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THE USEFULNESS OF P^* MEASURES FOR JAPAN AND GERMANY

Linda S. Kole and Michael P. Leahy

ABSTRACT

This paper develops measures of long-run equilibrium price levels (P^*) for Japan and Germany following the approach used for the United States by Hallman, Porter, and Small [1991]. Under this approach, P^* is determined by potential output, equilibrium velocity, and the amount of money in the economy. Constructing P^* for these foreign countries is more complicated than in the U.S. case because the velocities of the broad monetary aggregates (M2+CDs in Japan and M3 in Germany) exhibit clear downward trends in contrast to the relatively flat trend of U.S. M2 velocity. We utilize dynamic specifications of money demand to construct measures of equilibrium velocity and P^* for Japan and Germany. We then assess the explanatory power of deviations of actual prices from P^* in predicting the amount of inflationary potential in the Japanese and German economies. In general, we find that the P^* approach is useful in the analysis of German inflation, but that it is less promising for Japan than it has been for the United States.

The Usefulness of P^* Measures for Japan and Germany

Linda S. Kole and Michael P. Leahy[1]

Introduction

The long-run relationship between monetary aggregates and price levels has been complicated in recent years by financial innovation, deregulation, and the increasing degree of international capital market integration. Nonetheless, Hallman, Porter, and Small [1989, 1991 (HPS)] analyze the relationship between U.S. money and prices using a simple model based on what they term "P^*". They define P^* as the long-run equilibrium price level that current holdings of money would support if output and velocity were to settle down to their long-run values. HPS find that deviations of the actual price level from P^* has explanatory power for the acceleration of U.S. prices.

This paper discusses the construction of P^* measures for Japan and Germany and utilizes these measures to analyze the link between key money aggregates and prices.[2] Under the HPS approach, the difference between the actual and long-run price levels is seen as a major force behind inflation. We consider whether comparable deviations are valuable in forecasting German and Japanese inflation. In general,

1. The authors are staff economists in the Division of International Finance. This paper represents the views of the authors and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System or other members of its staff. We acknowledge helpful comments from Lewis Alexander, Neil Ericsson, Dale Henderson, David Howard, Will Melick, John Morton, Stephen O'Connell and Charlie Thomas. Joerg Dittmer, Maya Larson, and John Maluccio provided helpful research assistance.

2. We choose to confine our analysis to those monetary aggregates that are central to the conduct of German and Japanese monetary policy. The Bundesbank announces annual targets for M3 and the Bank of Japan forecasts growth of M2+CDs each quarter in order to promote price stability.

we find the P^* approach to be a useful tool in the analysis of German inflation of the 1970s and 1980s. However, we find the P^* analysis to be less promising for Japan than for the United States.

HPS define P^* in terms of the quantity equation or the identity MV = PY. Using lower case letters to stand for logarithms of variables, P^* is defined as:

$$(1) \quad p^* = m + v^* - y^*,$$

where m is the (log of the) current quantity of some measure of money, v^* is the long-run equilibrium velocity associated with that measure of money, and y^* is potential output. Noting that the velocity of U.S. M2 was relatively trendless for a thirty-year period beginning in 1955, HPS use the mean velocity during those years as V^*.

A crucial assumption underlying the HPS approach is that in the United States, V^* is constant, and therefore, P^* is proportional to M2 per unit of potential output, Y^*. Because in both Japan and Germany the income velocity of the key monetary aggregate exhibits a downward trend, the assumption of constant velocity must be relaxed. A paper produced in the Research and Statistics Department of the Bank of Japan [1990] specifies V^* as a trend and considers how deviations from this trend contribute to an explanation of inflation. The problem with this approach, however, is that without a good understanding of the factors underlying the trend in velocity, it is impossible to tell if and when the trend will come to an end. Herrmann and Toedter [1990] use two approaches to calculate the path of V^* for Germany: a trend, and a simple estimated money demand function. Our work is similar in spirit to the latter approach.

We take the analysis to the next logical step by trying to put some structure on trends in velocity by utilizing long-run money demand relationships that appear to be stable. This allows us to specify a path for V^* that is consistent with a steady-state relationship between money, prices, income, and other variables, but increases the complexity of the analysis. Whereas HPS must assume that the future income velocity of money remains constant, in our analysis it is necessary to make assumptions about the future course of the forcing variables that influence velocity. Our approach is thus contingent upon a richer set of assumptions.

We use dynamic specifications of money demand to construct measures of equilibrium velocity and P^* for Japan and Germany. We then assess the explanatory power of deviations of actual prices from P^* in predicting the amount of inflationary potential in the Japanese and German economies. In the next section we discuss the method used to construct P^* for Japan. Then, using the deviations of actual prices from P^*, we fit models in which inflation is a function of the price gap to see how well they perform in sample and in out-of-sample forecasts. The next two sections contain similar discussions and analyses for Germany. The last section summarizes our results and contains some concluding remarks.

Constructing P^* for Japan

One way to construct P^* is to use expression (1), which requires values for m, v^*, and y^*. To assemble an equilibrium velocity v^* for

Japan, we adapt a money demand specification developed by Corker [1989]. Corker's long-run money demand function is given by:

$$(2) \quad m - p = \tilde{\Theta}_1 + \Theta_2 w + (1-\Theta_2)y - \tilde{\lambda} \ln(1+i_g-i_o),$$

where $\tilde{\Theta}_1$, Θ_2, and $\tilde{\lambda}$ are functions of the short-run parameters Corker estimated;[3] w is (the log of) real wealth; i_g is the three-month gensaki rate, the nominal interest rate on three-month government bond repurchase agreements (at an annual rate, levels, not logarithms); and i_o is the own-rate on deposits (at an annual rate, levels). Evaluating the above expression at long-run equilibrium values w^*, y^*, i_g^*, and i_o^*, we obtain a value of p^* as the difference between current money holdings and the long-run equilibrium demand for real balances:

$$(3) \quad p^* = m - \{\tilde{\Theta}_1 + \Theta_2 w^* + (1-\Theta_2)y^* - \tilde{\lambda} \ln(1+i_g^*-i_o^*)\}.$$

Implicit in this construction of p^* is a definition of equilibrium velocity v^*:

$$(4) \quad v^* = y^* - \{\tilde{\Theta}_1 + \Theta_2 w^* + (1-\Theta_2)y^* - \tilde{\lambda} \ln(1+i_g^*-i_o^*)\}$$

$$= - \tilde{\Theta}_1 - \Theta_2(w^*-y^*) + \tilde{\lambda} \ln(1+i_g^*-i_o^*).$$

In this specification, equilibrium velocity changes as the long-run wealth-income ratio changes and as the long-run opportunity cost of money changes. Below, we examine alternative assumptions about the equilibrium

3. In Corker's specification, $\tilde{\Theta}_1 = 0.535$, $\Theta_2 = 0.535$, and $\tilde{\lambda} = 1.535$.

wealth-income ratio and opportunity cost to see how these assumptions affect the construction of v^* and p^*.

The next task is to obtain long-run equilibrium values for w, y, i_g, and i_o. We use the trend in Japanese real GNP (see the data appendix for a description of the source data for this and other variables) to obtain potential output, y^*. This trend breaks in the fourth quarter of 1973, to account for the oil shock. Prior to that quarter, potential output is assumed to grow at 8.2 percent per annum; afterward, potential output is assumed to grow at 3.8 percent. The top panel of chart 1 depicts the evolution of potential and actual real GNP.

We tried two methods to obtain long-run equilibrium wealth. In the first, we assumed that in long-run equilibrium, the wealth-income ratio is fixed. As shown in the bottom panel of chart 1, Japanese wealth-income ratios trend upwards throughout the sample period, roughly doubling from a low of about 18 in 1970 to a high of about 35 beginning in 1988. In the first method, we picked 35, a recent value, as an equilibrium ratio, and then calculated w^* as :

(5) $w^* = \delta + y^*$, where $\delta = \ln(35)$.[4]

The resulting series for w^* is plotted at the top of chart 2 along with the log of the real wealth variable used to estimate Corker's money demand function.

4. The actual value of δ affects only the constant term in the price gap equations, and, consequently, it does not matter for much of the subsequent analysis.

In the second method, we estimated w^* using a linear trend, as shown in bottom panel of chart 2. This method produces a wealth variable that grows at about 7-1/2 percent per year in sample.

$$(6)\quad w^* = 13.714 + 0.019*(\text{time}).$$

To calculate the long-run equilibrium interest rates, we assumed the following relationship between the long-run nominal rates:

$$(7)\quad i_o^* = \alpha + \beta\, i_g^*.$$

The market interest rate i_g is viewed as reflecting market pressures and the stance of monetary policy, and the deposit rate i_o is seen as adjusting, perhaps slowly, to some equilibrium relationship with the market rate. Equation (7) holds when i_g is in long-run equilibrium and i_o has fully adjusted to that level of i_g. These rates are shown in chart 3.

Because of reserve requirements, transactions taxes, account fees, etc., α need not equal zero and β need not equal unity. However, changes in regulations during the sample period may have hampered the ability of the data to reveal an underlying long-run relationship. Chart 3 shows that the own interest rate appears to have become more responsive to changes in the gensaki rate over the course of the sample. Therefore, following Moore, Porter, and Small [1990], we looked at scatter plots of the two interest rates, shown in chart 4, to see if any subperiods of the data might correspond more closely to an equilibrium relationship. None was apparent. The top panel of chart 4 shows a scatter plot of the whole sample from 1970 to 1989, and the bottom panel shows a subsample from 1985 to 1989. The subsample shows a counterclockwise shift in the relationship between the two rates. Beginning about 1987Q1, the slope and intercept appear to have shifted, perhaps corresponding to the deregulation of interest rates on large time deposits, which had begun about five quarters earlier. (See Ueda [1990], p. 187.)

We also fit lines to the data for the whole sample and three subsamples. These results are shown in table 1. While no clear choices for α and β emerge from these exercises, we favor the relationships associated with data from the most recent subperiod rather than from earlier subperiods or the whole sample, because it seems likely that earlier regulatory restraints on deposit rates could have obscured the equilibrium relationship between the two interest rates. As can be seen from the table, the coefficient estimate for α declines as the samples become more recent, while the coefficient estimate for β increases. One might expect that as deregulation progressed, restrictions that cause the deposit rates to move differently from market rates would be relaxed, and α and β would tend toward zero and one, respectively.

After choosing values for α and β, one can use them to calculate the equilibrium opportunity cost in the money demand function:

$$(8)\quad \tilde{\lambda}\ \ln(1+i_g^*-i_o^*) = \tilde{\lambda}\ \ln((1-\alpha)+(1-\beta)i_g^*)$$

$$= \tilde{\lambda}\ \ln[(1-\alpha)(1+(1-\beta)i_g^*/(1-\alpha))]$$

$$= \tilde{\lambda}\ \ln(1-\alpha) + \tilde{\lambda}\ \ln(1+(1-\beta)i_g^*/(1-\alpha)).$$

Substituting this term back into the money demand expression, we obtain:

$$(9)\quad p^* = m - \{[\tilde{\Theta}_1 - \tilde{\lambda}\ln(1-\alpha)] + \Theta_2 w^* + (1-\Theta_2)y^*$$

$$- \tilde{\lambda}\ln(1+(1-\beta)i_g^*/(1-\alpha))\}$$

$$= m - \{\Theta_1 + \Theta_2 w^* + (1-\Theta_2)y^* - \tilde{\lambda}\ \ln(1+(1-\beta)i_g^*/(1-\alpha))\},$$

where $\Theta_1 = \tilde{\Theta}_1 - \tilde{\lambda}\ \ln(1-\alpha)$.

Let $z^* = (1,\ w^*,\ y^*)'$ and $\Theta = (\Theta_1,\ \Theta_2,\ 1-\Theta_2)'$. Then (9) can be rewritten:

$$(10)\quad p^* = m - \Theta' z^* + \tilde{\lambda}\ \ln(1+(1-\beta)i_g^*/(1-\alpha))\}$$

$$\approx m - \Theta' z^* + \tilde{\lambda}(1-\beta)i_g^*/(1-\alpha)$$

$$= m - \Theta' z^* + \lambda\ i_g^*,$$

where $\lambda = \tilde{\lambda}(1-\beta)/(1-\alpha)$. When β is set equal to unity, i_o^* and i_g^* move together and the interest rate differential, or the opportunity cost of holding M2+CDs, is constant. In this extreme case, λ is zero, velocity does not respond to movements in interest rates, and movements in real interest rates and inflation expectations play no role in the calculation of p^*. If instead β is set equal to zero, deposit rates do not respond to changes in market rates, and changes in real interest rates or inflation expectations have their maximum effect on the calculation of p^*.

Since i_g^* is a nominal interest rate, it can be decomposed into a real interest rate, r_g^*, and expected inflation, $(p_{t+1}^{*e}-p^*)$. Then

$$(11)\quad p^* = m - \Theta' z^* + \lambda(r_g^* + 4(p_{t+1}^{*e}-p^*))$$

$$= m - \Theta' z^* + \lambda(r_g^* + x),$$

where the expected rate of inflation x is defined as $4(p_{t+1}^{*e}-p^*)$. The inflation term is annualized to make it comparable to the interest rate, which is also at an annual rate.

To obtain a reduced form for p^*, we need to make some assumption about how expectations are formed. For simplicity, we use the perfect foresight assumption. Setting $p_{t+1}^{*e} = p_{t+1}^*$ yields

$$(12)\quad p^* = m - \Theta' z^* + \lambda(r_g^* + 4(p_{t+1}^*-p^*)).$$

Solving for p^* in terms of the other variables yields

$$(13)\quad p_t^* = 1/(4\lambda+1) \sum_{i=0}^{\infty} (4\lambda/(4\lambda+1))^i [m_{t+i} - \Theta' z_{t+i}^* + \lambda r_{t+i}^*],$$

where the subscript g is suppressed on r^*. While this expression is a reduced form for p_t^*, p_t^* cannot be calculated directly from this expression due to the infinite sum. It is easier to calculate p_t^* in terms of x_t, using the following recursive method. Write x_t as:

$$(14)\quad x_t = 4(p^*_{t+1}-p^*_t).$$

Using (12), rewrite x_t in terms of x_{t+1}:

$$(15)\quad x_t = 4\{(m_{t+1}-m_t) - \Theta'(z^*_{t+1}-z^*_t) + \lambda(r^*_{t+1}-r^*_t) + \lambda(x_{t+1}-x_t)\}$$

$$= 4/(1+4\lambda)\{(m_{t+1}-m_t) - \Theta'(z^*_{t+1}-z^*_t) + \lambda(r^*_{t+1}-r^*_t) + \lambda x_{t+1}\}.$$

Setting inflation expectations to a quarterly rate shows that they are a weighted average of future one-step-ahead changes in the difference between nominal money balances and real money demand:

$$(16)\quad x_t/4 = 1/(1+4\lambda)\ \{(m_{t+1}-m_t) - \Theta'(z^*_{t+1}-z^*_t) + \lambda(r^*_{t+1}-r^*_t)\}$$

$$+ 4\lambda/(1+4\lambda)\ \{x_{t+1}/4\}.$$

With this formula and some assumptions about the future paths of money, income, wealth, and interest rates, expected inflation can be calculated recursively. Consider a sample of data m_t, z^*_t, and r^*_t, where $t=1,...,T$. Since x_t depends on values of these variables from dates beyond T, we must make some assumptions about their values. Assume that m_t, z^*_t, and r^*_t from dates T, T+1, T+2, . . . grow at the constant rates μ, γ, and ϕ. Then x_T is given by:

$$(17)\quad x_T = 4(\mu - \Theta'\gamma + \lambda\phi).$$

Earlier values for x_t can be calculated recursively using formula (15). After putting together a series x_t, p_t^* and v_t^* can be calculated as shown in equation (18) below:

$$(18)\quad p_t^* = m_t - \Theta' z_t^* + \lambda(r_t^* + x_t), \text{ and}$$

$$v_t^* = -\Theta_1 - \Theta_2(w_t^* - y_t^*) + \lambda(r_t^* + x_t).$$

Several specifications for p^* and v^* were computed. These can be grouped into categories that correspond to the method used to construct w^*. One group uses values of w^* calculated from equation (5), in which the long-run equilibrium wealth-income ratio is assumed to be fixed at 35. Another group uses the trend wealth variable from equation (6).

