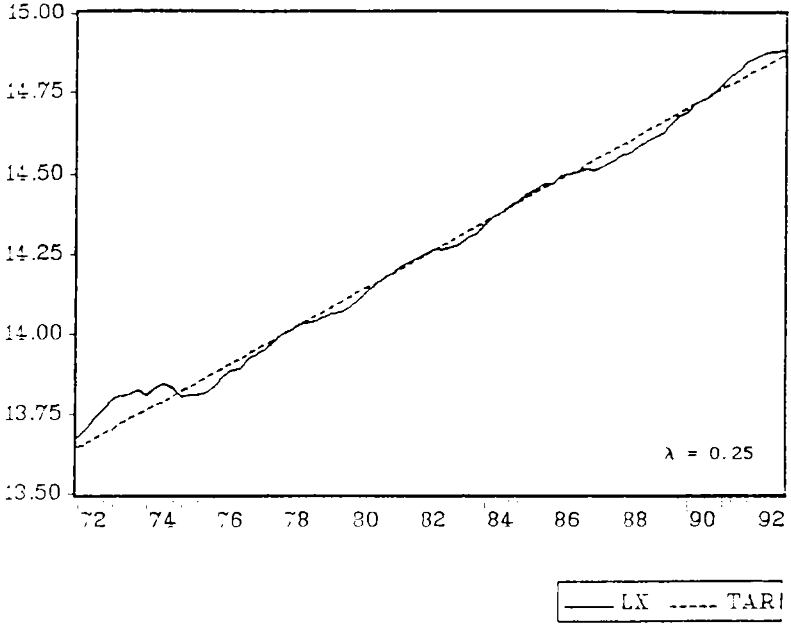


Figure 5

Results with Model (4)



targets. As explained in Section II, it is at least arguable that targets of the form $x_t^{**} = x_{t-1} + 0.01468$ are more appropriate than ones specified by x_t^*. Accordingly, results have been obtained for both models (3) and (4) using x_t^{**} targets, the RMSE values being reported in Table 4. There we see that Δx_t values are kept quite close to the 0.01468 target, variations around the latter having RMSE magnitudes close to 0.01. Since the actual historical RMSE for Δx_t relative to its mean value is 0.0138, there values indicate that the x_t^{**} targets lead to reduced growth rate variability, as well as values that would tend to reduce inflation to an insignificant magnitude. Figure 6 plots the simulated time paths (LX) for $\lambda = 0.25$ and $\lambda = 1.0$ together with the target paths x_t^{**} (TARM) and also, for reference purposes, x_t^* (TAR).

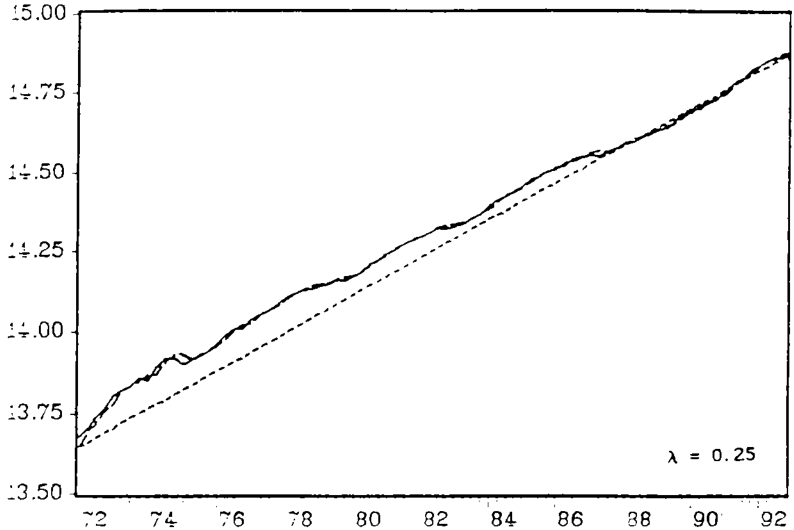
In addition to the foregoing results, there are also others that are favorable to the idea that x_t^{**} would be a desirable target. First, it is the case that variability of Δb_t instrument settings is reduced relative to that required with the x_t^* target; this is illustrated in Figure 7, where DLB denotes the simulated Δb_t path in each case and DLBA denotes the actual historical values. Second, it is possible to increase the value of the policy response parameter λ with a reduced danger of instrument instability. While $\lambda = 1.0$ generated explosive oscillations in model (4) with the x_t^* target, values as high as $\lambda = 3.0$ perform satisfactorily when the x_t^{**} target is used.

One objection that might be raised to use of the x_t^{**} growth rate target, instead of x_t^*, is that it permits x_t to drift away from x_t^*--which implies that the (log) price level is permitted to drift up or down over time. The plots in Figure 6 might seem to suggest that there is some sort of tendency for x_t to be driven back to x_t^* as time passes. But, unfortunately, the tendency that is indeed present in these counterfactual historical simulations is spurious; it results from the fact that the e_{4t} residuals in

Figure 6

Results with Target Variable x_t^{**}

Model (4)



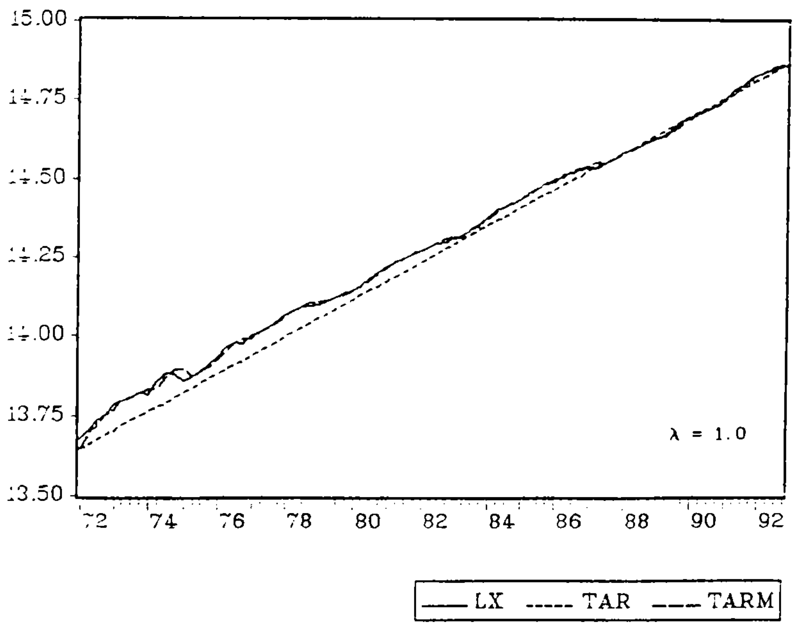
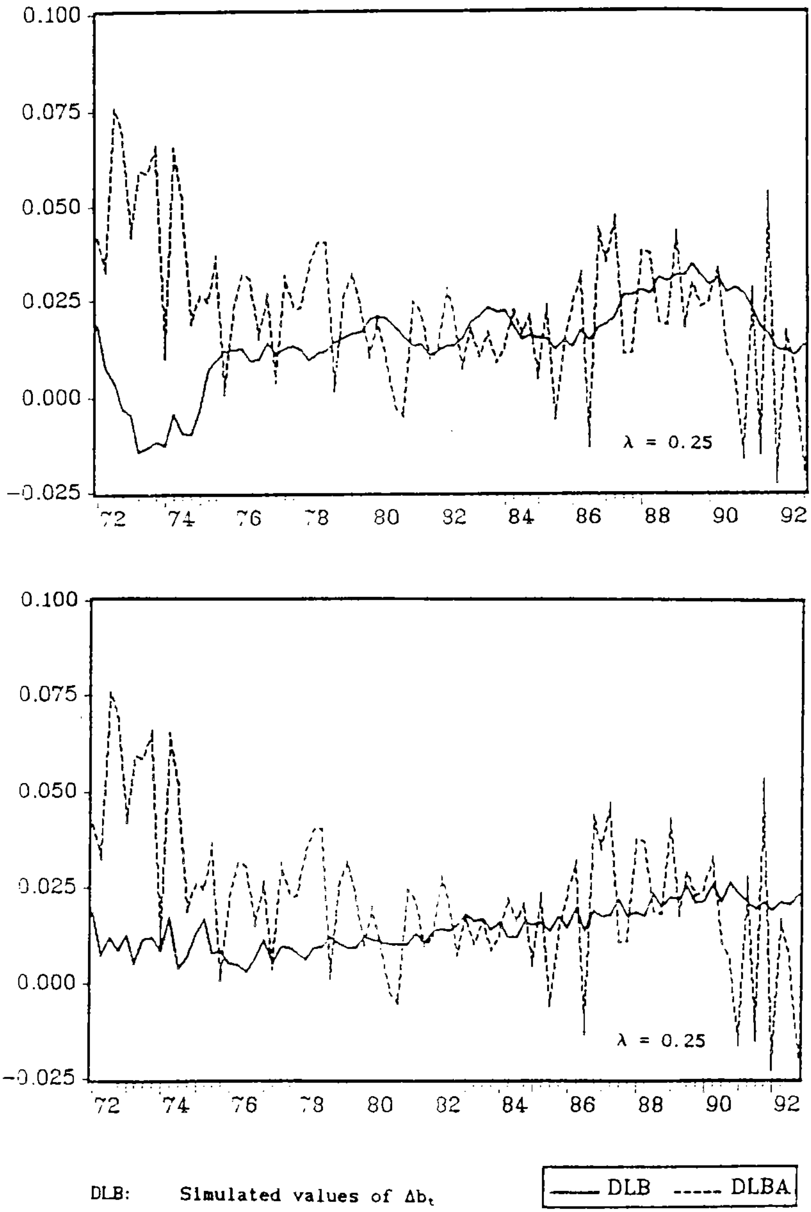


Figure 7
Instrument Variability with x_t^{*} and x_t^{**}
Targets, Respectively: Model (4)



DLB: Simulated values of Δb_t

DLBA: Actual values of Δb_t

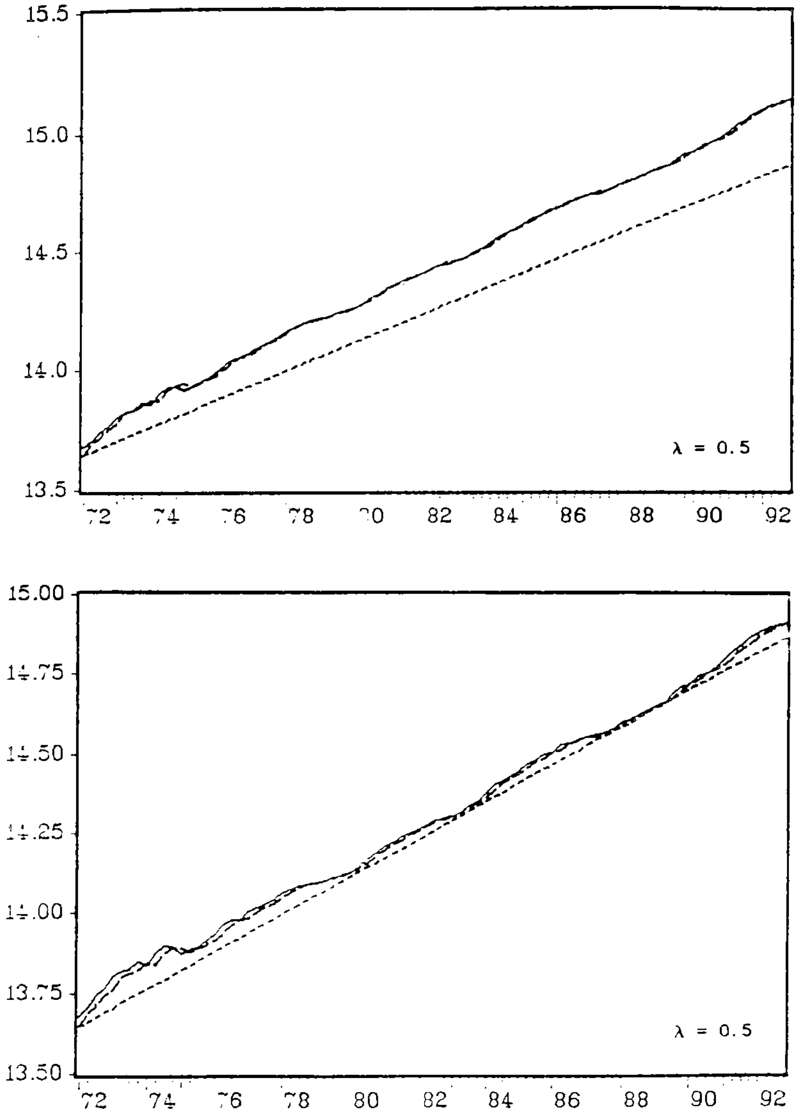
(4) are estimated by means of a regression procedure, which implies that their sum must equal zero.[24] But that would not be true in a proper stochastic simulation or in reality (over any finite time span).

So the rule with x_t^{**} targets will not prevent the price level from drifting away from its value at the time the rule is adopted. I have argued previously that such a tendency should not be considered to be a major problem, provided that the average drift magnitude is zero and the price level innovation term does not have a large variance. But it is an undesirable feature of the x_t^{**} target, nevertheless, so it seems worthwhile to consider a third target specification that is a weighted average of x_t^{*} and x_t^{**}. Accordingly, some relevant results are reported in Table 5 for a target variable defined as $x_t^{*a} = 0.2x_t^{*} + 0.8x_t^{**}$. There RMSE values for model (4) are reported not only relative to the x_t^{*a} values that are used as targets, but also relative to paths for x_t^{*} and x_t^{**}. It will be seen that deterioration with respect to the x_t^{**} targets is extremely small. But performance relative to the x_t^{*} path is substantially enhanced in comparison to the case with the x_t^{**} targets, for which the comparable RMSE values are 0.0644, 0.0544, and 0.0426 with λ set at 0.25, 0.50, and 1.0, respectively. Thus these results are highly encouraging.

Another way in which the tendency for the x_t^{*a} target to pull x_t back to a fixed path can be illustrated, within the context of counterfactual historical simulations, by misspecifying the constant term in the policy rule (2). Suppose, for example, that we set that constant equal to 0.01968 while keeping the value 0.01468 in the definition of x_t^{*a}. Then there would be a tendency for x_t to grow at a faster rate than x_t^{*}, were it not for the partial dependence of x_t^{*a} on x_t^{*}. The effects are shown in Figure 8, where the two plots are obtained with the x_t^{**} and x_t^{*a} targets, $\lambda = 0.5$ being used in both cases. The contrast in "path reverting" tendencies is striking.

Figure 8

Results Using x_t^{*e} and x_t^{*a} Targets and Misspecification, Model (4)



TARA: Path of x_t^{*a}, simulated

LX ----- TAR ——— TARA

The attractiveness of a weighted-average target such as x_t^{*a} is, therefore, quite substantial. In the next two sections, accordingly, the x_t^{*a} target will be used as the basis for investigation of the robustness of performance of policy rule (2). Thus we shall be conducting experiments analogous to those of this section, while focusing on x_t^{*a} targets, in a variety of multivariate models of the economy, these being utilized in place of the regression models (3) or (4).

V. Results with VAR Models

In this section we begin to investigate the robustness of our rule's performance by conducting simulations with a number of vector-autoregression (VAR) models. Since such models are not structural, we have no firm basis for believing that their parameters would be invariant to alternative policy regimes. They provide a useful starting point, however, for consideration of issues such as the effect of including or excluding certain variables. And in practice it may be the case that parameter responses to regime changes are not large.

Accordingly, simulation exercises analogous to those described above for models (3) and (4) have been conducted with a variety of VAR systems. In each case the procedure was to estimate parameters and residuals over the sample period 1964.2-1992.4 for a VAR system that includes Δb_t as one of the variables. Then a 84 period simulation for 1972.1-1992.4 was conducted by using initial conditions pertaining to 1972.1 and feeding in estimated shocks to the equations generating values for all the variables except Δb_t, values of the latter being generated by policy rule (2). The target variable x_t^{*a} was used with five alternative values of the feedback parameter λ in each VAR specification. In these systems nominal GNP was not itself included as one of the variables but logarithms of real GNP (y_t) and the price deflator (p_t) were, so simulated values of x_t could easily be calculated as $x_t = y_t + p_t$ and compared with the x_t^{*a} target path. Results of these comparisons can be summarized by means of RMSE statistics, as mentioned above. Such statistics are reported for six VAR systems in Table 6.

As in my U.S. study, the smallest VAR considered includes the three variables Δb_t, Δy_t, and Δp_t. Four lagged values were included for each variables--the same being true, it should be said, for all of the VAR systems. The RMSE statistics for this three-variable system are reported in the first row of Table 6. The results are not very different, as it will be seen, from those with regression model (4).[25]

The second VAR model adds a short-term interest rate to the variables of the previous system. The interest rate selected for inclusion was the 3-month bill rate, but observations on the latter are not available before 1972.1. Accordingly, values of the overnight call rate were spliced on[26] for the earlier period at the estimation stage; no such step was needed at the simulation stage since the simulations begin with 1972.1. In terms of the RMSE statistics (and the x_t^{*a} target) the values are slightly better than in the previous VAR model, as the reader will see from line 2 of Table 6. Results using the call rate throughout (not reported) are almost identical.

The third and fourth VAR systems each add one variable to those of line 2. In the first case the additional variable is Δg_t, where g_t is the log of real government spending on goods and services. In the second, the additional variable is Δs_t, where s_t is the log of Japan's exchange rate relative to the U.S. dollar (expressed as Yen per dollar). In this case, the bilateral rate is used (as in many other studies) as a crude proxy for an average foreign exchange rate, since no "effective" rate is regularly published by official agencies in Japan. As Table 6 shows, these two 5-variable systems yield RMSE results almost identical to those of the 4-variable system that proceeded them.[27]

Finally, the last line of Table 6 pertains to a system in which Δq_t replaces Δs_t, q_t being the log of the real exchange rate relative to the United States.[28] In addition, four lags of the variable Δy_t^* were included as regressors in the other equations of the VAR system, with y_t^* denoting the log of real GDP in the United States.[29] No equation was estimated for this variable, however, which was taken to be exogenous. The RMSE results are very close to those of line 1. In all of these VAR systems, then,

Table 6

Results for Japan, 1972-1992, with VAR Models

RMSE Values with Target x_t^{*a}

Variables in VAR System	Value of λ: 0.00	0.25	0.50	1.00	2.00
1. $\Delta y_t, \Delta p_t, \Delta b_t$	0.0167	0.0132	0.0125	0.0122	0.0118
2. $\Delta y_t, \Delta p_t, \Delta b_t, R_t$	0.0156	0.0129	0.0123	0.0120	0.0118
3. $\Delta y_t, \Delta p_t, \Delta b_t, R_t, \Delta g_t$	0.0156	0.0130	0.0124	0.0122	0.0120
4. $\Delta y_t, \Delta p_t, \Delta b_t, R_t, \Delta s_t$	0.0155	0.0128	0.0123	0.0120	0.0118
5. $\Delta y_t, \Delta p_t, \Delta b_t, R_t, \Delta q_t$ plus Δy_t^* exogenous	0.0165	0.0134	0.0128	0.0124	0.0119

performance of the rule (2) with target variable x_t^{**} is highly satisfactory over a rather wide range of values for the parameter λ, with the best performance resulting with λ in the range from 1.0 to 2.0.

VI. Results with Classical and Keynesian Models

We now turn to models that are intended to be "structural" in the sense that they pertain to alternative theories concerning the source of business cycle fluctuations. Our versions are extremely small in scale and are not derived by explicit maximization analysis, but are designed so as to represent the main features of important and competing theoretical schools of thought. Following McCallum (1988) three general types of models will be considered: the real business cycle type, the monetary misperceptions type, and one representative of a sticky-price "Keynesian" position.