Within each group, we considered five different interest rate relationships, shown in table 2. Four correspond to the values of α and β estimated over the four sample periods shown in table 1. The fifth sets α equal to zero and β equal to unity; it captures the extreme case in which long-run velocity does not respond to movements in interest rates and inflation expectations play no role in the calculation of p^* and v^*. The table also shows the parameter values required to compute p^* (see equation (18)) that result from using different values of α and β. The parameter Θ_2 is not affected by variations in α or β.

The coefficient on inflation expectations, λ, varies widely with the different values for α and β. The largest value of λ used is 1.32, obtained when α and β are estimated using the full sample period. With later sample periods, however, the size of λ declines toward zero. When

α and β are assumed to be zero and one, λ equals zero. This pattern corresponds to the idea that inflation expectations affect the opportunity cost of holding money only to the extent that the nominal interest rate on M2+CDs fails to respond as fully as the gensaki rate to a change in inflation expectations.

Some aspects of all specifications were the same. The period of calculation was 1970Q1 to 1989Q4. Beyond 1989Q4, Japanese M2+CDs is assumed to grow at an annual rate of 10 percent, and real GNP and real wealth are assumed to grow at annual rates of 4 percent. The long-run equlibrium real interest rate is assumed to be constant forever from 1970Q1 at 4 percent.

The values of p^* and v^* do not vary much across the interest rate specifications. As shown in chart 5, the different values of p^* and v^* within each group of wealth variables are highly correlated. The values calculated using the assumption of a constant wealth-income ratio are shown in the top panels, and the values calculated using trend wealth are shown in the bottom panels. While the intercepts vary across interest rate specifications, the movements are similar. The values of p^* generally increase smoothly, with a faster rate of growth prior to 1973 than afterward. The simple correlation between any p^* pair in the top left panel is at least 0.99. The corresponding values of v^*, shown in the top right panel, are generally flat. These fail to pick up the trend in velocity but may represent the long-run equilibrium velocity the economy is moving toward. In the lower panel, v^* fits the trend in velocity better.[5] Accordingly, the values of p^* show a bit more variation during the sample period, although, again, the p^* series are highly correlated across the interest rate specifications, with simple correlations of at least 0.96.

Because the results do not appear to be sensitive to the particular values of α and β used, for presentational simplicity we focus on results for only two of the p^* and v^* candidates, using the α and β estimated over the most recent sample period. This interest rate specification seems most likely to reflect the long-run relationship between the interest rates. Thus, only two alternative specifications, corresponding to the two assumptions about w^*, will be considered below.

The top panels of charts 6 and 7 show how p^* moves relative to actual prices p. Similar comparisons for the output and velocity components of p^* are shown below. Under either specification, during most of the two decades shown, P^* has been below the GNP deflator, which, according to the hypothesis, indicates downward pressure on the rate of inflation. Near the end of the period in both specifications and in 1972 and 1973 in the trend wealth specification (chart 7), P^* was above the deflator, which should indicate upward pressure on the rate of inflation. Output below potential and V^* below actual velocity are also supposed to indicate reduced pressures on the rate of inflation.

5. One might question the usefulness of explaining the trend in velocity with a trend in wealth. However, the two are linked, and it may be easier to find explanations for both by looking to demographics and life-cycle models of saving. If a major proportion of the Japanese population is due to retire around the same time, there may be some significant aggregate wealth accumulation in Japan during the working years of that sector of the population. This could drive the aggregate wealth-income ratio higher until the sector retires.

Augmented Dickey-Fuller tests were conducted to determine whether p and p^* are cointegrated. Rejection of the null hypothesis constitutes statistical evidence that a long-run relationship between p and p^* might exist. Using an Engle-Granger procedure, however, we cannot reject the null hypothesis at a 5 percent level of significance that p and either specification of p^* are not cointegrated. While this result casts some doubt on the major premise of the P^* approach for Japan, it might also be a reflection of our sample size and the low power of the augmented Dickey-Fuller test. Nonetheless, P^* may still provide a useful benchmark in the forecast of inflationary pressures for Japan.

P^* as an Inflation Indicator for Japan

If P^* is to be useful as an inflation indicator, the price gap, defined as the difference between the log of the GNP deflator and p^*, should vary inversely with the actual inflation rate. When the price gap declines, the model predicts prices will tend to rise faster or fall slower toward the equilibrium, p^*. When the gap increases, prices should tend to rise slower or fall faster. In chart 8, which uses the assumption of a constant wealth-income ratio, one can see that the gradual narrowing of the price gap corresponds to the general decline in the rate of inflation over the sample period. In addition, corresponding to the sharp rise in inflation in 1973 associated with the oil shock is a large decline in the price gap. Furthermore, as the price gap approaches zero and becomes negative in the late 1980s, inflation in Japan appears to have turned upward. Chart 9 shows a similar pattern under the trend assumption for w^*.

Because in this study we wish to investigate the usefulness of the P^* approach described by HPS, we have restricted our search for suitable specifications of the relationship between the price gap and inflation to types they considered. The particular dynamic specification we settled on for Japan is given by:

$$(19) \quad \Delta\pi_t = \kappa + \alpha(p_{t-1} - p^*_{t-1}) + \sum_{j=1}^{4} \beta_j \Delta\pi_{t-j},$$

where $\pi_t = p_t - p_{t-1}$ and $\Delta\pi_t = \pi_t - \pi_{t-1}$.[6] The coefficient estimates for this model when p^* and v^* are calculated under the assumption of a constant wealth-income ratio are shown in the column marked "price-gap model" of table 3, along with a variety of diagnostic statistics. While the coefficient on the price-gap term has a negative sign, it is not significantly different from zero at standard levels of significance.

The unrestricted version of this model is:

$$(20) \quad \Delta\pi_t = \kappa + \alpha_1(v_{t-1} - v^*_{t-1}) + \alpha_2(y_{t-1} - y^*_{t-1}) + \sum_{j=1}^{4} \beta_j \Delta\pi_{t-j} + \gamma\pi_{t-1}.$$

The unrestricted version does not constrain the coefficients on the output and velocity gaps to be equal in magnitude but opposite in sign.

6. Stephen O'Connell has pointed out to us that the basic specification used by HPS neglects the possibility that changes in inflation might also be explained by accelerations in p^* (i.e., by $\Delta\pi^*$, using notation analogous to that described above). Presumably, when $p=p^*$, inflation may still change as p^* accelerates.

Furthermore, it includes the lagged level of the inflation rate to allow a test of the hypothesis that the dependent variable should be the level of inflation rather than the change in the level of the inflation rate. Based on the t-statistic on γ, the change specification appears to be favored by the data. However, the coefficient on the velocity gap is insignificant and of the wrong sign. An F-test of the hypothesis that both restrictions are valid is rejected at the 1 percent level of significance.

This specification of p^* fails to generate a price-gap model that performs as well in sample as the more traditional output-gap model:

$$(21) \quad \Delta\pi_t = \kappa + \alpha_2(y_{t-1} - y^*_{t-1}) + \sum_{j=1}^{4} \beta_j \Delta\pi_{t-j}.$$

The coefficient on the output gap is positive and significant, and an F-test of the hypothesis that the coefficients on π_{t-1} and the velocity gap are zero is not rejected.

Table 4 shows coefficient estimates for the price-gap and unrestricted models when p^* is constructed under the trend wealth assumption. For comparison purposes, it also repeats the coefficient estimates for the output-gap model. The price-gap model performs much better in sample under this specification. The coefficient on the price gap is negative and significant, and an F-test that the restrictions are valid is not rejected. While this test result indicates that the restrictions imposed by the price-gap model appear to be satisfactory, the restrictions imposed by the output-gap model, one of which is that the coefficient on the velocity gap is zero, are not rejected either.

Following Hallman, Porter, and Small, we subjected all these models to a battery of diagnostic tests. The results are shown in the tables. For all specifications, serial correlation of various orders does not appear to be a problem. Tests for heteroskedasticity in which squared residuals are regressed on the right-side variables and their squares, on fitted values, and on a trend do not indicate problems either. Tests for including one further lag in the dependent variable and for including one further lag in the dependent variable and the other right-side variables indicate that the dynamic specification of the models appears to be adequate.

The one diagnostic that does indicate a problem is the Fisher [1922] test that the coefficients estimated over the first half of the sample stay the same when estimated over the second half. The price-gap model with the assumption of a constant wealth-income ratio appears to perform particularly badly in this respect, although coefficient stability in the price-gap model under the trend-wealth assumption and in the output-gap models is also rejected at the 1 percent levels. Recursive estimation shows that the estimates of coefficients for the price gaps and for the constant tend to move significantly over time, possibly indicating a change in the structure of the inflation process in Japan from the 1970s to the 1980s.

The models were also evaluated with other tests of parameter constancy (based on Chow [1960]) that rely on the forecasting ability of the specifications. Chart 10 shows the results of a sequence of F-tests based on rolling estimation and 1-step-ahead forecasts. The F-statistics are scaled by critical values at the 5 percent level so that this level of significance is represented by a straight line at unity. All three specifications break down in 1980. A possible explanation for this breakdown is that potential output dropped sharply immediately following the 1979 oil price shock but the trend measure of potential output use in the analysis fails to account for this decline.

Some traditional gauges of the forecasting accuracy of these specifications are presented in table 5. Each of the three models was put through three sequences of estimation and forecasting. We estimated the models over sample periods beginning in 1971Q3 and ending in 1986Q4, 1987Q4, or 1988Q4. Coefficients from the estimation over the short sample were then used to calculate forecasts over a three-year horizon. Similarly, coefficients from the estimation over the middle and long samples were used to calculate forecasts over two- and one-year horizons, respectively. Three types of summary statistics on the forecast errors are shown in the table: average error, mean absolute error, and root mean squared error.

In general, all three models tend to forecast $\Delta\pi$ too high, as evidenced by the uniformly negative average errors. The output-gap model performed the best in terms of root mean squared error at all horizons and in terms of mean absolute error at the three-year horizon. The price-gap model under the assumption of a constant wealth-income ratio performed the best in terms of mean absolute error at the two- and one-year horizons, although the margin between this model and the output-gap model in terms of mean absolute error was relatively small. On the other hand, the price-gap model that uses the trend-wealth assumption performed the worst of the three models across all horizons and all measures, with root mean squared errors and mean absolute errors on the order of 20 to 30 percent larger than the best performer.

To sum up the results of the P^* analysis for Japan, the in-sample and out-of-sample performances of the price-gap models are mixed. In sample, the model using the price gap calculated under the trend-wealth assumption performs better than the alternative price-gap model. However, out of sample, its performance is decidedly worse. On the other hand, the output-gap model appears to perform as well or better than either of the price-gap models both in and out of sample.

<u>Constructing P^* for Germany</u>

The first step in the construction of P^* for Germany was the specification of a long-run money demand function that could be used to calculate V^*. Although there exists a literature on the demand for German M3, there is little work that has included wealth or the "own" rate of interest as explanatory variables. For purposes of comparison, we started with a specification similar to Corker's [1989] for Japanese M2+CDs. This specification allows the demand for broad money to be influenced by both transactions and portfolio motives. Chart 11 shows that, as is the case for Japanese M2+CDs, the income velocity of German M3, the main targeted aggregate in Germany, has exhibited a secular decline since 1970. One factor underlying this trend is the rise in the wealth/income ratio between 1970 and 1989, as shown in the bottom half of Chart 11.

Between 1970 and 1989, the ratio of wealth to GNP rose from 2.9 to 4.25 (about 46 percent) while the GNP velocity of M3 fell from .62 to .48 (about 23 percent). Ceteris paribus, these changes would imply an elasticity of M3 velocity with respect to the the wealth/income ratio of about 1/2. In Japan, the wealth/income ratio almost doubled between the

1970 and 1989 while the GNP velocity of M2+CDs fell about 40 percent, which implies an elasticity similar to the German one.[7]

To some extent, the rise in the wealth/income ratio in Germany and Japan is a result of the postwar reconstruction that led to a period of very rapid growth in these countries. Other factors, such as demographic patterns, have also played a role. The aging of the populations in both Germany and Japan has caused national savings rates to be high in these countries during the last few decades.

Unlike Japanese savings instruments, German savings accounts have had market-related interest rates for some time. All regulations of German interest rates were abolished in 1967, reducing the incentive for financial innovations such as those which changed the nature of the relationship between Japanese interest rates during the sample period. Chart 12 shows the evolution of the interest rate on public sector bonds (the rate relevant for marginal portfolio decisions) and the "own" rate of interest on German M3. The "own" rate on M3 was constructed as a weighted average of the rates on various components of M3, where the weights were the proportion of M3 accounted for by each component.[8]

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7. The U.S. net worth/GNP ratio has remained relatively stable during this period, which may be one explanation behind the apparent stability of U.S. M2 velocity.

8. German M3 consists of M1, notes and coin plus sight deposits (which do not bear interest in Germany), plus time deposits with maturities less than 4 years (TD), plus savings deposits at statutory notice (SD). The own rate of return on M3 was defined as:

$$i_o = (TD/M3)*i_{TD} + (SD/M3)*i_{SD}$$

where i_{TD} is the rate on 1-3 month time deposits and i_{SD} is the average rate of interest on savings deposits at statutory notice. There are 3 published rates for time deposits with agreed maturities of 1 to 3 months: those for deposits of less than DM 100,000, from DM 100,000 to DM 1 million, and from DM 1 million to DM 5 million. Because a breakdown

(Footnote continues on next page)

To calculate the long-run relationship between interest rates in Germany, we estimated the following equation from 1970Q2 to 1989Q4:

$$(22)\quad i_{o,t} = \underset{(-.87)}{-.002} + \underset{(16.50)}{.854 i_{o,t-1}} + \underset{(2.10)}{.076 i_{pb,t}}$$

SER = .00423 $\bar{R}^2 = .814$ Durbin's h = -.546

The long-run equilibrium relationship implied by equation (22) is:

$$(23)\quad i_o^* = -.016 + .518 i_{pb}^*$$

The relationship between the two interest rates described by equation (22) is stable over most of the sample period. The bottom panel of Chart 12 shows the 1-step-ahead Chow statistics scaled by the critical values of the F distribution at a 5 percent level of significance. There is only one rejection, in the second quarter of 1975; for the rest of the sample the Chow tests do not reject the hypothesis of parameter stability.

(Footnote continued from previous page)
of time deposits into size of deposit was unavailable, we used the rate for the intermediate size of deposits as a proxy for the average rate of return on all time deposits. We could not adjust for the maturity of time deposits, because average rates on deposits with maturities from 3 months to 4 years were not readily available. About two thirds of all time deposits are in the short-term category, but the lack of inclusion of the more medium-term rates may introduce a degree of measurement error into i_o if the longer-term time deposits are closer substitutes to public bonds.

To estimate the demand for German real balances, we utilized a dynamic specification of the error-correction form. The dependent variable was the change in the log of real M3 balances, $\Delta(m-p)$. Included in the equation were: 4 lags of the dependent variable, the change in log real GNP (Δy) and several of its lags, the change in log real wealth (Δw), the inflation rate as measured by the rate of change of the log GNP deflator (π), various combinations of the interest rates and their rates of change, and two error correction terms. The error correction terms used were lags of the ratio of money to GNP (m-p-y) and the ratio of money to wealth (m-p-w).

The variables that had the most explanatory power were the error correction term relating money and income, the constant, the level of inflation and the lagged level of the interest rate on public bonds. The level of the own interest rate never entered the equation significantly, and the rate of change of real GDP and wealth and their lags usually were insignificant and entered with the wrong sign.[9] Specifications including the interest rate differential and its rate of change did worse than those including the rate on public bonds. The parsimonious versions of these results are presented in table 6.

9. We also estimated a simple partial adjustment model in which the dependent variable, real M3, was regressed on its own lag, a constant, i_{pb}, i_o, real GNP (y), and the wealth/GNP ratio (w-y). All variables were in logs except for interest rates which entered in levels or in the form ln(1+i). In these regressions, the own interest rate entered with the wrong (negative) sign, albeit an insignificant one. This could be due to measurement error in that a weighted-average measure of the return on M3 may not represent the marginal return on the asset that is relevant for shifts into and out of M3. Using the level of an interest rate instead of ln(1+i) made little difference. Both idiff = $i_{pb} - i_o$ and ln(1+idiff) entered significantly with the correct sign although the fit worsened slightly, as did Durbin's h statistic.

The first column presents the case for the best specification we found. Both error correction mechanisms enter with the correct sign, and the lagged level of i_{pb} and inflation are significant. The steady-state solution of this equation (i.e. when $\Delta(m-p) = 0$) leads to the following long-run demand for money:

$$(24) \quad m - p = \tilde{\Theta}_1 + \Theta_2 w + \Theta_3 y - \tilde{\lambda}\ln(1+i_{pb}) - \phi\pi$$

where $\tilde{\Theta}_1 = .28$, $\Theta_2 = .4$, $\Theta_3 = .6$, $\tilde{\lambda} = 1.09$ and $\phi = 5.15$.[10]

The second column of the table shows that using the differential between i_{pb} and i_o instead of i_{pb} alone slightly worsens the fit of the equation, although the major parameters of interest do not change substantially.

The last column addresses the problem of the level of inflation entering the dynamic equation significantly. The inclusion of inflation in the long-run money demand relationship above complicates the p^* analysis by making p^* a second-order process.[11] To avoid a messier solution algorithm, we tried an additional specification for German money demand that depends on the change in inflation rather than the level. As the last column indicates, this variable seems to do fairly well as a proxy for π. This variable falls out of the long-run solution, so that

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10. These values were similar to the long-run parameters of money demand that were implied by the simple partial adjustment regressions. The implied long-run elasticities are reasonable and seem to be in the range of those found in previous studies.

11. Equation (12) becomes:

$$p^* = m - \Theta' z^* + \lambda(r_g^* + 4(p_{t+1}^* - p^*)) + \phi(p^* - p_{t-1}^*)$$

In some specifications of the demand for M3, an additional lag of inflation entered the equation, which would make the process a third-order one.

we can continue to use the first-order solution method presented for the case of Japanese P^*. With this restriction on the model, the long-run parameters of interest are: $\tilde{\Theta}_1 = -.01$, $\Theta_2 = .57$, $\Theta_3 = .43$, and $\tilde{\lambda} = .98$

This parameterization was used for the calculation of P^* for Germany. Note that we are using the money demand specification that excludes i_o, which is equivalent to setting $\alpha = \beta = 0$. In addition, beyond the fourth quarter of 1989, we assume M3 growth at an annual rate of 6 percent, and real GNP growth of 4 percent. Real wealth is also assumed to grow at 4 percent so that the wealth/income ratio remains at 4-1/4. The real interest rate was set equal to 4 percent for the entire sample and was assumed to remain there susbequently.

P^* as an Inflation Indicator for Germany

Chart 13 shows the evolution of inflation along with the price gap, $p-p^*$. Note that the gap was negative, $p < p^*$, for most of the 1970s and then positive during almost all of the 1980s. The latter period was one of significant disinflation in Germany; the 4-quarter rate of inflation fell from 5.5 percent in the second quarter of 1980 to 2.2 percent in the last quarter of 1989. Chart 14 decomposes the price gap into its components: $y-y^*$ and $v-v^*$. The output gap turned negative in 1981 and did not reenter the positive range until 1985, which undoubtedly contributed to the German disinflation in the last decade. However, between mid-1987 and the end of 1989, German real GNP has grown more rapidly than potential, without being manifested in an appreciable rise in inflation (at least in sample). As the lower panel shows, the widening of the velocity gap more than offset the movements in the output gap. There is a clear positive trend in the velocity gap during the sample period. This trend is due in part to the measure of real wealth we have used to proxy for w^*. As with Japan, we calculated w^* using a linear trend estimated for the period between 1960Q1 and 1989Q4:

$$(25) \quad w^* = 5.891 + 0.0129*(\text{time})$$

The top panel of Chart 15, labelled Case A, shows the actual and fitted values of real wealth that correspond to equation (25). This measure of w^* seems to underestimate real wealth in the 1970s, but exceeds actual real wealth from 1983 on. Using a simple time trend to describe the evolution of real wealth may be one source of the trend in the velocity gap. The growing tendency to overestimate equilibrium wealth over the sample could be associated with an increasing tendency to underestimate v^* and overestimate the velocity gap.

The trend in the velocity gap could also be a result of an omitted variable in the long-run money demand relationship we have estimated. Some of the literature on German money demand indicates that for certain periods, foreign interest rates (specifically U.S. rates) and the rate of change of the DM's exchange rate are significant explanatory variables in dynamic equations for German money demand. Whether these variables are important to long-run money demand and thus, to v^*, is beyond the scope of this paper. However, if we have omitted variables that should have been included in our specification, they may have a trend which contributes to the trend in the velocity gap.

When we tried the HPS unrestricted specifications (see equation 21) for the influence of the price gap on inflation or the rate of change of inflation, we were always able to reject those specifications with $\Delta\pi$

as the dependent variable. In other words, π_{t-1} always entered the equations significantly. Therefore, we settled on specifications where the dependent variable was the rate of inflation. The unrestricted equation we estimated was:

$$(26) \quad \pi_t = \kappa + \alpha_1(v_{t-1} - v^*_{t-1}) + \alpha_2(y_{t-1} - y^*_{t-1}) + \sum_{j=1}^{4} \beta_j \pi_{t-j}$$

Table 7 presents the results.

The first column presents the results for the price-gap model. This is the restricted version of equation (26) where $\alpha_1 = -\alpha_2$. The price gap is significant and has the correct negative sign. The equation explains almost a third of the quarterly variation of inflation over the sample period and passes all of the diagnostic tests, except for two tests for heteroskedasticity. When the squared residuals are regressed on a constant and either a trend or $\hat{\pi}$, the tests reject the hypothesis of independence. The χ^2 statistic is significant at the 5 percent level for the regression on fitted values of π, and at the 1 percent level for a regression on a time trend. The evidence of heteroskedasticity indicates that we have not achieved the most efficient estimate of our coefficients. The fact that the squared residuals have a negative trend could mean that there is an omitted variable (possibly the foreign interest rate or the exchange rate) that was important early in the sample, that becomes less important over time.

The second column presents estimates for the unrestricted model. Both the velocity gap and the output gap enter significantly with the correct (opposite) signs. The standard error of the equation is only slightly lower than that of the price-gap equation, and some of the diagnostic statistics indicate a slightly less efficient specification. Although the absolute values of the coefficients on these gaps do not at first glance appear to be equal, the F-test of this equation versus the restricted form in column 1 could not reject the restriction. Finally, the third column presents results for the output-gap model. This model fits somewhat worse than either of the previous two models, indicating that the P^* analysis has improved our ability to explain German inflation in sample.

Next we repeat the analysis for the case where long-run money demand and V^* depend on the interest rate differential rather than i_{pb}. As shown in table 8, the estimates for coefficients of interest in the unrestricted and price-gap models did not change significantly, but the standard errors of the regressions were slightly larger. The velocity gap is significant at the 10 percent level, but not at the 5 percent level as before. In the remainder of the paper, we drop the interest rate differential from the analysis, because of the somewhat mysteriously better fit achieved by excluding it.

As in the Japanese analysis, the results of the German P^* equations depend crucially upon what is assumed about the equilibrium path of wealth or w^*. In an attempt to explain some of the trend in the velocity gap discussed above and to examine the robustness of the results to different assumptions, we reestimated the w^* equation over a shorter time period that coincides with our sample of 1970Q1 to 1989Q4. This yielded the following relationship:

$$(27) \quad w^* = 6.086 + 0.0109*(\text{time})$$

The actual and fitted values of real wealth corresponding to this equation are shown in the bottom panel of Chart 15, labeled Case B.

Table 9 presents the results for the P^* analysis using the different assumption on w^*. It is evident that the fit of the equations has worsened and that both the price-gap and velocity-gap variables in the unrestricted form enter less significantly into the equations. The coefficient estimates on lagged levels of inflation appear to have changed the most. Judging by the size of the standard errors, there is virtually no improvement of the P^* analysis over the output-gap model. One cannot reject the restrictions implied by the price-gap model compared to the unrestricted version, but neither can one reject the hypothesis that the velocity gap has nothing to add to the output-gap specification.

When we performed augmented Dickey-Fuller tests to determine whether p and p^* were cointegrated, the results were inconclusive. We could not consistently reject the hypothesis of a unit root for the residuals from the regression of p^* on p (or vice versa), a constant and a trend. However, we came closer to rejecting the null hypothesis than in the Japanese case, and, for certain lag structures, we did reject at a 5 percent level of significance. Perhaps owing to the low power of these tests, the results did not differ substantially across the different measures of p^*.

Chart 16 compares the two versions of the price-gap model and the output-gap model in terms of Chow tests based on a sequence of F-tests calculated from rolling estimation and 1-step-ahead forecasts. Again, the F-statistics are scaled by critical values so that the 5 percent level of significance is represented by a straight line at unity.

Neither of the two price-gap specifications break down, providing some evidence of parameter stability. The output-gap specification has only one rejection in late 1978.

Table 10 presents information on the forecasting performance of the different indicators for German inflation. We estimated the three models (the price-gap model with the two different trend wealth assumptions and the output-gap model) over sample periods beginning in 1971Q3 and ending in 1986Q4, 1987Q4, or 1988Q4. The estimated coefficients were used to calculate 12-quarter, 8-quarter and 4-quarter forecasts, respectively. These forecast errors are summarized by the statistics shown in the table: average error, mean absolute error, and root mean squared error.

The table shows that the price-gap model with w^* estimated over the longer sample (case A) had the best forecast performance across all three horizons and based on all of the different criteria. Over the 12-quarter horizon, all of the models tended to overpredict inflation, as evidenced by the uniformly negative average errors. However, the price-gap model (case A) had a root mean squared forecast error 15 percent below the price-gap model (case B) and 33 percent below that of the output-gap model. At the shorter forecast horizons, the first price-gap model tended to slightly underpredict inflation in contrast to a continuation of the tendency to overpredict by the other two models. The difference between the root mean squared errors associated with the two versions of the price-gap model diminished somewhat: case A had errors about 8 percent less than case B over both the two- and one-year horizons. In contrast, the forecasting performance of the price-gap models and the output-gap models further diverged over the shorter horizons. For both the 4-quarter and 8-quarter forecasts the output-gap model had root mean squared errors more than twice those associated with case A of the price-gap model. The output-gap model performed the worst of the three models across all horizons and all measures, indicating some degree of gain in forecasting inflation as a result of the P^* approach.

In summary, the results of the construction of a P^* measure for Germany indicate that using trend wealth estimated from 1960Q1 to 1989Q4 as a key variable underpinning V^* performs better, both in and out of sample, than the alternative measures considered. In addition, the long-run money demand specification that is favored seems to be one that uses the interest rate on public bonds, rather than the differential between i_{pb} and i_o as the opportunity cost of holding M3. The German analysis differs from both the HPS results for the United States and the Japanese results we discuss above because, it is the level of inflation, rather than its rate of change, that the price gap influences most strongly. This has interesting implications for the different dynamic behavior of inflation across countries with respect to inflationary shocks.

Conclusion

The P^* approach worked well for Germany in that it improved our ability to explain or forecast inflation during the period of 1971Q3 to 1989Q4. However, our approach did not yield a consistently superior forecasting tool for changes in inflation in Japan. The simpler and more traditional output-gap approach to forecasting inflationary pressures appears to work as well or better than the Japanese P^* approaches we tried.

When V^* is not constant, the attractiveness of the P^* approach to forecasting inflationary pressures diminishes. We have adapted the simple model of P^* put forth by Hallman, Porter, and Small to allow for fluctuations in V^* consistent with models of money demand and perfect foresight. These adaptations increased the complexity of the approach in terms of calculation of P^* and in making assumptions about the future long-run equilibrium paths of the determinants of money demand and velocity.

The assumption that the income velocity of U.S. M2 is constant gives a very special case in which P^* is easily calculated and a useful indicator for analyzing and forecasting inflation. In this paper, we have shown that trends in the wealth/income ratio in Japan and Germany are key determinants behind the trends in income velocity in these countries, which are crucial to the construction of P^*. One possible extension to the paper is to formalize a model of the evolution of wealth based on demographics and other economic factors in order to more fully understand these trends. Another extension would be to move from a partial equilibrium to a general equilibrium model of P^*. However, the design of a general equilibrium model of the aggregate price level is a tall order and would move us even further away from the usefulness of P^* as a simply constructed leading indicator of inflation.

DATA APPENDIX

Japan

Much of the data used for the calculation of Japanese P^* was kindly provided by Robert Corker. We updated his sample through the end of 1989 using various issues of the Bank of Japan's Economic Statistics Monthly. All of the series were seasonally adjusted except for the interest rates.