As was mentioned in footnote 3 and more generally in Section II, the main difference among these competing types of macroeconomic models concerns the specification of their aggregate-supply or Phillips-curve sectors. Consequently, the approach here will follow that of my previous work in relying upon a single specification for the aggregate demand portion of the model--i.e., for the relation describing the quantity of output that would be demanded at a given price level for consumption, investment, government, and net-export purposes together. The present investigation will depart from my U.S. study, however, by including additional variables that are presumed determinants of net export flows--that is, by recognizing more explicitly the role of international economic interactions.

In my U.S. study, the principal determinants of aggregate demand were taken to be real money balances and government purchases, with the former represented by price-level-deflated magnitudes of the monetary base.[30] The relation was estimated in first-differenced, logarithmic form with one lag of each variable included to reflect dynamics. For the sake of comparison, a similar relation has been estimated for Japan. It was found that additional lagged terms were significant, and were accordingly included. Least squares estimates for the sample period 1964.2-1992.4 are as follows:

$$(5)\quad \underset{(.002)}{\Delta y_t = 0.0024} + \underset{(.093)}{0.082\Delta y_{t-1}} + \underset{(.091)}{0.191\Delta y_{t-2}}$$

$$+ \underset{(.056)}{0.102(\Delta b_t - \Delta p_t)} + \underset{(.058)}{0.095(\Delta b_{t-1} - \Delta p_{t-1})}$$

$$+ \underset{(.059)}{0.133(\Delta b_{t-2} - \Delta p_{t-2})} + \underset{(.078)}{0.110\Delta g_t} + \underset{(.077)}{0.099\Delta g_{t-1}} + e_{5t}$$

$$R^2 = 0.272 \qquad SE = 0.0099 \qquad DW = 2.11$$

Here the degree of explanatory power, as expressed by the R^2 and SE statistics, is very nearly the same as in the U.S. case, but is spread over more quarters so individual t-statistics are somewhat smaller.

Now we add Δq_t and Δy^*_t variables, with q_t and y^*_t denoting logs of the yen/dollar real exchange rate and the U.S. level of real GDP. Both variables should in theory enter with positive signs. As it happens, none of the Δq_t or Δg^*_t terms (dated t through t-2) enter significantly, but the most satisfactory relationship is the following, which includes Δq_{t-1} and Δy^*_t:

$$(6)\quad \underset{(.002)}{\Delta y_t = 0.0017} + \underset{(.093)}{0.072\Delta y_{t-1}} + \underset{(.091)}{0.195\Delta y_{t-2}}$$

$$+ \underset{(.057)}{0.092(\Delta b_t - \Delta p_t)} + \underset{(.058)}{0.089(\Delta b_{t-1} - \Delta p_{t-1})} + \underset{(.059)}{0.125\Delta b_{t-2} - \Delta p_{t-2})}$$

$$+ \underset{(.079}{0.125\Delta g_t} + \underset{(.078)}{0.118\Delta g_t} + \underset{(.017)}{0.016\Delta q_{t-1}} + \underset{(.104)}{0.146\Delta y^*_t} + e_{6t}$$

$$R^2 = 0.288 \qquad SE = 0.0099 \qquad DW = 2.12$$

Here the international variables, while not "statistically significant," do add a bit of explanatory power and have the proper signs. The point estimates in (6) are, accordingly, used in all the simulation described in this section, with the e_{6t} residuals being incorporated as estimates of shocks to aggregate demand.[31]

Next we consider the aggregate-supply portion of the three competing theories. In the case of the real business cycle (RBC) approach it is not necessary to estimate any additional relations. That situation stems primarily from the fact that the RBC approach suggests that real variables are block exogenous with respect to monetary variables.[32] Accordingly, we take real output movements to be exogenous, which implies that the role of (6) is to determine the price level. In addition, we also take Δg_t, Δq_t, and Δy^*_t movements to be exogenous.[33] Thus the simulation exercise uses (2) and (6) to generate sequences of values for b_t and p_t with y_t, g_t, q_t, and y^*_t set equal to their actual historical values. Implied values for x_t are then calculated as $x_t = p_t + y_t$ and can be compared with target path values for the purpose of RMSE computations, etc.

The second of the three approaches is intended to represent the monetary misperceptions theory, developed by Lucas (1972) (1973). As the most notable empirical implementations of this theory were those of Barro (1977) (1978), the formulation in McCallum (1988) was based on Barro's work to a considerable extent, with money-growth surprises--measured empirically as residuals from an estimated equation designed to explain money growth rates--taken to be important determinants of real output. Monetary base measures were used instead of M1, however, and a more stringent specification of the output equation was employed--one formulated in terms of Δy_t and including Δy_{t-1} as an explanatory variable.[34]

An attempt to apply this same strategy in the present study of the Japanese economy yields, however, rather unsatisfactory results from the perspective of the theoretical approach at hand. Specifically, the surprise values of Δb_t, denoted $\Delta \tilde{b}_t$, do not have any substantial explanatory power in the equation for output:[35]

$$(7)\quad \Delta y_t = \underset{(.003)}{0.020} - \underset{(.002)}{0.0104} d7_t + \underset{(.096)}{0.057}\Delta y_{t-1}$$

$$+ \underset{(.058)}{0.077}\Delta\tilde{b}_t + \underset{(.060)}{0.035}\Delta\tilde{b}_{t-1} + \underset{(.060)}{0.099}\Delta\tilde{b}_{t-2} + e_{7t}$$

$$R^2 = 0.253 \qquad SE = 0.010 \qquad DW = 2.04$$

Here $d7_t$ is a dummy variable reflecting the break in average output growth rates between the 1960s and 1970s.[36] As can readily be seen, the coefficients attached to the $\Delta\tilde{b}_t$ terms are small in value both absolutely and in relation to their standard errors. Indeed, they sum to only 0.21 whereas the comparable figure in the U.S. study was 0.85.[37] It can be anticipated, therefore, that the simulation results for this model would be almost the same as those for the RBC approach.[38] These simulations will not be conducted, consequently, partly for reasons that will become apparent shortly.

Finally, we turn our attention to a specification more representative of moderately Keynesian views. In particular, the comparable specification in my U.S. study was designed to represent a streamlined version of the wage-price sector of the well-known MPS econometric model (which is used by the Federal Reserve's Board of Governors as its "official" quarterly econometric model). In that model, nominal wage changes are dependent, via an expectational Phillips-curve relation, on measures of capacity utilization and expected inflation. Prices then adjust gradually toward values implied by the prevailing level of wages and "normal" labor productivity growth.[39] In the present implementation, the first of these two relations was initially estimated as follows:

$$(8)\quad \Delta w_t = \underset{(.003)}{0.0098} + \underset{(.205)}{0.640}\tilde{y}_t - \underset{(.207)}{0.468}\tilde{y}_{t-1} + \underset{(.203)}{1.021}\Delta p_t^e + e_{8t}$$

$$R^2 = 0.235 \qquad SE = 0.0216 \qquad DW = 2.70$$

Here w_t is the log of nominal wages (including special payments) in manufacturing, seasonally adjusted, while the expected inflation rate is proxied by the average rate of actual inflation over the previous two quarters. Also, $\tilde{y}_t$ is the deviation of y_t from a fitted trend, with a trend break after 1971.4. The results differ to some extent from those obtained for the United States, in that wage movements are more responsive to prevailing values of the capacity variable $\tilde{y}_t$ and the coefficient on Δp_t^e is closer to 1.0.

A notable feature of the Japanese wage determination process, however, is the spring <u>shunto</u>, during which time most contracts are arranged for the next year. The existence of this feature suggests that some modification to equation (8) might be appropriate. Merely adding a second-quarter dummy ds_t has no appreciable effect, but that would not seem to be the right way to proceed in any case. More appropriate, arguably, would be to let the slope coefficient attached to $\tilde{y}_t$ be different in the second quarter than in the other three quarters. To reflect this type of effect, one can add an additional variable defined as the product $ds_t\tilde{y}_t$, since $\beta_1\tilde{y}_t + \beta_2 ds_t\tilde{y}_t = (\beta_1 + \beta_2 ds_t)\tilde{y}_t$ which makes β_1 the coefficient for all seasons except the second when it equals $\beta_1 + \beta_2$. Re-estimation with this feature incorporated for $\tilde{y}_t$ and $\tilde{y}_{t-1}$ yields the following:

$$(9)\quad \Delta w_t = \underset{(.003)}{0.0095} + \underset{(.204)}{0.593}\tilde{y}_t + \underset{(.278)}{0.506}ds_t\tilde{y}_t - \underset{(.217)}{0.492}\tilde{y}_{t-1} - \underset{(.274)}{0.236}ds_{t-1}\tilde{y}_{t-1} + \underset{(.201)}{1.046}\Delta p_t^e + e_{9t}$$

$$R^2 = 0.272 \qquad SE = 0.0213 \qquad DW = 2.62$$

These results are supportive of the idea that the shunto effect is important: the implied slope coefficients are much larger for the second quarter. The equation still has rather low explanatory power and negatively autocorrelated residuals, but those weaknesses are to a considerable extent related to an extremely large residual in the second quarter of 1974 (following the first oil shock). It would be possible to obtain "better" statistics, consequently, by including a dummy variable for that quarter. But the abnormal wage behavior of that quarter constitutes a shock of the type that actual monetary policy is occasionally required to deal with. Accordingly, equation (9) with no 1974.2 dummy will be utilized in our policy simulations.