Real money balances were defined as the stock of M2+CDs deflated by the GNP deflator, wealth consisted of financial wealth defined from flow-of-funds data as the average of beginning and end period stocks of the sum of the total positive financial assets of the personal and corporate sectors. The opportunity cost of holding money was defined as the 3-month Gensaki rate minus the average return on holding money, a weighted average of the interest rates on 3-month CDs and the guideline 3-month deposit rate. (See Corker (1989) for a detailed description of the construction of the own rate.) A regression of real GNP from 1965Q2 to 1989Q4 on a constant and a trend, broken in 1973Q4 to allow for the oil price shock, was used to generate a series for potential output:

$$y^* = 11.8 + .021*(\text{time}) - .011*(\text{trendshift})$$

where trendshift was a dummy variable that was zero until the fourth quarter of 1973 and then set equal to a trend thereafter.

Germany

Most of the data used to estimated German money demand and to calculate P^* came from various issues of the Monthly Report of the

Deutsche Bundesbank. All of the data were seasonally adjusted except for the interest rates.

The construction of the own rate on German M3 was discussed in the text. The interest rate on public-sector bonds that was used was the monthly average secondary market yield on all public-sector bonds where the yields of redeemable issues were based on the mean remaining maturity and group yields were weighted by the amounts issued. This interest rate series was used because it was the one for which the largest number of observations was available.

The measure of wealth was constructed from flow-of-funds data published In the May and October issues of the Monthly Report of the Deutsche Bundesbank. We started with the net assets (excluding housing) of the household sector at the end of 1989, and generated a series of the stock of wealth back to 1960 using data on the financial surplus/deficit of households (or the change in net assets). Since this data is only available on a biannual basis, we interpolated to give us an admittedly smooth version of quarter household wealth. This series was deflated by the GNP deflator to give us a measure of real wealth.

Finally, we seasonally adjusted a measure of German potential output kindly provided by Karl-Heinz Toedter at the Bundesbank. The Bundesbank measure, which is used in the econometric model of the Bundesbank, is based on estimates from a CES production function.

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Table 1

OLS Estimates for α and β
(t-statistics in parentheses)

Sample Period	α	β	$\bar{R}^2$	SER
70Q3 - 89Q4	0.015 (11.9)	0.152 (9.1)	0.517	0.0039
85Q1 - 89Q4	0.006 (1.4)	0.337 (4.0)	0.468	0.0039
85Q1 - 86Q4	-0.003 (-1.73)	0.456 (13.2)	0.961	0.0009
87Q1 - 89Q4	-0.021 (-4.3)	1.009 (8.8)	0.873	0.0020

Table 2

Parameter values for different values of α and β

Specification #	α	β	Θ_1	Θ_2	λ_0
1	0.015	0.152	0.558	0.535	1.321
2	0.006	0.337	0.544	0.535	1.024
3	-0.003	0.456	0.530	0.535	0.832
4	-0.021	0.894[12]	0.503	0.535	0.160
5	0.0	1.0	0.535	0.535	0.0

12. The fact that the coefficient estimate of 1.009, shown in table 1, is slightly greater than one, indicates the possibility that the interest rate on M2+CDs may have become more responsive by the end of the sample than the gensaki rate to shocks to the Japanese money market, thus reversing the roles of the two interest rates described in the text. However, estimates of one or somewhat less are not statistically distinguishable from the 1.009. So, to avoid complicating further the theory and to present an interesting alternative to specification 5, we have set β in this specification one standard deviation below its estimated value of 1.009.

Table 3a

Estimates for Japan
of Price Gap Model under the Assumption of Constant Equilibrium
Wealth-Income Ratio and of Output Gap Model
(t-statistics in parentheses)

Variable	price-gap model, equation 20	unrestricted model, equation 21	output-gap model, equation 22
Intercept	.002 (.91)	-.001 (-0.6)	.0003 (.47)
$p_{t-1}-p^*_{t-1}$	-.00686 (-1.1)	...	...
$v_{t-1}-v^*_{t-1}$	...	.0113 (1.3)	...
$y_{t-1}-y^*_{t-1}$	...	.112** (3.3)	.118** (3.8)
$\Delta\pi_{t-1}$	-.264* (-2.3)	-.272* (-2.1)	-.355** (-3.3)
$\Delta\pi_{t-2}$	.271* (2.3)	.290* (2.3)	.211 (1.9)
$\Delta\pi_{t-3}$	-.115 (-.99)	-.118 (-.92)	-.198 (-1.8)
$\Delta\pi_{t-4}$	-.252* (-2.2)	-.249* (-2.1)	-.303** (-2.9)
π_{t-1}	...	-.117 (-1.2)	...
Regression statistics			
$\bar{R}^2$	.172	.304	.305
Standard error	.00670	.00614	.00614
Sample	71Q3 - 89Q4	71Q3 - 89Q4	71Q3 - 89Q4
Sample size	74	74	74

* Significant at the 5 percent level
** Significant at the 1 percent level
T-tests on lagged $\Delta\pi$ terms are two-tailed; all others are one-tailed.

Table 3b

Diagnostic Statistics for Price Gap Model under the Assumption of a Constant Equilibrium Wealth-Income Ratio and for Output Gap Model
(degress of freedom in parentheses)

	price-gap model, equation 20		unrestricted model, equation 21		output-gap model, equation 22	
Diagnostic statistics						
Serial correlation						
1st-order	.044	(1)	1.748	(1)	1.845	(1)
4th-order	.016	(1)	.480	(1)	.489	(1)
1st- to 4th-order	2.577	(4)	4.924	(4)	4.476	(4)
1st- to 8th-order	6.777	(8)	6.947	(8)	6.538	(8)
Heteroskedasticity						
With ind. vars.	14.34	(11)	19.92	(15)	10.76	(11)
With fitted values	1.07	(1)	.86	(1)	1.13	(1)
With trend	1.74	(1)	.92	(1)	.54	(1)
One further lag in						
$\Delta\pi$	.39	(1)	.49	(1)	1.360	(1)
all r.h.s. vars.	5.55	(2)	3.62	(4)	2.842	(2)
Fisher test	6.01**	(6,62)	3.58**	(8,58)	3.214**	(6,62)

F-tests

hypothesis	statistic	
$\alpha_1 = -\alpha_2, \quad \gamma = 0$	7.45**	(2,66)
$\alpha_1 = 0, \quad \gamma = 0$	0.97	(2,66)

** Significant at the 1 percent level

Table 4a

Estimates for Japan
of Price Gap Model under the Trend Wealth Assumption and of Output Gap Model
(t-statistics in parentheses)

Variable	price-gap model, equation 20	unrestricted model, equation 21	output-gap model, equation 22
Intercept	.003** (2.5)	.002 (1.4)	.0003 (.47)
$p_{t-1}-p^*_{t-1}$	-.0513** (-3.6)	...	...
$v_{t-1}-v^*_{t-1}$	...	-.0282 (-1.3)	...
$y_{t-1}-y^*_{t-1}$	...	.0894* (2.3)	.118** (3.8)
$\Delta\pi_{t-1}$	-.371* (-3.3)	-.341* (-2.8)	-.355** (-3.3)
$\Delta\pi_{t-2}$	.148 (1.3)	.191 (1.6)	.211 (1.9)
$\Delta\pi_{t-3}$	-.174 (-1.6)	-.157 (-1.3)	-.198 (-1.8)
$\Delta\pi_{t-4}$	-.277* (-2.6)	-.264* (-2.3)	-.303** (-2.9)
π_{t-1}	...	-.0555 (-.77)	...
Regression statistics			
$\bar{R}^2$	.291	.304	.305
Standard error	.00620	.00614	.00614
Sample	71Q3 - 89Q4	71Q3 - 89Q4	71Q3 - 89Q4
Sample size	74	74	74

* Significant at the 5 percent level
** Significant at the 1 percent level
T-tests on lagged $\Delta\pi$ terms are two-tailed; all others are one-tailed.

Table 4b

Diagnostic Statistics for Price Gap Model under the Trend Wealth Assumption
and for Output Gap Model
(degress of freedom in parentheses)

	price-gap model, equation 20		unrestricted model, equation 21		output-gap model, equation 22	
Diagnostic statistics						
Serial correlation						
1st-order	.048	(1)	.826	(1)	1.845	(1)
4th-order	.029	(1)	.135	(1)	.489	(1)
1st- to 4th-order	2.585	(4)	3.618	(4)	4.476	(4)
1st- to 8th-order	7.015	(8)	5.082	(8)	6.538	(8)
Heteroskedasticity						
With ind. vars.	8.50	(11)	17.15	(15)	10.76	(11)
With fitted values	.96	(1)	.93	(1)	1.13	(1)
With trend	.23	(1)	.23	(1)	.54	(1)
One further lag in						
$\Delta\pi$	1.42	(1)	1.18	(1)	1.36	(1)
all r.h.s. vars.	2.57	(2)	3.92	(4)	2.84	(2)
Fisher test	3.42**	(6,62)	3.88**	(8,58)	3.21**	(6,62)

F-tests

hypothesis	statistic	
$\alpha_1=-\alpha_2,\ \gamma=0$	1.60	(2,66)
$\alpha_1=0,\ \gamma=0$	0.96	(2,66)

** Significant at the 1 percent level

Table 5

Forecasting Performance of Alternative Inflation Indicators for Japan

	average error	mean absolute error	root mean squared error
12 Quarter forecast			
Price Gap with Constant Wealth-Income Ratio	-0.0036	0.0056	0.0071
Price Gap with Trend Wealth	-0.0057	0.0068	0.0084
Output Gap	-0.0015*	0.0055**	0.0065***
8 Quarter forecast			
Price Gap with Constant Wealth-Income Ratio	-0.0012*	0.0039**	0.0056
Price Gap with Trend Wealth	-0.0039	0.0049	0.0070
Output Gap	-0.0014	0.0040	0.0055***
4 Quarter forecast			
Price Gap with Constant Wealth-Income Ratio	-0.0006*	0.0023**	0.0046
Price Gap with Trend Wealth	-0.0019	0.0029	0.0053
Output Gap	-0.0009	0.0024	0.0044***

---

* Minimum average error across the models

** Minimum mean absolute error across the models

*** Minimum root mean squared error across the models

Table 6

Estimates of German Demand for Real M3, $\Delta(m-p)$, 1970Q2 - 1989Q4

Variable	dynamic model, with i_{pb} and π	dynamic model, with idiff and π	dynamic model, with i_{pb} and $\Delta\pi$
Intercept	.063 (3.34)	.072 (3.610)	...
$\Delta(m-p)_{t-1}$	.002 (.04)	.043 (.56)	.241 (3.04)
$\Delta(m-p)_{t-3}$	.126 (1.70)	.169 (2.30)	.117 (1.65)
$\Delta(m-p)_{t-4}$	.095 (1.30)	.123 (1.65)	.105 (1.48)
$\Delta\ln(1+i_{pb})$	.229 (1.20)	...	.305 (1.69)
$\ln(1+i_{pb})_{t-1}$	-.245 (-3.28)	...	-.274 (-4.80)
$\Delta\ln(1+idiff)$	...	.229 (1.46)	...
$\ln(1+idiff)_{t-1}$	...	-.292 (-3.05)	...
$m_{t-1}-p_{t-1}-y_{t-1}$	-.135 (-6.02)	-.140 (-6.12)	-.115 (-5.68)
$m_{t-1}-p_{t-1}-w_{t-1}$	-.09 (-2.31)	-.074 (-1.94)	-.150 (-6.27)
π	-1.16 (-7.51)	-1.15 (-7.23)	...
$\Delta\pi$	...	...	-.79 (-7.24)
Regression statistics			
$\bar{R}^2$	.581	.559	.766
Standard error	.00660	.00677	.00669
Durbin's h	.877	1.12	-1.310
Sample size	79	79	79

Table 7

Estimates of Equations for Alternative German Inflation Indicators (Case A)

Variable	price-gap model $\alpha_1 = -\alpha_2$ equation (26)	unrestricted model, equation (26)	output-gap model $\alpha_1 = 0$ equation (26)
Intercept	.008 (3.76)	.006 (2.60)	.003 (1.66)
$p_{t-1}-p^*_{t-1}$	-.068 (-3.23)	...	...
$v_{t-1}-v^*_{t-1}$	...	-.049 (-1.97)	...
$y_{t-1}-y^*_{t-1}$	...	.101 (3.20)	.092 (2.88)
π_{t-1}	-.230 (-2.14)	-.231 (-2.16)	-.192 (-1.79)
π_{t-2}	.093 (.93)	.118 (1.17)	.180 (1.99)
π_{t-3}	.138 (1.39)	.191 (1.803)	.279 (2.85)
π_{t-4}	.231 (2.34)	.282 (2.70)	.361 (3.65)
Regression statistics			
$\bar{R}^2$	.313	.322	.294
Standard error	.00539	.00535	.00546
Sample	71Q3 - 89Q4	71Q3 - 89Q4	71Q3 - 89Q4
Sample size	74	74	74
Diagnostic statistics			
Serial correlation			
1st-order	.580 (1)	.848 (1)	.431 (1)
4th-order	.494 (1)	.902 (1)	2.675 (1)
1st- to 4th-order	1.628 (4)	2.113 (4)	3.592 (4)
1st- to 8th-order	2.710 (8)	4.995 (8)	6.871 (8)
Heteroskedasticity			
With ind. vars.	16.70* (11)	19.68* (13)	17.53** (11)
With fitted values	5.76** (1)	5.97** (1)	7.83** (1)
With trend	7.81** (1)	7.54** (1)	7.98** (1)
One further lag in			
π	.05 (1)	.06 (1)	.22 (1)
all r.h.s. vars	.91 (2)	5.68 (3)	4.23 (2)
Fisher test	.80 (6,62)	1.16 (7,60)	1.51 (6,62)

Table 8

Estimates of Equations for Alternative German Inflation Indicators Using a Money Demand Specification Depending on the Interest Rate Differential

Variable	price-gap model $\alpha_1 = -\alpha_2$ equation (26)		unrestricted model, equation (26)		output-gap model $\alpha_1 = 0$ equation (26)	
Intercept	.006 (3.48)		.005 (2.41)		.003 (1.66)	
$p_{t-1}-p^*_{t-1}$	-.076 (-3.20)		...		...	
$v_{t-1}-v^*_{t-1}$	...		-.053 (-1.73)		...	
$y_{t-1}-y^*_{t-1}$	...		.099 (3.13)		.092 (2.88)	
π_{t-1}	-.208 (-1.96)		-.212 (-1.99)		-.192 (-1.79)	
π_{t-2}	.124 (1.29)		.142 (1.45)		.180 (1.99)	
π_{t-3}	.173 (1.79)		.214 (2.07)		.279 (2.85)	
π_{t-4}	.261 (2.69)		.301 (2.92)		.361 (3.65)	
Regression statistics						
$\bar{R}^2$	.311		.314		.294	
Standard error	.00540		.00539		.00546	
Sample	71Q3 - 89Q4		71Q3 - 89Q4		71Q3 - 89Q4	
Sample size	74		74		74	
Diagnostic statistics						
Serial correlation						
1st-order	.193	(1)	.448	(1)	.431	(1)
4th-order	.724	(1)	1.112	(1)	2.675	(1)
1st- to 4th-order	1.483	(4)	1.964	(4)	3.592	(4)
1st- to 8th-order	2.487	(8)	4.297	(8)	6.871	(8)
Heteroskedasticity						
With ind. vars.	15.92	(11)	19.10	(13)	17.53	(11)
With fitted values	5.93*	(1)	6.42*	(1)	7.83**	(1)
With trend	7.52**	(1)	7.45**	(1)	7.98**	(1)
One further lag in						
π	.23	(1)	.03	(1)	.22	(1)
all r.h.s. vars	1.78	(2)	5.84	(3)	4.23	(2)
Fisher test	.80	(6,62)	1.28	(7,60)	1.51	(6,62)