The second equation of the MPS-type wage-price sector reflects, as mentioned above, partial adjustment of prices. Our version is estimated in first differences, obviating the need for a trend term to reflect productivity growth, as follows:

$$(10)\quad \Delta p_t = \underset{(.001)}{0.0014} + \underset{(.036)}{0.216}\Delta w_t + \underset{(.066)}{0.465}\Delta p_{t-1} + e_{10t}$$

$$R^2 = 0.578 \qquad SE = 0.0075 \qquad DW = 2.48$$

These results are not too bad, but experimentation revealed that lagged values of Δw_t would enter strongly and reduce residual autocorrelation. Accordingly, the following estimates were adopted for use in the simulations:

$$(11)\quad \Delta p_t = \underset{(.001)}{-0.0009} + \underset{(.032)}{0.180}\Delta w_t + \underset{(.037)}{0.112}\Delta w_{t-1} + \underset{(.033)}{0.153}\Delta w_{t-2} + \underset{(.085)}{0.234}\Delta p_{t-1} + e_{11t}$$

$$R^2 = 0.657 \qquad SE = 0.0068 \qquad DW = 1.96$$

In the case of this Phillips-type model, the counter-factual policy simulations use equations (2), (6), (9), and (11) to generate time paths for b_t, y_t, p_t, and w_t with x_t and RMSE values calculated as before.

Before turning to the simulation results, it may be useful to emphasize the relatively innocuous nature of econometric weaknesses in the estimation of the models in this section.[40] It is not crucial that least squares estimates of (9) and (11), for example, are subject to simultaneity bias or even that identification of (9) and (11) is questionable. The object in estimation is not to build a case that any one of the models is "true" but merely to obtain numerical representations, which are consistent with Japanese data for 1972-1992, of alternative theories of macroeconomic behavior. One could in principle just assign conjectured parameter values, provided that the shock estimates were based on those values. That parameter values are generally consistent with the data, and that shock estimates are not too far from white noise, is guaranteed in a relatively straightforward way by our approach.

Let us now consider, then, results of the simulations conducted using policy rule (2) and the target variable x_t^{*a} with the two structural models representing RBC and Keynesian theories. In the former case the rule's performance turns out to be quite poor. It succeeds in keeping nominal GNP growing at a rate of about 6 percent on average, implying an inflation rate close to the 2 percent target, but period-to-period variability is rather high. Indeed, as the first row of Table 7 shows, variability relative to the x_t^{*a} target path is increased as λ is raised above 0.25.

The reason for the poor performance in this model can be understood in the following way. If base growth rates Δb_t were kept constant, then equation (6) would (with Δy_t exogenous) generate inflation values in accordance with

(12) $\Delta p_t = 0.018 - 0.967\Delta p_{t-1} - 1.359\Delta p_{t-2}$ + exogenous terms

where 0.967 is obtained as the ratio 0.089/0.092, etc. But obviously this is a stochastic difference equation that generates highly explosive oscillations. Feedback from the policy rule (6) is somewhat helpful but cannot properly stabilize the behavior of Δp_t because lagged values of Δb_t

Table 7

Results with RBC and Phillips Curve Models

RMSE Values with x_t^{*a} Target, 1972-1992

Equations in Model	RMSE Rel. to	Value of λ: 0.00	0.25	0.50	1.00	2.00
(6) (2)	x_t^{*a}	0.0199	0.0186	0.0200	0.0362	expl.
(6) (2)	x_t^{*}	0.0563	0.0357	0.0288	0.0273	expl.
(6) (9) (11) (2)	x_t^{*a}	0.0133	0.0123	0.0117	0.0106	0.0099
(6) (9) (11) (2)	x_t^{*}	0.0486	0.0390	0.0341	0.0259	0.0183

Note: expl. denotes explosive oscillations.

appear in (6) and work in the wrong direction part of the time, due to the oscillations. Performance was much better in my U.S. study because the analog of (12) was almost stable and lagged Δb_t values were not so important.

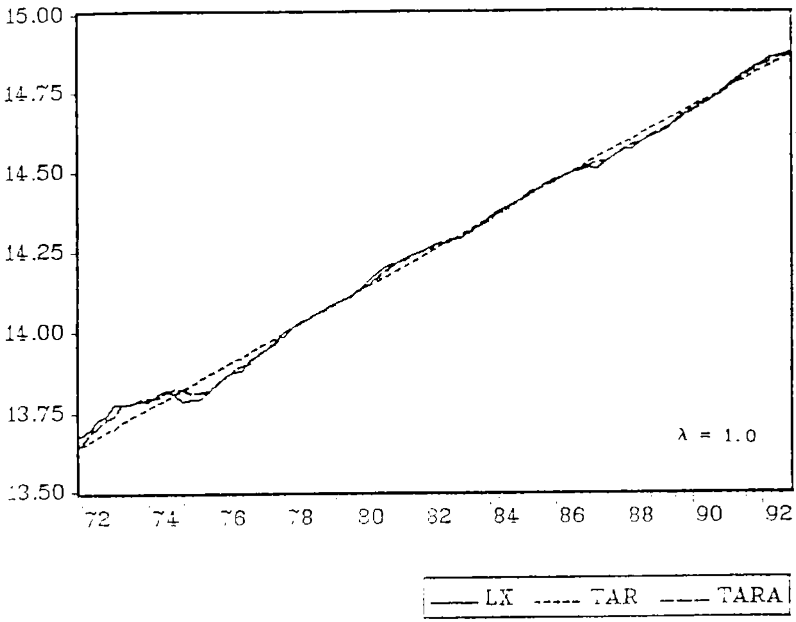
When we turn to the Keynesian or Phillips Curve model incorporating equations (6), (9), and (11), by contrast, the rule's performance is excellent. Specifically, the RMSE values reported in Table 7 compare very favorably with those given in Table 5 and 6 for the single-equation and VAR models, with performance improving with increased values of λ up through 2.0. Visual inspection of Figure 9 confirms the favorable evaluation of performance in this case.

Similar results were obtained, furthermore, when the Phillips curve model was extended to include the growth rate of the price of imported oil (plus coal and gas products) as an additional explanatory variable in the wage and price equations (9) and (11). The additional variable enters significantly, and without much effect on the other coefficient estimates, in both equations. Simulations then resulted in the following RMSE values in place of those reported in line 3 of Table 7: 0.0133, 0.0118, 0.0113, 0.0104, and 0.0104.

What conclusions are appropriate, then, on the basis of the results in this section? In a sense the results are disappointing since they do not reflect the robustness across model specifications that was found in the case of the United States. But the model in which the rule's performance is excellent is the only one of the three considered that is at all consistent with the Japanese data. It was mentioned above that the monetary misperception specification was empirically unsatisfactory and it can easily be argued that the same is true for the RBC model. For the role of (6), as explained above, is to determine values of p_t and thus Δp_t. But there is no discernible tendency for quarterly inflation rates in Japan to behave in an oscillatory fashion, much less in the explosive oscillatory manner implied by equation (12). So, the evidence of this section is in fact consistent with the idea that policy rule (2) would perform well in the Japanese context. The poor performance in the RBC model can be discounted since the latter is highly unstatisfactory, empirically. It would be useful to conduct other robustness exercises, for example by using a Taylor-style contracting scheme for Δw_t in place of relation (9), but to do so is beyond the scope of the present study.

Figure 9

Results with Phillips Curve Model and x_t^{*a} Target



VII. Results with an Interest Rate Instrument

Having obtained simulation results for policy rule (2) in a variety of models, we now wish to explore the possibility of hitting x_t^{*a} targets with a rule that specifies quarterly settings of an interest rate instrument (or operating variable). It was argued above that design of an interest rate rule is more difficult than when the monetary base is used as the instrument, but it should be possible to effect at least some stabilization. In any event, the results of attempts of this type should be instructive. Throughout this section, the experiments will be conducted using the four-variable VAR system described in Section V as the model of nominal GNP determination.[41] The simulation, that is, will be conducted using VAR equations that have Δy_t, Δp_t, and Δb_t as their dependent variables.

The fourth equation in each simulation system will, then, be a policy rule specifying values of R_t, the 3-month bill rate. But what specification should be used for such an equation? Despite the ambiguity noted above concerning levels of interest rates, there is a presumption shared by practitioners and most researchers that changes in interest rates have temporary effects that are qualitatively clear cut. An increase in R_t, that is, reflects a tightening of monetary policy that should reduce the magnitude of x_t relative to the value that it would have assumed if R_t had not been changed. Similarly, decreases in R_t should tend to increase x_t.

A natural starting point for the current investigation is provided, then, by a simple rule of the form

$$(13)\quad R_t = R_{t-1} - \lambda_1(x_{t-1}^{*a} - x_{t-1}),$$

where λ_1 is some positive policy parameter. Simulations have been conducted using (13) with different λ_1 values in conjunction with VAR equations for

Δy_t, Δp_t, and Δb_t. As in previous sections, residuals from these last three equations were fed in as estimates of shocks that occurred over the simulation period of 1972.1-1992.4.[42]

The results of the simulations using rule (13) are very poor. As the figures in the first row of Table 8 show clearly, the RMSE values relative to x_t^{*a} decline only slightly as λ_1 is increased from zero, with the smallest values occurring for λ_1 around 0.002-0.005 and then increasing steadily with higher settings of λ_1. Furthermore, the minimum RMSE values are about 0.1, roughly eight times as large as those in Table 6. Time plots analogous to Figures 6 or 9 are too unattractive to be shown.

Somewhat better results are obtained, however, if the <u>change</u> in $x_{t-1}^{*a} - x_{t-1}$ is used in the rule in place of its level. In this case the rule calls for the interest rate R_t to be decreased when target misses are growing rather than when they are large. In particular, various values of λ_2 are considered in the rule

$$(14)\ R_t = R_{t-1} - \lambda_2(\Delta x_{t-1}^{*a} - \Delta x_{t-1}).$$

RMSE results are reported in the second row of Table 8. There it can be seen that performance is substantially improved, with RMSE values falling slightly below 0.03 for λ_2 values between 0.5 and 0.75. In particular, a λ_2 value of 0.60 results in a RMSE of 0.0287. A plot of the simulated path is shown in Figure 10.

Even these improved results are rather poor, however, in comparison with those obtained with the b_t instrument in previous sections. Not only are the best RMSE values over twice as large, but also there is more sensitivity to the policy parameter value utilized. In Table 6, for example, good results are obtained for λ settings ranging over a factor of 10 whereas λ_2 values

Table 8

Results with R_t Instrument and VAR Model No. 2

RMSE Values with Target x_t^{*a}, 1972-1992

Policy Rule	Value of $100\lambda_1$ or λ_2					
	0.00	0.25	0.50	0.75	1.00	5.00
Eq. (13)	0.1717	0.1039	0.1088	0.1425	0.1712	0.3586
Eq. (14)	0.1717	0.0462	0.0310	0.0361	0.1987	expl.