Table 9

Estimates of Equations for Alternative Inflation Indicators (Case B)

Variable	price-gap model $\alpha_1 = -\alpha_2$ equation (26)		unrestricted model, equation (26)		output-gap model $\alpha_1 = 0$ equation (26)	
Intercept	.005 (3.00)		.004 (2.04)		.003 (1.66)	
$p_{t-1}-p^*_{t-1}$	-.088 (-2.87)		...		...	
$v_{t-1}-v^*_{t-1}$	...		-.051 (-1.18)		...	
$y_{t-1}-y^*_{t-1}$	...		.105 (3.11)		.092 (2.88)	
π_{t-1}	-.168 (-1.57)		-.185 (-1.72)		-.192 (-1.79)	
π_{t-2}	.170 (1.78)		.249 (2.48)		.180 (1.99)	
π_{t-3}	.215 (2.23)		.331 (3.25)		.279 (2.85)	
π_{t-4}	.294 (3.03)		.004 (2.04)		.361 (3.65)	
Regression statistics						
$\bar{R}^2$	.293		.298		.294	
Standard error	.00546		.00545		.00546	
Sample	71Q3 - 89Q4		71Q3 - 89Q4		71Q3 - 89Q4	
Sample size	74		74		74	
Diagnostic statistics						
Serial correlation						
1st-order	.139	(1)	.382	(1)	.431	(1)
4th-order	.931	(1)	1.536	(1)	2.675	(1)
1st- to 4th-order	1.772	(4)	2.393	(4)	3.592	(4)
1st- to 8th-order	2.801	(8)	5.015	(8)	6.871	(8)
Heteroskedasticity						
With ind. vars.	16.11**	(11)	18.52**	(13)	17.53**	(11)
With fitted values	6.68**	(1)	6.91**	(1)	7.83**	(1)
With trend	7.89**	(1)	7.67**	(1)	7.98**	(1)
One further lag in						
π	.05	(1)	.33	(1)	.22	(1)
all r.h.s. vars	.39	(2)	5.47	(3)	4.23	(2)
Fisher test	.71	(6,62)	1.55	(7,60)	1.51	(6,62)

Table 10

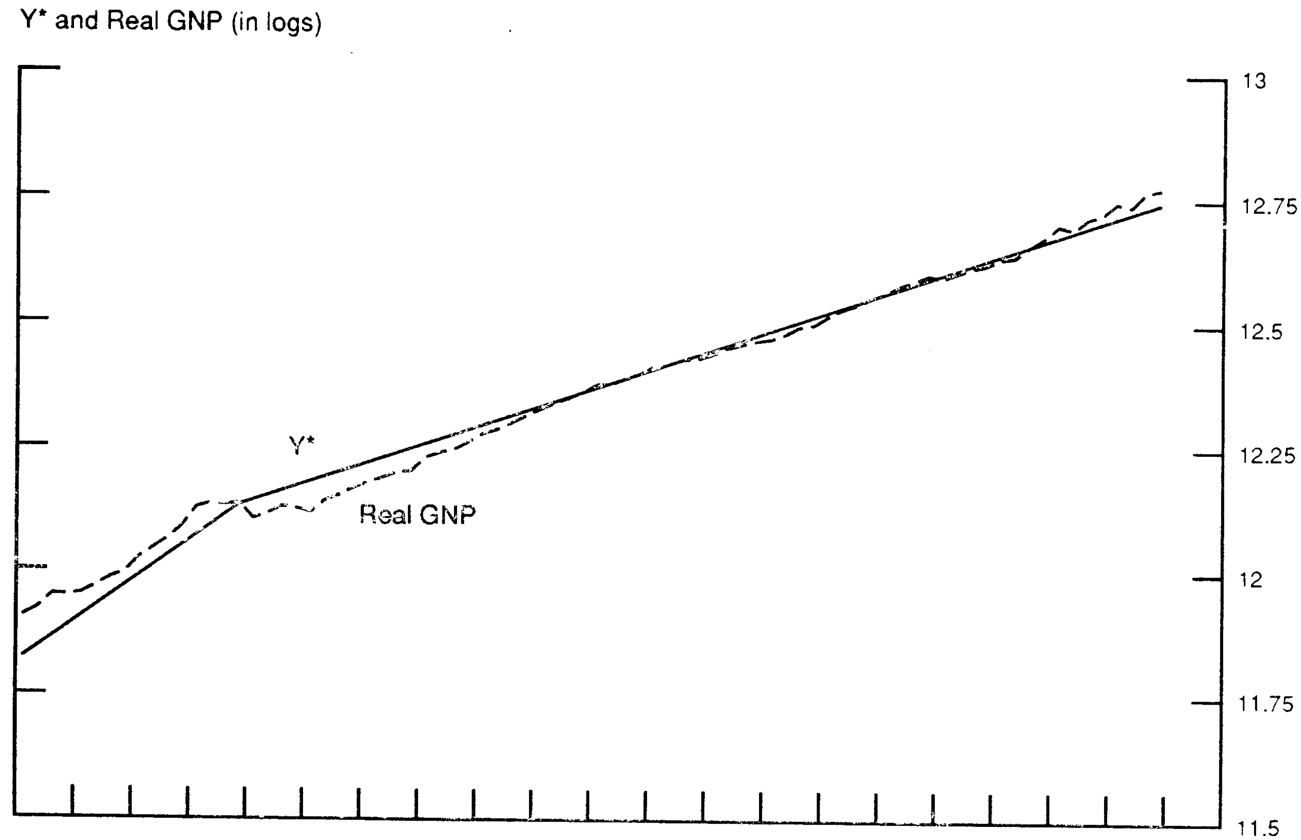
Forecasting Performance of Alternative Inflation Indicators for Germany

	average error	mean absolute error	root mean squared error
12 Quarter forecast			
Price Gap with Trend Wealth Estimated from 1960Q1-1989Q4	-0.0021*	0.0028**	0.0038***
Price Gap with Trend Wealth Estimated from 1970Q1-1989Q4	-0.0031	0.0035	0.0044
Output Gap	-0.0046	0.0048	0.0057
8 Quarter forecast			
Price Gap with Trend Wealth Estimated from 1960Q1-1989Q	0.0001*	0.0016**	0.0021***
Price Gap with Trend Wealth Estimated from 1970Q1-1989Q4	-0.0009	0.0020	0.0022
Output Gap	-0.0028	0.0033	0.0042
4 Quarter forecast			
Price Gap with Trend Wealth Estimated from 1960Q1-1989Q4	0.0005*	0.0018**	0.0025***
Price Gap with Trend Wealth Estimated from 1970Q1-1989Q4	-0.0008	0.0025	0.0027
Output Gap	-0.0040	0.0052	0.0055

* Minimum average error across the models
** Minimum mean absolute error across the models
*** Minimum root mean squared error across the models

Chart 1

JAPANESE POTENTIAL OUTPUT AND WEALTH/GNP



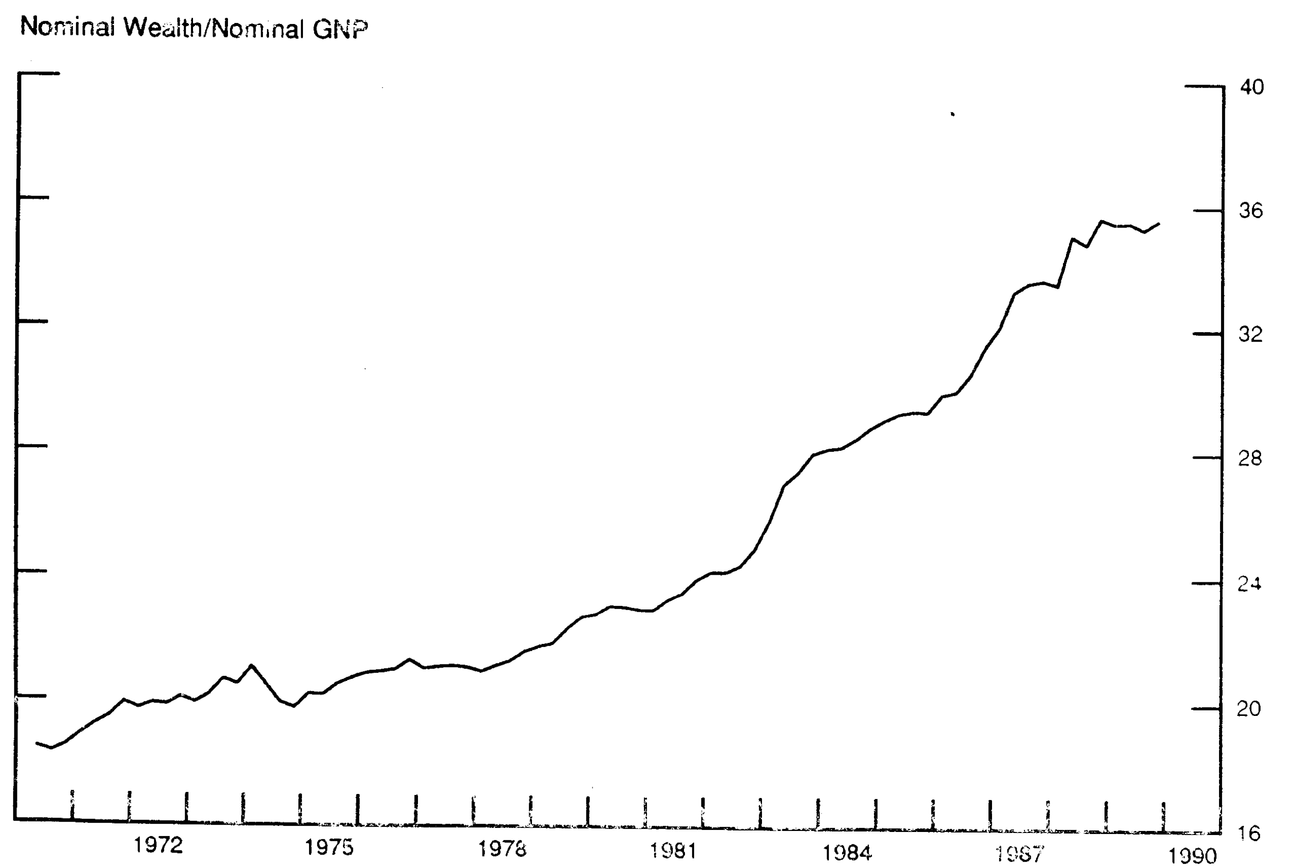


Chart 2

JAPANESE REAL WEALTH
(IN LOGS)

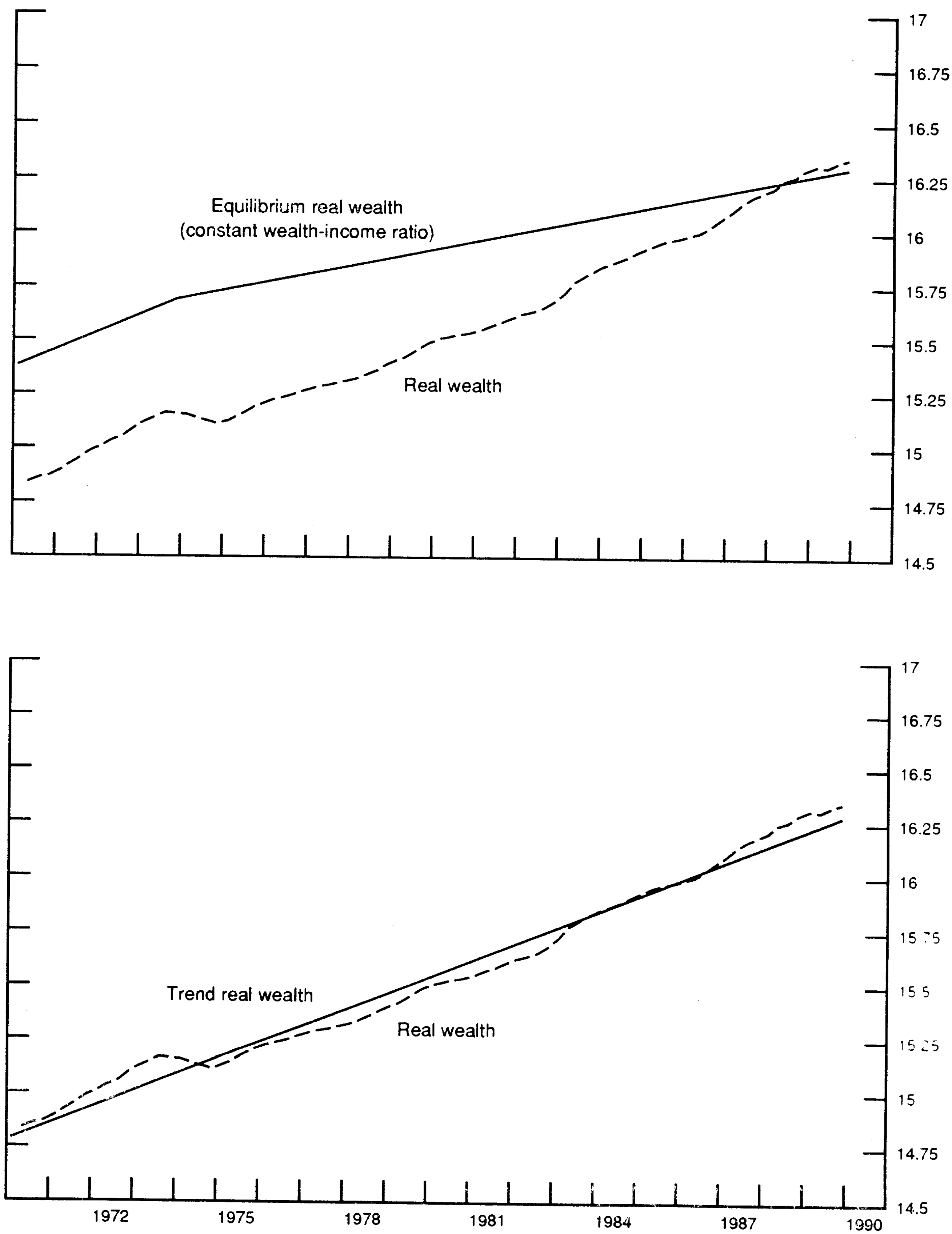


Chart 3

JAPANESE INTEREST RATES

Own Interest Rate on M2+CDs and Gensaki Rate

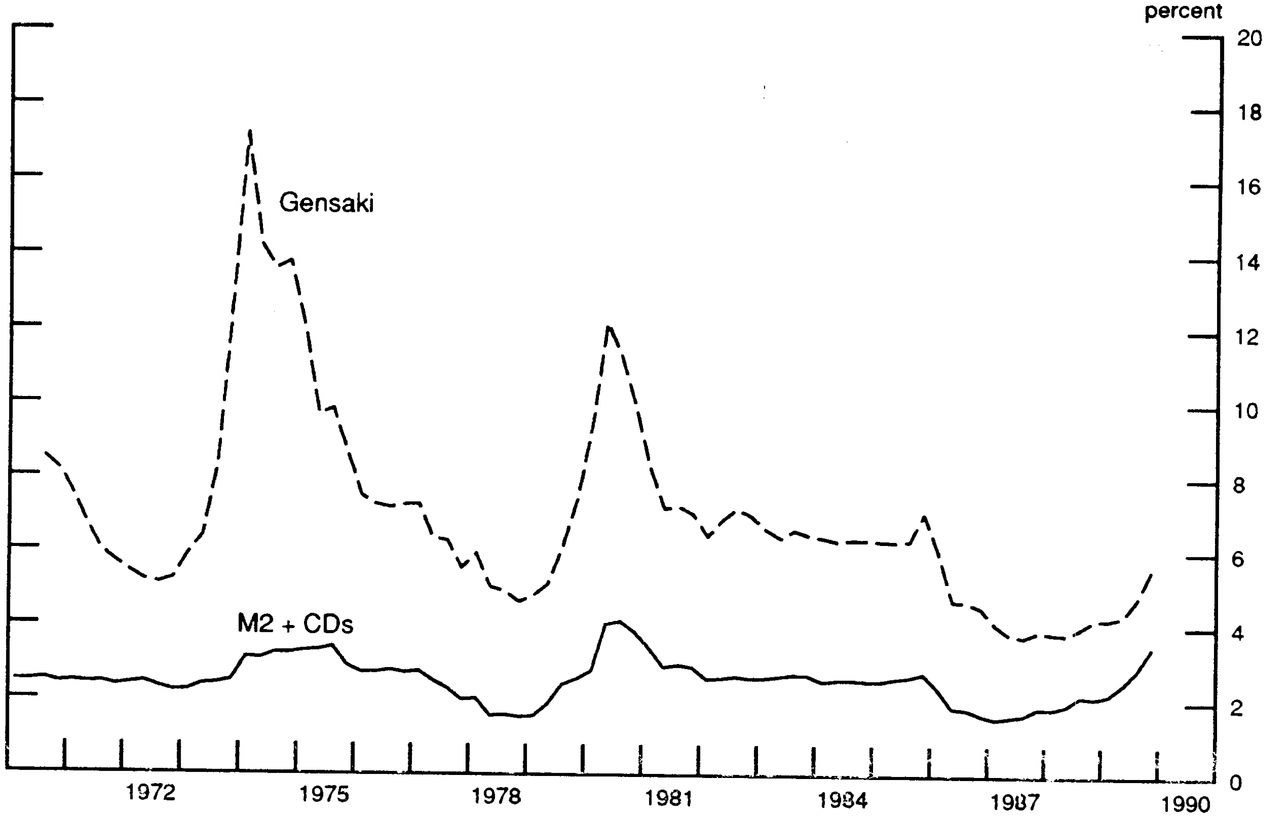
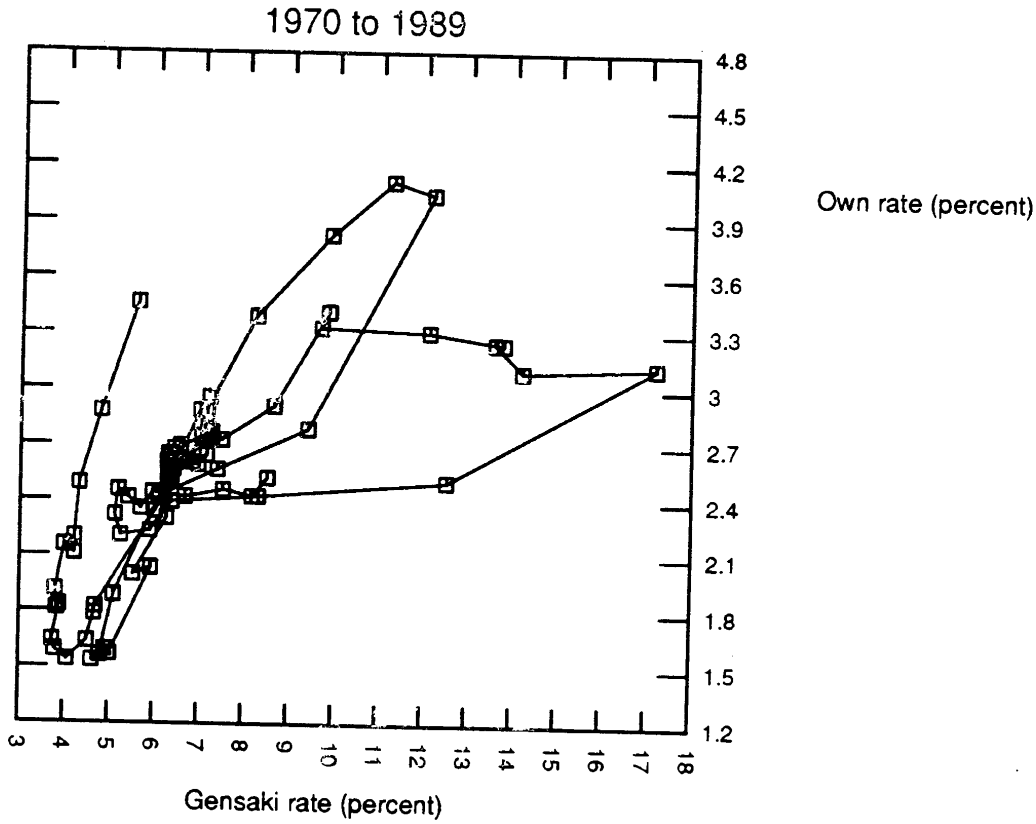


Chart 4

RELATION OF RATES ON JAPANESE M2+CDs AND GENSAKI



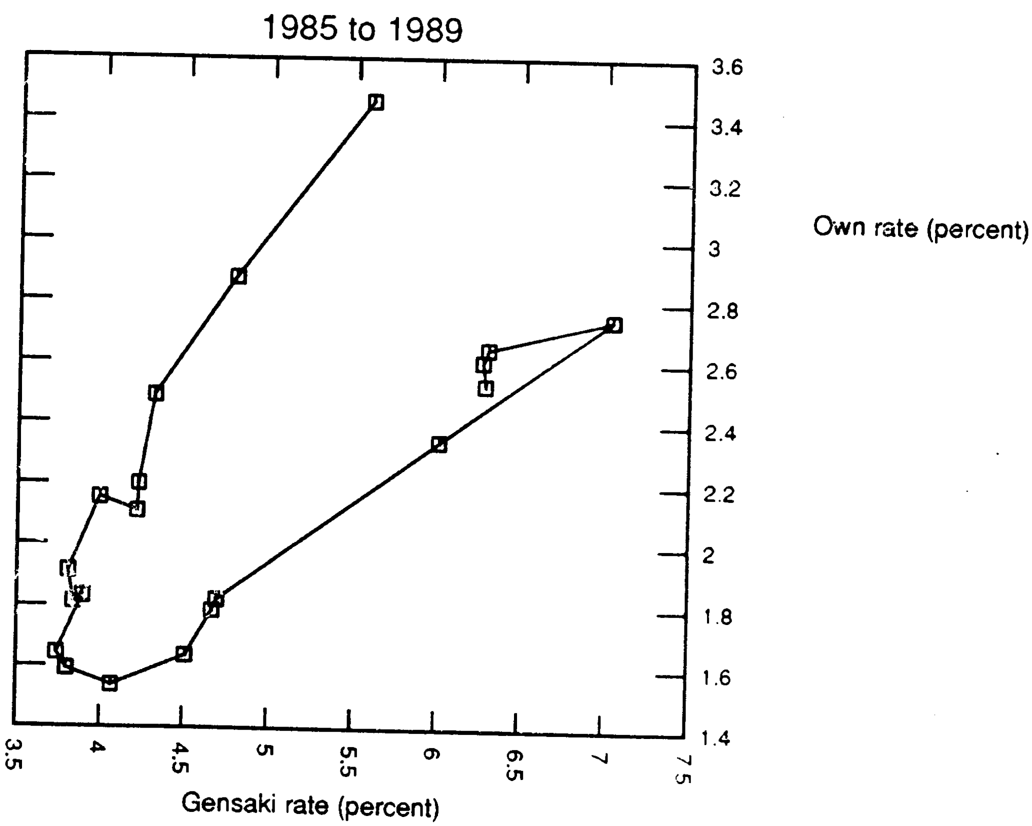


Chart 5

CANDIDATE SPECIFICATIONS FOR P* AND V*
(IN LOGS)

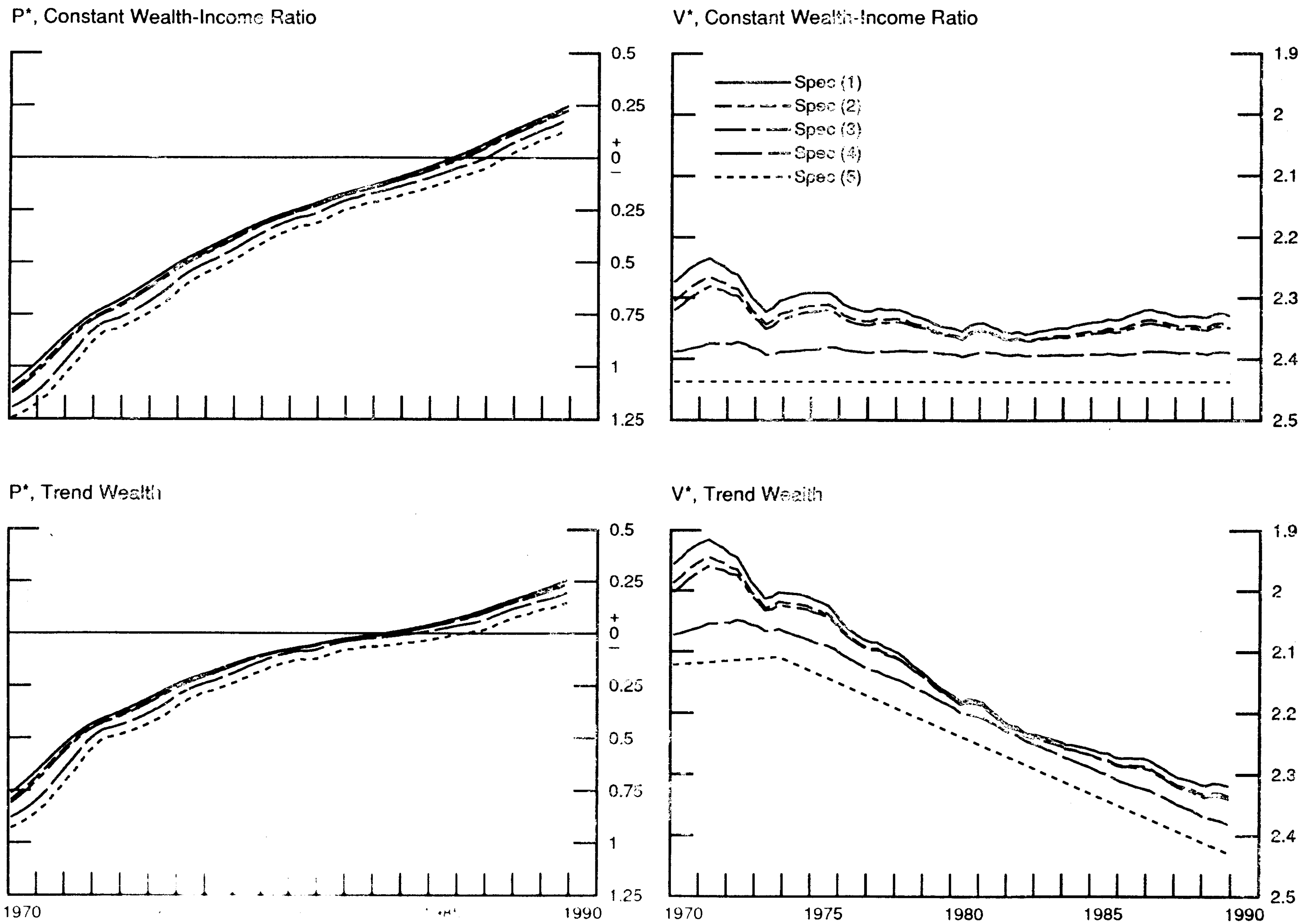


Chart 6

COMPARISON OF JAPANESE P* AND COMPONENTS

(Constant Wealth-Income Ratio)

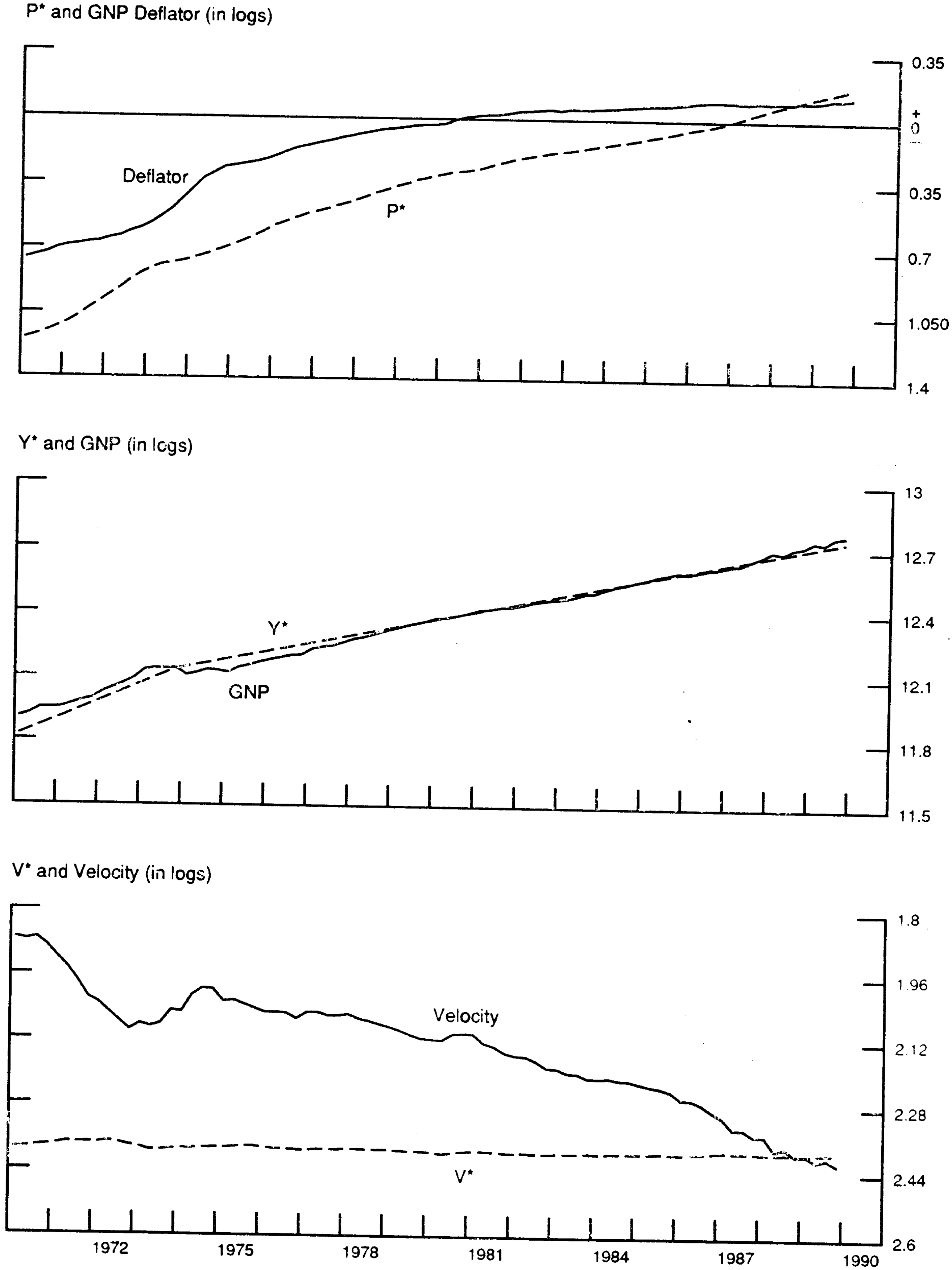


Chart 7

COMPARISON OF JAPANESE P* AND COMPONENTS

(Trend-Wealth)