Table 9

Results with VAR Model No. 2 and Rule (15)

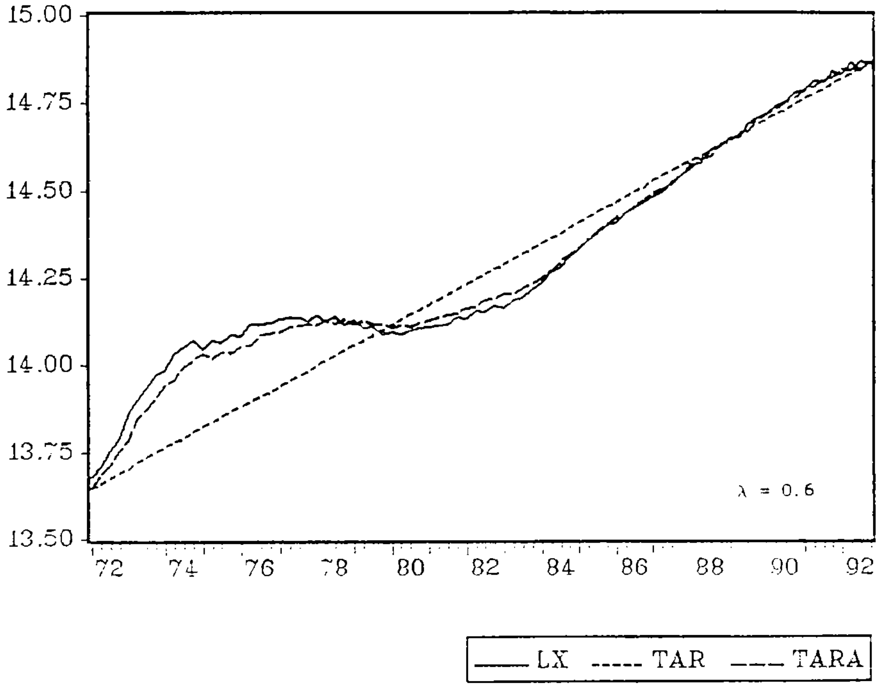
RMSE Relative to x_t^{*a} with x_t^{**} Target

Value of λ_1	Value of λ_2		
	0.3	0.4	0.5
0.2	.0276	.0261	.0272
0.3	.0255	.0247	.0306
0.4	.0264	.0268	.0424

Figure 10

Results with Interest Rate Instrument

Rule (14), VAR Model No. 2



need to be kept within the range 0.3-0.8 for reasonable results with rule (14).

It might be thought that inclusion of both types of discrepancy terms, as in the equation

$$(15)\ R_t = R_{t-1} - \lambda_1(x^{**}_{t-1} - x_{t-1}) - \lambda_2(\Delta x^{**}_{t-1} - \Delta x_{t-1})$$

but with x_t^{*a} instead of x_t^{**}, could improve performance relative to (14). It turns out, however, that the optimal value for λ_1 is in that case so close to zero that for all practical purposes (15) provides no improvement over (14) when x_t^{*a} is the target variable.

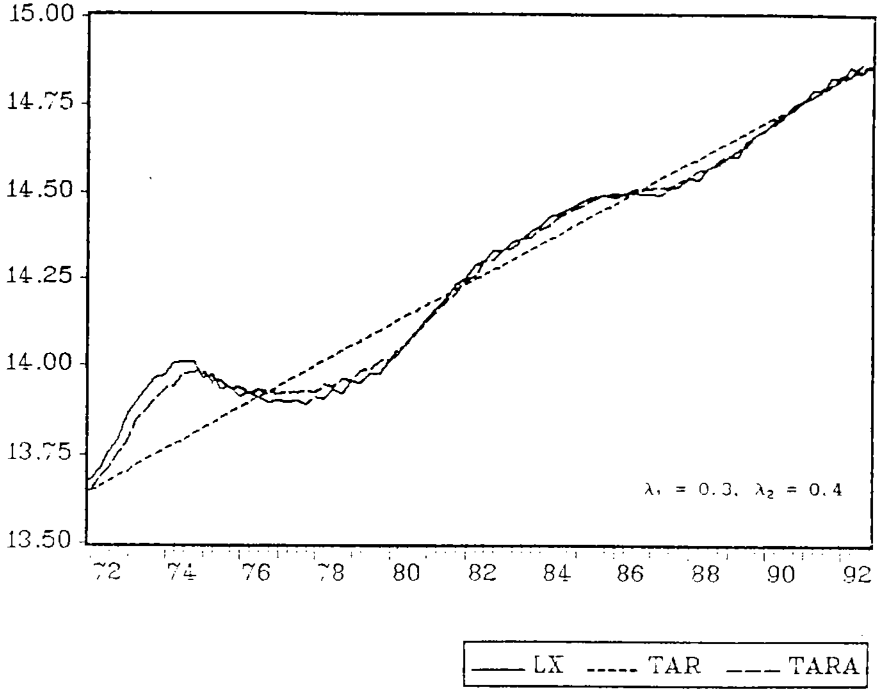
Experimentation reveals, on the other hand, that formulation (15) does function better than either (13) or (14) when x_t^{**} is used as the target variable. Rather surprisingly, performance is somewhat better even in terms of the RMSE criterion relative to x_t^{*a}! That fact is demonstrated in Table 9, which reports RMSE values for this preferred criterion over a range of λ_1 and λ_2 magnitudes. Also see Figure 11. The improvement is not sufficient, nevertheless, to alter the relative attractiveness of the R_t and Δb_t instruments.

In a final attempt to achieve improved performance with the R_t instrument, a simulation (with $\lambda_1 = 0.3$, $\lambda_2 = 0.4$) was conducted like those of Table 9 but with the residuals from the estimated VAR equation for R_t used in place of the residuals in the VAR Δb_t equation.[43] This step was taken to determine whether the poorer performance with the R_t instrument could be attributed to the fact that the estimated shocks are more variable in the VAR equation for Δb_t than in the VAR equation for R_t, since the latter equation is included when Δb_t is the instrument. A significant amount of improvement was in fact obtained, the RMSE for x_t^{**} falling from the Table 9 value of

Figure 11

Results with Interest Rate Instrument

Rule (15), VAR Model No. 2



0.0247 to 0.0201. But even in this case, it remains true that performance with the R_t instrument is substantially less satisfactory--according to our simulations--than when Δb_t values are specified by the rule. A more complete investigation would be desirable, but on the basis of these results the Δb_t variable appears to be, from the macroeconomic perspective, the more effective instrument.[44]

VIII. Concluding Discussion

To begin this final section, it may be useful briefly to consider how Bank of Japan policy actions would have been different if policy rule (2) had been in effect over the period 1972.1-1992.4. How, in other words, would base growth rates--Δb_t values--have evolved in comparison with the actual historical record? The answer is provided for two models, and a λ value of 0.5 with the x_t^{*a} target, in the two panels of Figure 12. The top panel pertains to the Phillips-curve or Keynesian model of Section VI while the bottom panel is for the four-variable VAR model. In each of these, DLB denotes the simulated path of Δb_t generated by the rule whereas DLBA denotes the actual historical path. Looking first at the top panel, we see that Δb_t variability under the rule would have been less than occurred in actuality--base growth rates would have evolved more smoothly.[45] In addition, we see that on average base growth rates would have been substantially lower than they were over the years 1972-74 and also that base growth rates would have been somewhat higher recently, i.e., during 1990-1992. During the remainder of the period, there are no extended spans during which actual policy departed strongly and systematically from that called for by the rule, although actual policy was slightly more expansive over 1987 and 1988.[46]

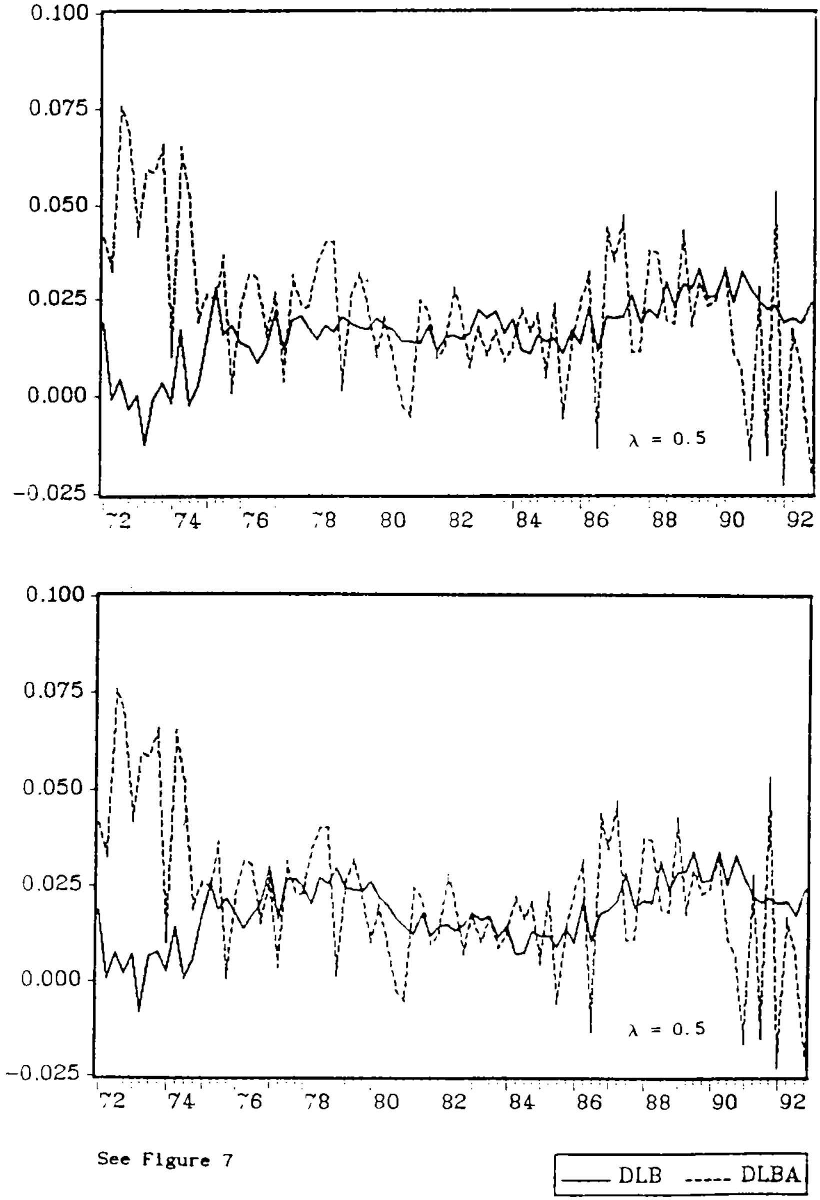
Turning to the bottom panel of Figure 12, we obtain the same impression as before. Indeed, the two plots are extremely similar. That similarity is highly desirable, from the perspective of a proponent of rule (2), since it suggests that policy actions dictated by that rule are not overly sensitive to model specification--at least for this two-model comparison.