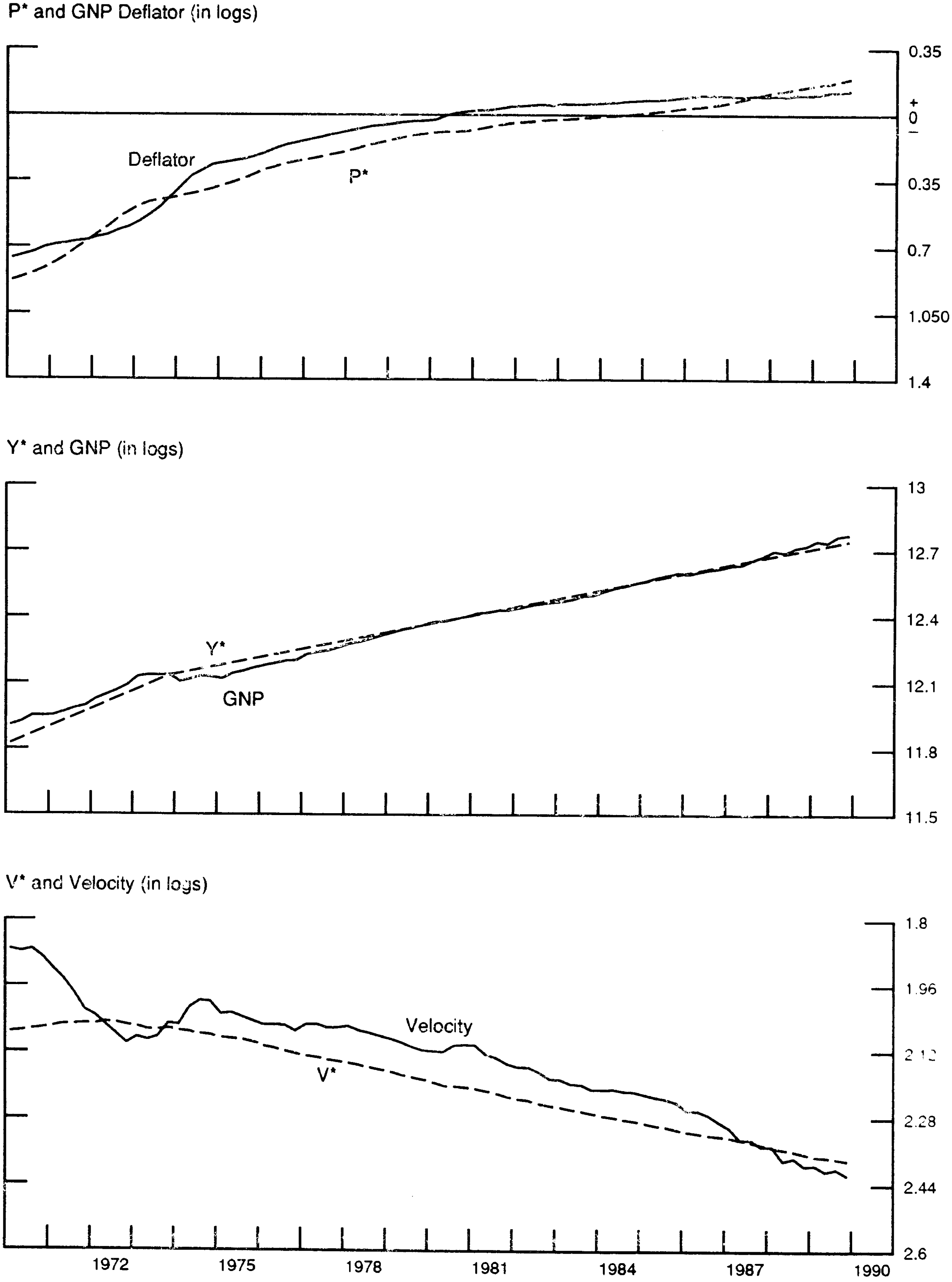
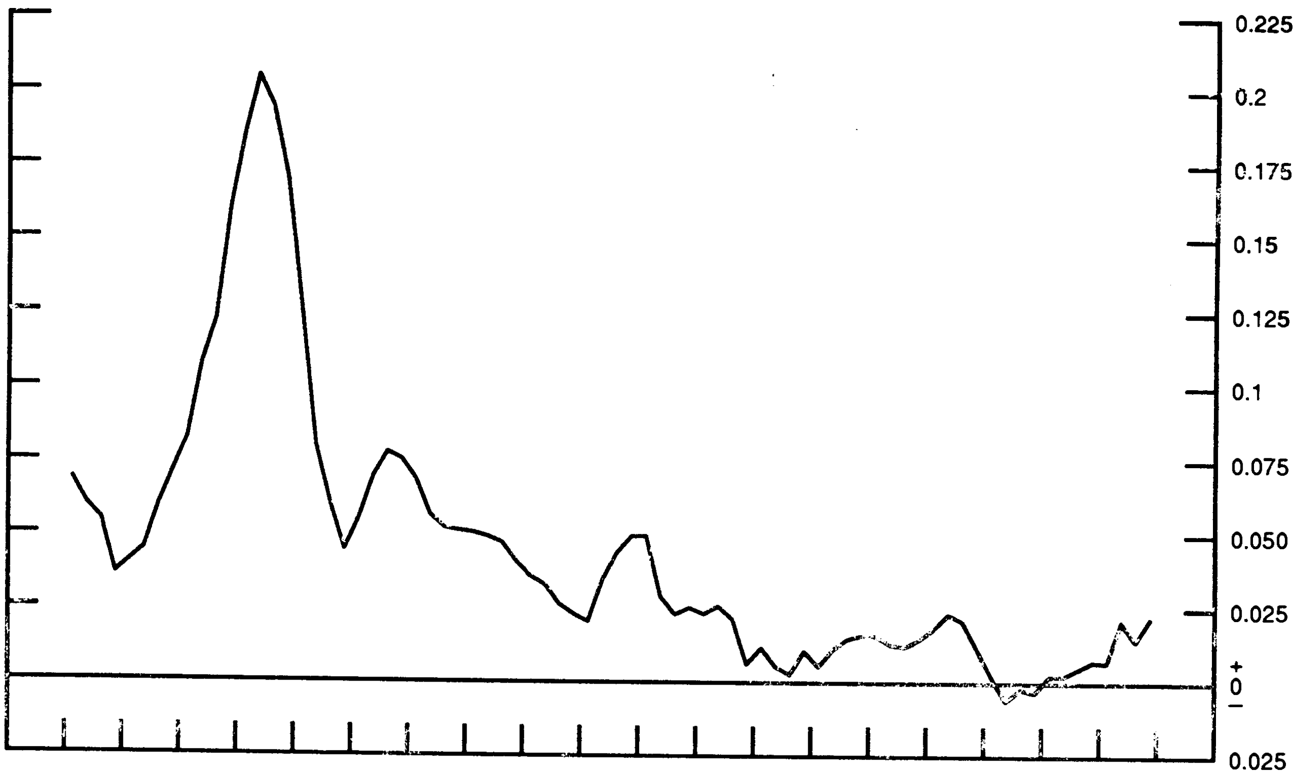


Chart 8

JAPANESE INFLATION AND PRICE GAP

Constant Wealth-Income Ratio

4 Quarter Change in GNP Deflator (in logs)



Price-Gap (p-p*), Constant Wealth-Income Ratio

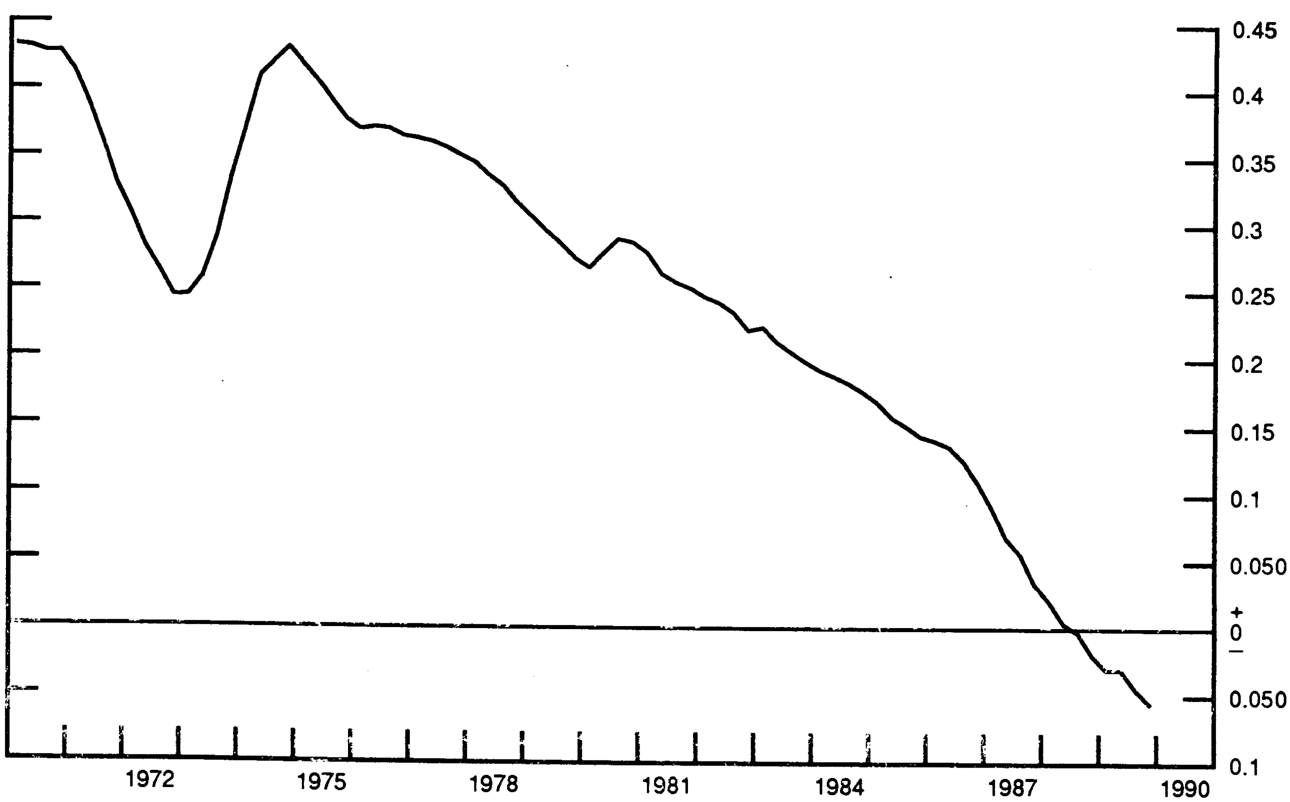
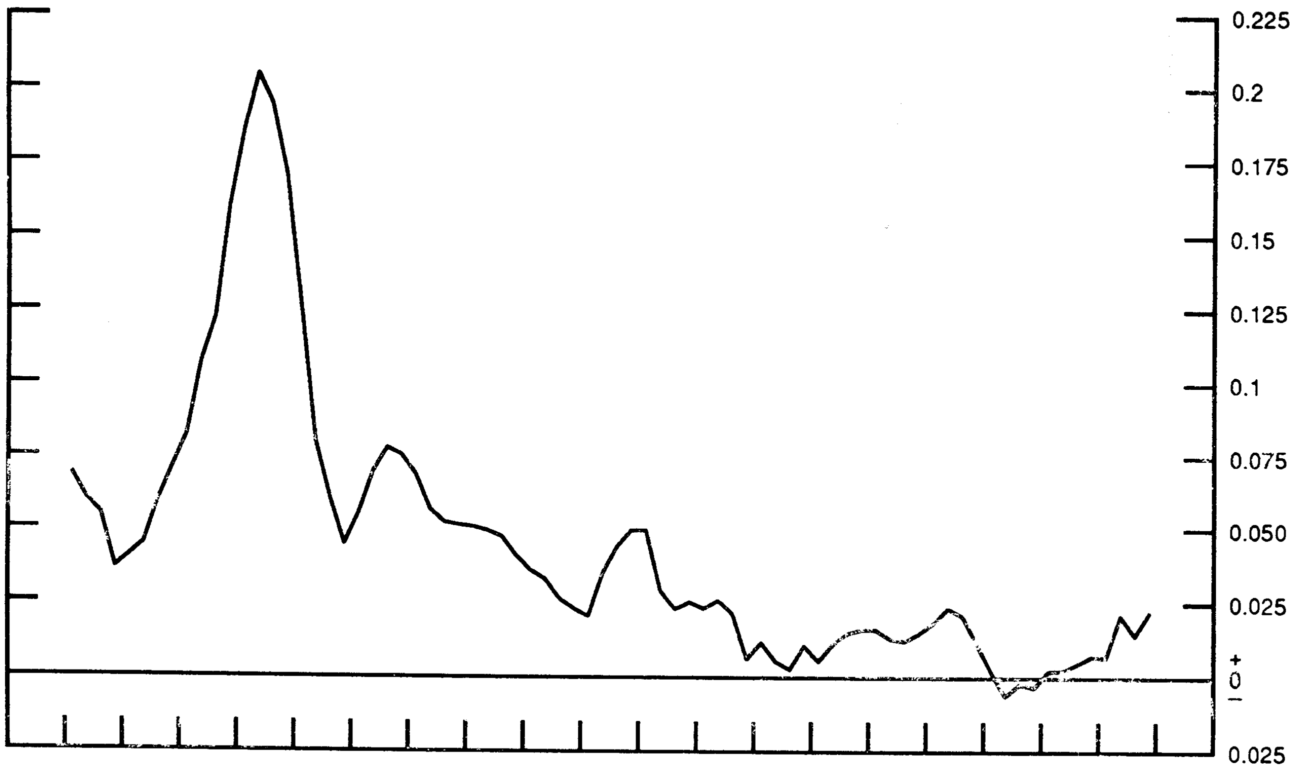