A second rule vs. actual comparison is presented in Figure 13 where, for the four-variable VAR model and λ = 0.5, simulated (R3), and actual (R3A) time paths are plotted for the three-month bill rate. Here the striking--and surprising--feature of the comparison is the similarity between simulated and

Figure 12

Rule vs. Actual Paths of Δb_t with Phillips Curve Model and

VAR Model No. 2



See Figure 7

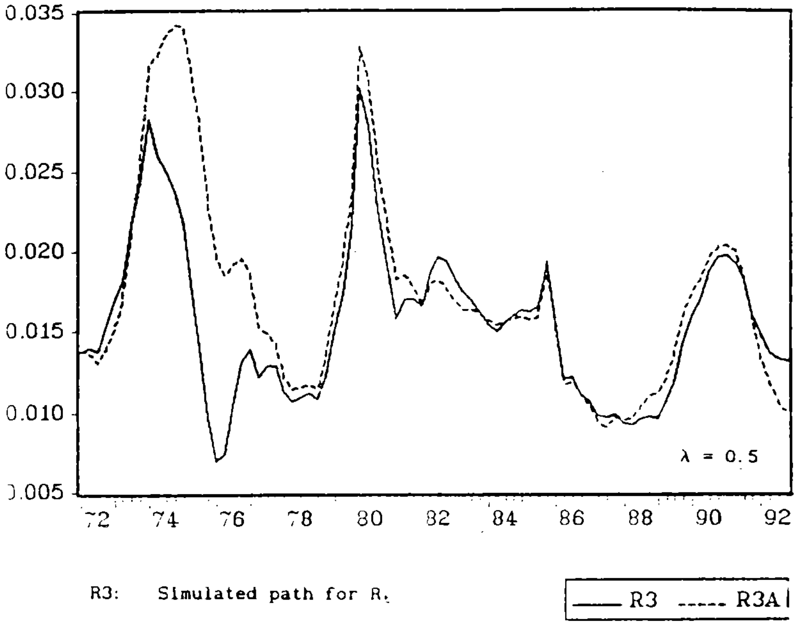
actual time paths. Only for the years 1973-1977 is there a truly major discrepancy, with actual rates being much higher--presumably reflecting the higher values of actual over simulated inflation rates. This similarity suggests that, in terms of quarterly averages, variability of short-term interest rates would not be greatly increased by adoption of a policy rule like (2).[47] Another implication of Figure 13 is that the prevailing level of the three-month bill rate is not a reliable indicator of policy ease or tightness. Neither the greater tightness during 1972-73 nor the greater ease during 1990 called for by rule (2) (and apparent in Figure 12) is reflected in the interest rate comparison of Figure 13.

Next, some discussion should be provided concerning the possible vulnerability of our various results to objections of the type emphasized in the famous policy-evaluation "critique" of Lucas (1976). That need is heightened by the absence of any explicit maximization analysis in justification of the equations of the various models that are treated as policy-invariant in the simulation exercises. It is important, I would argue, that Lucas-critique objections be taken seriously--but also that they not be applied indiscriminately. Lucas's famous critique is best thought of not as a methodological imperative, but partly as a reminder of the need to use policy-invariant relations for simulation purposes and especially as a provider of striking examples in which policy invariance would seem improbable. Explicit maximization analysis can in some cases be helpful in designing models intended to possess policy invariance, but is neither necessary nor sufficient.

In that regard, a major difficulty in the construction of an invariant model for monetary policy analysis is the profession's lack of understanding, mentioned above, of the connection between monetary and real variables. As I have argued previously, "flexible price models appear to be inconsistent with

Figure 13

Rule vs. Actual Paths of Interest Rate VAR Model No. 2



R3: Simulated path for R_t

R3A: Actual path of R_t

the behavior of actual economies, while existing sticky-price models do not conform to the dictates of the equilibrium approach..." (1990, p.21). Given this situation, the most promising strategy would seem to be to consider a variety of models in the hope that one will be reasonably well specified (and therefore relatively immune to the critique) and that all will give similar results; that is the strategy applied in McCallum (1988) and attempted here. It has transpired, of course, that the hoped-for robustness to model specification does not hold so well in the case of Japan, as the RBC model yields rather poor performance. Indeed, that is the main reason for disappointment over the RBC and monetary-misperception results--that they damage the possibility of claiming robustness to model specification as a defense against Lucas-critique objections.

A second line of defense is also present, however, in another feature of the research design. This line involves the performance criterion utilized in the simulation experiments, which is expressed in terms of the proximity of nominal GNP to its target path. In particular, this approach resists any tendency to examine simulated paths of real GNP and/or the price level separately. The reason again involves the profession's poor understanding of aggregate supply or Phillips-curve behavior, i.e., the wage-price sector of macroeconomic models. The relevant point here is that, because of the crucial role of unexpected components or surprise terms in the wage-price sector of most models, this is the sector that would seem to be most susceptible to the Lucas critique. But this is also the sector that determines how changes in nominal GNP are divided into inflation and real growth components. Accordingly, there is reason to believe that the behavior of p_t and y_t are more susceptible to the Lucas critique than is the behavior of x_t, nominal GNP. One should have more confidence, that is, concerning performance measures pertaining to x_t alone than to ones that involve the separate behavior of p_t and y_t on a period-by-period basis. So, while my defense against Lucas-critique objections is not as successful as in the case of my U.S. studies, some defense is provided by the basic strategy utilized in the study.

In conclusion, a brief summary may be useful. The study was designed to determine whether a monetary policy rule such as equation (2)--with nominal income targets and a monetary base instrument--could produce good macroeconomic performance in Japan. In selecting the precise nominal income targets to be used, it was suggested that there are fairly persuasive reasons for preferring constant growth rate targets instead of ones given by a single preset path--in other words, for treating past misses as bygones. A weighted average of these two types of target path, with a larger weight for the growth rate targets, would seem even more attractive, moreover.

With target paths taken to be of this last type, simulations with a variety of models indicated that good performance would have been obtained over the period 1972.1-1992.4. Specifically, nominal GNP values would have been kept close to target paths that would have avoided major fluctuations and yielded low inflation rates. Even though Japanese monetary policy has been rather successful over the past 15 years or so, the simulation paths are more attractive than the actual historical record.

These favorable simulation outcomes were obtained with two single-equation atheoretic models of nominal GNP determination, with several small vector autoregression systems, and with one small "structural" model featuring a wage-price sector similar to that of the Federal Reserve's quarterly MPS model. Less successful results were obtained with a model of the real business cycle type, but it was argued that the latter does not provide a sensible depiction of the Japanese economy. Using one of the vector autoregression systems, some additional experiments were conducted with an interest rate replacing monetary base growth rates as the instrument variable, but the outcomes were significantly less desirable.[48]

All in all, the simulation results are quite supportive of the idea that a policy rule with nominal spending targets could function successfully in Japan. Actual central banks are unlikely to officially adopt such rules, of course, but consideration of their implications could nevertheless prove helpful in practice, especially during times when traditional indicators are providing conflicting signals.[49] Such times may arise in the future for the Bank of Japan as financial liberalization, technical innovation, and globalization phenomena continue to occur.

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Appendix A

Description of Data Series

The following tabulation gives, for each variable used in the paper, the name of the original data series, in most cases as it appears in the Bank of Japan's publication, Economic Statistics Monthly. These series were obtained from the Bank of Japan's computerized data base. In the listing below, "sa" means that the BOJ series was obtained in seasonally adjusted form while "satsp" means that the series was seasonally adjusted by the author, before taking logs, by means of the Micro TSP ratio-to-moving-average routine. Also, "log" means the natural logarithm of the series indicated.

x_t: log of Gross National Product, Nominal sa (100 million yen)

y: log of Gross National Product, Real (At 1985 Prices), sa

p_t: $x_t - y_t$

R_t: Bill rate, 3 months; average of three monthly-average values
For dates before 1972, the original series is
Call rate, Collateralized Overnight, average of three monthly-average values, series spliced for 1972.1.

g_t: log of Government Expenditures, real (1985) Prices), sa

s_t: log of Yen/Dollar spot rates, average of monthly-average values

q_t: log of (Yen/Dollar spot rate times ratio of U.S. to Japanese GNP deflator price indexes)

y^*_t: log of U.S. Gross National Product, Real (1987 Prices)

w_t: log of Wage Index, Manufacturing, including special pay, sa

pim_t: log of Import Price Index for Petroleum, Coal, and Natural Gas, average of monthly-average value

b_t: log of adjusted monetary base, satsp

Adjusted monetary base is the sum of quarterly values of "Cash Currency Issued" and BOJ "Deposits from Deposit-Money Banks," with the latter adjusted for changes in reserve requirement ratios as described in Section III. For both series, the quarterly values are averages of end-of-month values. Also see Appendix B.

Appendix B

Here we tabulate the Required Reserve Ratios used to adjust end-of-month values of "Deposits from Deposit Money Banks" in calculating the adjusted monetary base series. For each month, the adjusted reserves series is the above Deposits series multiplied by 2.5 and divided by the Required Reserve Ratio prevailing at the end of that month. The first column gives the date on which the corresponding ratio became effective. The ratios are percentages applying to "Other Deposits" at banks of the largest size currently recognized.

Date	Ratio	Date	Ratio
Sept. 11, 1959	1.50	Sept. 1, 1973	3.75
Oct. 1, 1961	3.00	Jan. 1, 1974	4.25
Nov. 1, 1962	1.50	Nov. 16, 1975	3.75
Dec. 16, 1963	3.00	Feb. 1, 1976	3.00
Dec. 16, 1964	1.50	Oct. 1, 1977	2.50
July 16, 1965	1.00	Mar. 1, 1980	3.25
Sept. 5, 1969	1.50	Apr. 1. 1980	3.75
Jan. 16, 1973	2.00	Nov. 16, 1980	3.25
Mar. 16, 1973	3.00	Apr. 1, 1981	2.50
June 16, 1973	3.25	Oct. 16, 1991	1.30

Also, the quarterly series are presented in the following table. There RESADJQ is the adjusted reserves (deposits at BOJ) series, CASH is the Cash Currency Issued series, BASE is their sum, BASES is the seasonally adjusted value of BASE, and LB is the log of BASES.

obs	RESADJQ	CASH	BASE	BASES	LB
1963.1	1124.444	15568.00	16692.44	16677.96	9.721843
1963.2	1251.667	16296.00	17547.67	17632.74	9.777513
1963.3	1273.333	16564.00	17837.33	18236.02	9.811154
1963.4	1188.889	18491.00	19679.89	19173.40	9.861279
1964.1	1384.722	18207.00	19591.72	19574.72	9.881994
1964.2	1403.611	18841.00	20244.61	20342.76	9.920481
1964.3	1359.722	19328.00	20687.72	21150.12	9.959401
1964.4	1703.333	21252.00	22955.33	22364.55	10.01523
1965.1	2117.778	21103.00	23220.78	23200.62	10.05193
1965.2	1849.444	21485.00	23334.44	23447.57	10.06252
1965.3	2105.833	21749.00	23854.83	24388.02	10.10185
1965.4	2349.167	23693.00	26042.17	25371.94	10.14140
1966.1	3537.500	23745.00	27282.50	27258.82	10.21313
1966.2	3614.167	24352.00	27966.17	28101.75	10.24359
1966.3	2930.000	24860.00	27790.00	28411.14	10.25454
1966.4	2536.667	27088.00	29624.67	28862.24	10.27029
1967.1	3337.500	27418.00	30755.50	30728.80	10.33296
1967.2	2940.000	28118.00	31058.00	31208.58	10.34845
1967.3	3025.000	29122.00	32147.00	32865.52	10.40018
1967.4	3711.667	31806.00	35517.67	34603.57	10.45171
1968.1	4645.833	31846.00	36491.83	36460.16	10.50398
1968.2	4054.167	33199.00	37253.17	37433.78	10.53033
1968.3	3965.000	34017.00	37982.00	38830.95	10.56697
1968.4	4205.833	37205.00	41410.83	40345.07	10.60522
1969.1	4689.167	37535.00	42224.17	42187.52	10.64988
1969.2	3795.833	39114.00	42909.83	43117.87	10.67169
1969.3	3783.055	40328.00	44111.05	45096.99	10.71657
1969.4	3733.333	44287.00	48020.33	46784.47	10.75331
1970.1	4796.111	44932.00	49728.11	49684.95	10.81346
1970.2	4308.333	46438.00	50746.33	50992.36	10.83943
1970.3	4126.111	47726.00	51852.11	53011.07	10.87826
1970.4	4673.333	51768.00	56441.33	54988.74	10.91488
1971.1	5856.111	52275.00	58131.11	58080.65	10.96959
1971.2	5226.111	53570.00	58796.11	59081.16	10.98667
1971.3	5608.889	55394.00	61002.89	62366.38	11.04078
1971.4	6083.333	59713.00	65796.34	64102.98	11.06825
1972.1	7098.333	59762.00	66860.34	66802.30	11.10949
1972.2	6669.444	61943.00	68612.45	68945.09	11.14107
1972.3	7467.778	65330.00	72797.78	74424.90	11.21755
1972.4	7692.778	74139.00	81831.78	79725.73	11.28635
1973.1	8322.083	74832.00	83154.09	83081.91	11.32758
1973.2	9257.137	78469.00	87726.14	88151.45	11.38681
1973.3	9180.410	82239.00	91419.41	93462.74	11.44532
1973.4	11057.33	91432.00	102489.3	99851.63	11.51144
1974.1	10598.63	90319.00	100917.6	100830.0	11.52119
1974.2	11658.82	95484.00	107142.8	107662.3	11.58675
1974.3	11811.57	98941.00	110752.6	113228.0	11.63716
1974.4	11874.90	106506.0	118380.9	115334.2	11.65559
1975.1	12449.22	106143.0	118592.2	118489.3	11.68258
1975.2	13193.92	107595.0	120788.9	121374.5	11.70664
1975.3	13130.79	110051.0	123181.8	125935.1	11.74352
1975.4	12183.37	117127.0	129310.4	125982.4	11.74390

obs	RESADJQ	CASH	BASE	BASES	LB
1976.1	11561.33	117389.0	128950.3	128838.4	11.76631
1976.2	13133.89	119261.0	132394.9	133036.8	11.79838
1976.3	12911.39	121176.0	134087.4	137084.4	11.82835
1976.4	12283.89	130515.0	142798.9	139123.8	11.84312
1977.1	13168.33	129851.0	143019.3	142895.2	11.86987
1977.2	12980.28	129687.0	142667.3	143359.0	11.87311
1977.3	13115.28	131584.0	144699.3	147933.5	11.90452
1977.4	13736.67	141551.0	155287.7	151291.1	11.92696
1978.1	15023.00	139888.0	154911.0	154776.5	11.94974
1978.2	18754.00	140635.0	159389.0	160161.8	11.98394
1978.3	18410.00	144656.0	163066.0	166710.7	12.02402
1978.4	19201.67	158913.0	178114.7	173530.7	12.06411
1979.1	18947.33	154837.0	173784.3	173633.5	12.06470
1979.2	18726.67	158521.0	177247.7	178107.0	12.09014
1979.3	19334.33	160563.0	179897.3	183918.3	12.12225
1979.4	20352.67	172644.0	192996.7	188029.7	12.14435
1980.1	20554.64	169449.0	190003.6	189838.7	12.15393
1980.2	21225.78	171519.0	192744.8	193679.3	12.17396
1980.3	22111.33	169151.0	191262.3	195537.3	12.18351
1980.4	20653.13	179503.0	200156.1	195004.9	12.18078
1981.1	19631.79	174476.0	194107.8	193939.3	12.17530
1981.2	21283.67	176511.0	197794.7	198753.6	12.19982
1981.3	21468.33	177145.0	198613.3	203052.6	12.22122
1981.4	21246.33	189122.0	210368.3	204954.2	12.23054
1982.1	22324.67	185350.0	207674.7	207494.4	12.24286
1982.2	23657.67	188713.0	212370.7	213400.3	12.27092
1982.3	23066.67	189945.0	213011.7	217772.8	12.29121
1982.4	22343.67	202738.0	225081.7	219288.9	12.29815
1983.1	24199.33	199199.0	223398.3	223204.4	12.31584
1983.2	25052.33	199179.0	224231.3	225318.5	12.32527
1983.3	24752.33	199388.0	224140.3	229150.1	12.34213
1983.4	25609.67	211587.0	237196.7	231092.1	12.35057
1984.1	28699.33	205504.0	234203.3	234000.0	12.36308
1984.2	28140.00	209950.0	238090.0	239244.3	12.38524
1984.3	29461.67	208259.0	237720.7	243034.0	12.40096
1984.4	30923.67	223854.0	254777.7	248220.6	12.42207
1985.1	30762.67	218715.0	249477.7	249261.1	12.42626
1985.2	29635.00	224339.0	253974.0	255205.3	12.44982
1985.3	28461.67	219542.0	248003.7	253546.9	12.44330
1985.4	26991.33	236701.0	263692.3	256905.9	12.45646
1986.1	31249.00	231918.0	263167.0	262938.6	12.47968
1986.2	31096.00	239129.0	270225.0	271535.1	12.51185
1986.3	24073.67	237922.0	261995.7	267851.6	12.49819
1986.4	24891.00	262467.0	287358.0	279962.5	12.54241
1987.1	26845.67	263176.0	290021.7	289769.9	12.57684
1987.2	30860.00	271379.0	302239.0	303704.3	12.62381
1987.3	29577.00	270693.0	300270.0	306981.4	12.63454
1987.4	30291.67	288276.0	318567.7	310368.9	12.64552
1988.1	31857.33	290657.0	322514.3	322234.4	12.68303
1988.2	32848.00	299896.0	332744.0	334357.2	12.71996
1988.3	37331.33	295952.0	333283.3	340732.6	12.73885
1988.4	40820.00	315222.0	356042.0	346878.8	12.75673

obs	RESADJQ	CASH	BASE	BASES	LB
1989.1	41628.00	320887.0	362515.0	362200.3	12.79995
1989.2	38219.67	328551.0	366770.7	368548.9	12.81733
1989.3	42806.00	328477.0	371283.0	379581.6	12.84682
1989.4	44622.67	353927.0	398549.7	388292.5	12.86951
1990.1	46194.33	352087.0	398281.3	397935.6	12.89405
1990.2	47537.67	362045.0	409582.7	411568.4	12.92773
1990.3	53141.00	353721.0	406862.0	415955.9	12.93833
1990.4	53067.00	376267.0	429334.0	418284.5	12.94392
1991.1	48462.33	363175.0	411637.3	411280.0	12.92703
1991.2	49651.67	371380.0	421031.7	423072.9	12.95530
1991.3	46990.00	360237.0	407227.0	416329.0	12.93923
1991.4	68295.52	382852.0	451147.5	439536.7	12.99348
1992.1	59050.00	370590.0	429640.0	429267.1	12.96983
1992.2	58614.11	375886.0	434500.1	436606.7	12.98679
1992.3	59883.98	370022.0	429906.0	439514.9	12.99343
1992.4	54010.26	388241.0	442251.3	430869.3	12.97356

Footnotes

[1]Studies by other economists that apply, extend, or investigate this rule include Flood and Isard (1988), Hall (1990), Judd and Motley (1991, 1993), Hafer, Haslag, and Hein (1990), Hess, Small, and Brayton (1992), and Dueker (1993). Some critical comments have been put forth by Friedman (1988, 1990), with a reply by the author following in the first of these publications.

[2]Previous studies concerning policy rules for Japan include West (1993), McNelis and Yoshino (1992), and Taylor (1988).

[3]This point has been emphasized in theoretical studies by Bean (1983), Henderson (1992), and others. Such results are highly model-dependent, however, with the relative desirability of different target variables depending on the relative variances of different shocks, the serial correlation properties of these shocks, the relative magnitudes of static supply and demand elasticities, and the precise specification of the dynamic Phillips-curve mechanism. The latter is one of the most poorly understood relations in all of macroeconomics, incidentally: there are multitudes of competing theories but none that combines empirical validity with a sound theoretical basis, involving maximization analysis incorporating individuals' objectives and constraints.

[4]As mentioned above, this assertion will be discussed below (in Section III).

[5]This ambiguity will be discussed in Section III.

[6]That it is not feasible was argued by Axilrod (1985) in a comment on a study, principally devoted to other topics, by the present author.

[7]Who cannot experiment with actual economies.

[8]Values are reported for only one of the several VAR systems.

[9] For comparative purposes, it might be noted that the RMSE value for the actual historical path is 0.771, over 30 times as large as the cases with the rule and moderate λ. Another relevant comparison, since a large part of the 0.771 value reflects inflation rather than variability, is the RMSE value for the actual historical path relative to a fitted trend line; that value is 0.0854.

[10] Here I am classifying effects of the "Lucas Critique" type as second-order. My defense against this type of criticism is explained below in Section VIII.

[11] Specifically, legal requirements pertain to average balances held over month-long periods that begin two weeks after the end of the month-long accounting periods over which deposits are averaged to determine the magnitude of the required reserves. The accounting periods, not the maintenance periods, coincide with calendar months. See Okina (1993) for more details.

[12] Over the period 1967-1987, excess reserves averaged only 0.14 percent of required reserves (Ueda, 1993, fn. 9).

[13] Another possibility would be for reserve requirements to be reduced to a level that is not binding, so that all banks would normally hold an appreciable volume of excess reserves. This would not necessarily reduce the effectiveness of money stock control since required reserves are irrelevant for that purpose when an interest rate instrument is employed. On this subject see Goodfriend and Hargraves (1983) and McCallum and Hoehn (1983).

[14] Overnight interbank interest rates might still tend to fall to exceedingly low values on the final day of a reserve maintenance period if reserve requirements turn out to be smaller than expected.

[15] Another possible adjustment mechanism is that banks would begin to fail more frequently to meet their legal reserve requirements. The explicit penalties imposed with such a failure are not very large--see Okina (1993, p.50)--and the non-pecuniary costs might diminish as failures became more common. Actually, both types of adjustment would probably occur to some extent.

[16] This point has been recognized by Ueda (1993, fn. 9).

[17] In the work described in the previous section, the adjusted base series utilized is that prepared and published by the Federal Reserve Bank of St. Louis.

[18] Thus the implicit inflation target in this study is approximately 2 percent per year whereas it is zero in my previous studies. This difference should not be interpreted as expressing a belief that the BOJ is more willing than the Federal Reserve to accept inflation; my previous studies were not prepared for publication by the Federal Reserve.

[19] The values of the monetary base were obtained as described in Section III. Seasonal adjustment was effected by applying the ratio-to-moving-average technique as performed by the program Micro TSP (i.e., the multiplicative option) before taking logarithms. The moving average values span a year centered on the current observation.

[20] This sample period was used because some data series were not available prior to 1963.1 and 4 lags (in first differences) were needed for the vector autoregression models.

[21] The DW statistic is, of course, inappropriate for a formal test of residual autocorrelation in any equation that includes a lagged endogenous variable. Values are reported in this paper, nevertheless, to provide a general indication of the extent of first-order autocorrelation. Values reasonably close to 2.0 are a necessary condition for the absence of serial correlation problems.

[22] A word of explanation is needed for the initial date used in the simulations. While the exact quarter is somewhat arbitrary, one in the vicinity of 1972.1 is desirable because of (i) the end of the Bretton Woods System, (ii) a reduction in the average growth rate of real GNP, and (iii) the desirability of beginning with initial values of Δx_{t-1} that are not too far from the target value.

[23] It must be recognized, however, that removal of a linear trend for x_t may not yield an appropriate measure of smoothness. If a cubic trend is removed instead, the residual variability falls to 0.015.

[24] This is not quite correct, because the estimation and simulation periods are not the same. But since the latter is a large subset of the former, the practical point remains.

[25] Some readers may wish to ask why no study of the "unit-root" properties of the variables was conducted prior to estimation, as has become customary in recent years. The answer is that I subscribe to Cochrane's (1991) argument that general knowledge of a variable's behavior is more reliable than formal tests for determining these properties. Also relevant is the argument in Section IV of McCallum (1993), which suggests that quantity variables such as real GNP will include both trend-stationary and unit root components.

[26]The splice was effected by adding 0.000942 to each value of the call rate, that number being the excess of the bill rate over the call rate in 1972.1. Both rates are expressed in quarterly fractional units, i.e., percentage rates on an annual basis (as reported) are divided by 400.

[27]The regression equation for Δs_t was estimated over the reduced sample period 1972.1-1992.4, since the Bretton Woods system prevailed (more or less) until August 1971.

[28]The price indexes used are the two nations' GNP deflators.

[29]The point, of course, is to include a proxy for foreign income levels as these are relevant in most open-economy models.

[30]Thus the aggregate demand function implicitly incorporates banking sector relations reflecting the connection between the money stock and the monetary base.

[31]It might be asked why (6) is estimated by means of ordinary least squares rather than some instrumental variable estimator that might reduce possible simultaneity bias. The answer involves the difficulty in finding appropriate instruments. There are probably no variables of macroeconomic importance that are actually exogenous, so one is forced to turn to lagged endogenous variables. If relation (6) is estimated with a set of instruments that includes all lagged variables in (6) plus one additional lagged term for each variable the estimated relation features slightly reduced explanatory power, much larger standard errors, and a much larger point estimate of the parameter on the $\Delta b_t - \Delta p_t$ term. Because of this latter property, the use of this relation in the simulations would tend to sharply increase the stabilizing power of policy operations with Δb_t. Accordingly, use of (6) is considerably more conservative in the context of the present study.

[32]It is also being assumed here that any fiscal effects on output work through an intermediate impact on aggregate demand. This is a simplification of RBC views, but one that is perhaps justified by the approach's emphasis on technology shocks as the predominant source of cyclical fluctuations.

[33]In principle it would be more appropriate to determine Δq_t endogenously, but since it enters (6) so weakly the results would not be affected significantly.

[34]The money growth specification is simpler than Barro's; some justification is provided in footnote 22 of McCallum (1988).

[35]As in McCallum (1988), the $\Delta\tilde{b}_t$ values are residuals from an autoregression for Δb_t. In the Japanese case, a fourth-order AR is used because Δb_{t-4} provides more explanatory power by far than any other variable.

[36]The break is dated, rather arbitrarily, after 1971.4 to coincide with our simulation start-up date.

[37]That the Lucas-Barro model performs poorly for Japan will come as no surprise to readers of Okina (1986).

[38]In these simulations Δb_t, Δp_t, and Δy_t values would be generated by the equations (2), (6), and (7) with residuals fed in and with $\Delta\tilde{b}_t$ values in (7) set equal to zero since the monetary rule (2) is deterministic. The implied y_t values would therefore differ from the historical values only to the extent that $\Delta\tilde{b}_t$ terms are important.

[39]For relevant references, see McCallum (1988, p. 190).

[40]The present paragraph, like several other portions of the paper, is based on the discussion in McCallum (1988).

[41]The only models studied above that include interest rate variables are the VAR systems in lines 2-5 of Table 6. Since the results are similar in all four of these, the simplest system was adopted.

[42]Note that now residuals from the Δb_t equation in the VAR system are being fed in rather than R_t residuals as in Section V.

[43]The VAR equation for R_t is not itself used in this simulation, R_t values being generated by (15).

[44]It should be mentioned that Hess, Small, and Brayton (1992) have found that an interest rate instrument performs better than a base instrument in the Fed's MPS model, with the latter leading to dynamic instability. The form of R_t policy rule used in that study specifies deviations of R_t from its historical path, however, rather than changes from the previous quarter. The rule is not operational, therefore--it is not a rule that could be given to a policymaker to put into operation in real time. For an elaboration on this point, see McCallum (1992).

[45]Interestingly, the same is not true for the United States. More, not less, instrument variability would have been required by the rule.

[46]This last conclusion would be strengthened if one were to utilize a 4 percent per annum growth target for nominal GNP, rather than the 6 percent target used here, reflecting an implied inflation target of zero rather than 2 percent (in terms of the GNP deflator).

[47]The analysis says nothing, of course, about week-to-week or day-to-day variability.

[48]This statement pertains only to the macroeconomic perspective; it does not take account of possible effects on the economy's financial stability. I hope to consider effects of this type in a subsequent study.

[49] For a recent argument stressing the benefits of unofficial use in this way of a policy rule, see Taylor (1993).