Chart 9

JAPANESE INFLATION AND PRICE GAP

Trend Wealth

4 Quarter Change in GNP Deflator (in logs)



Price-Gap (p-p*), Trend Wealth

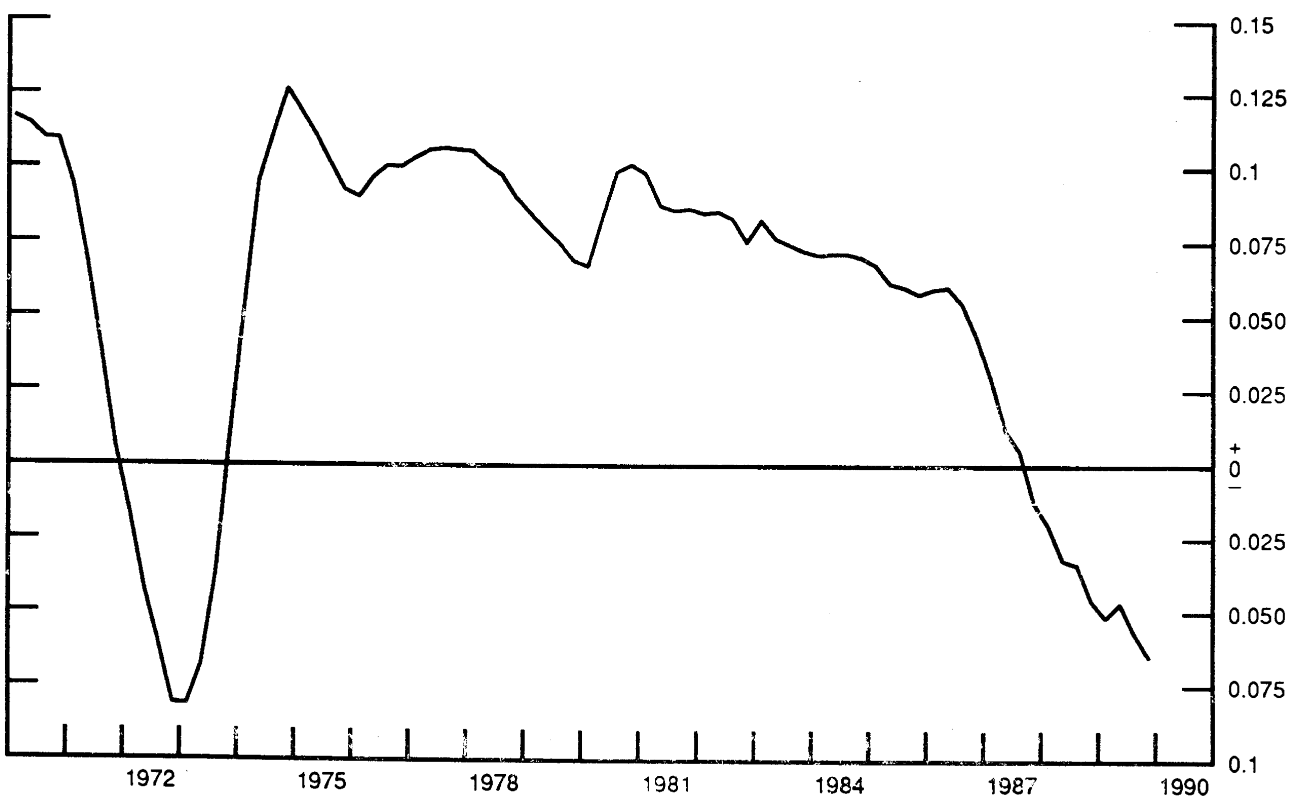


Chart 10

ONE-STEP-AHEAD CHOW TESTS FOR INFLATION INDICATORS

(Scaled by Critical Values of F-Distribution)

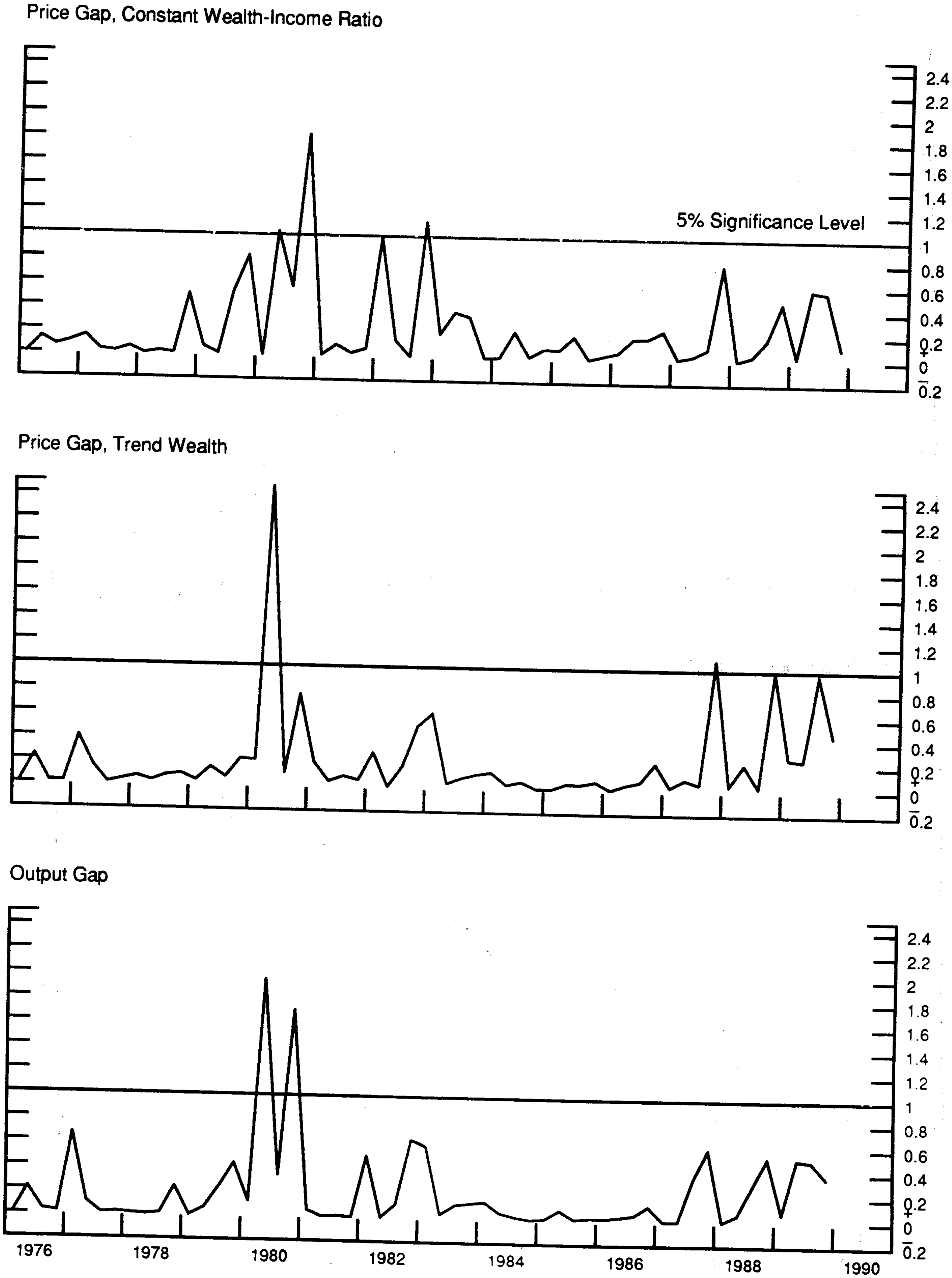


Chart 11

GERMAN M3 VELOCITY AND WEALTH/GNP

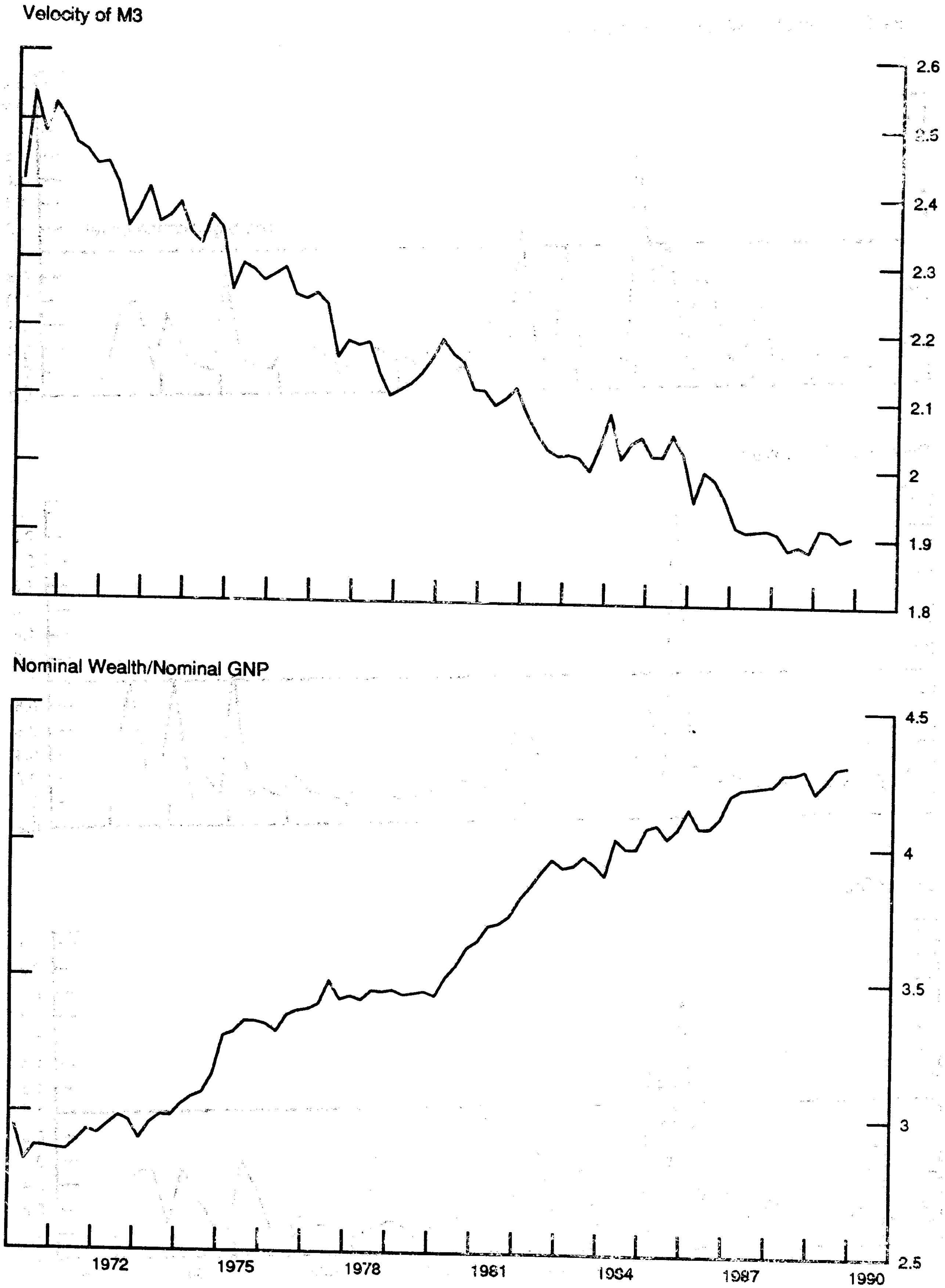
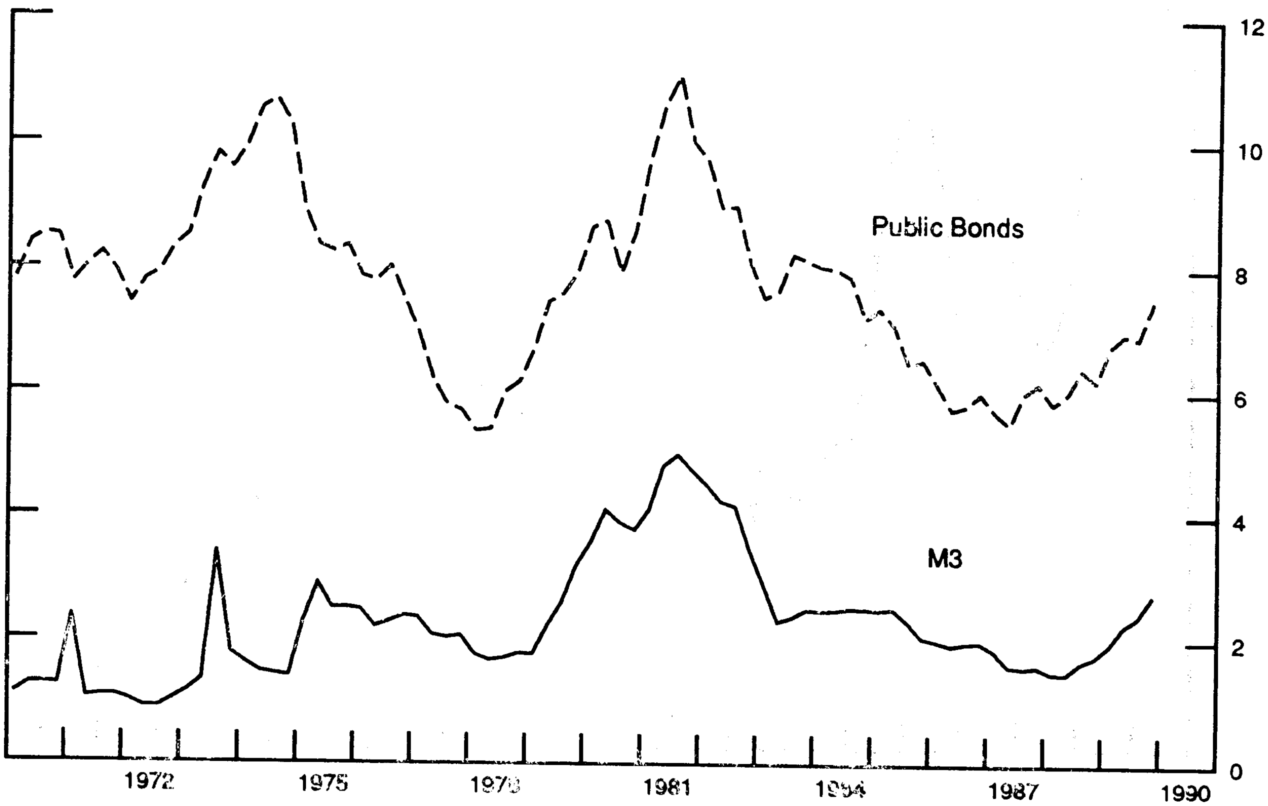


Chart 12

GERMAN INTEREST RATES

Own Interest Rate on M3 and Yield on Public Sector Bonds



One Step Ahead Chow Tests for German Interest Rate Relationship
(Scaled by Critical Values of F-Distribution)

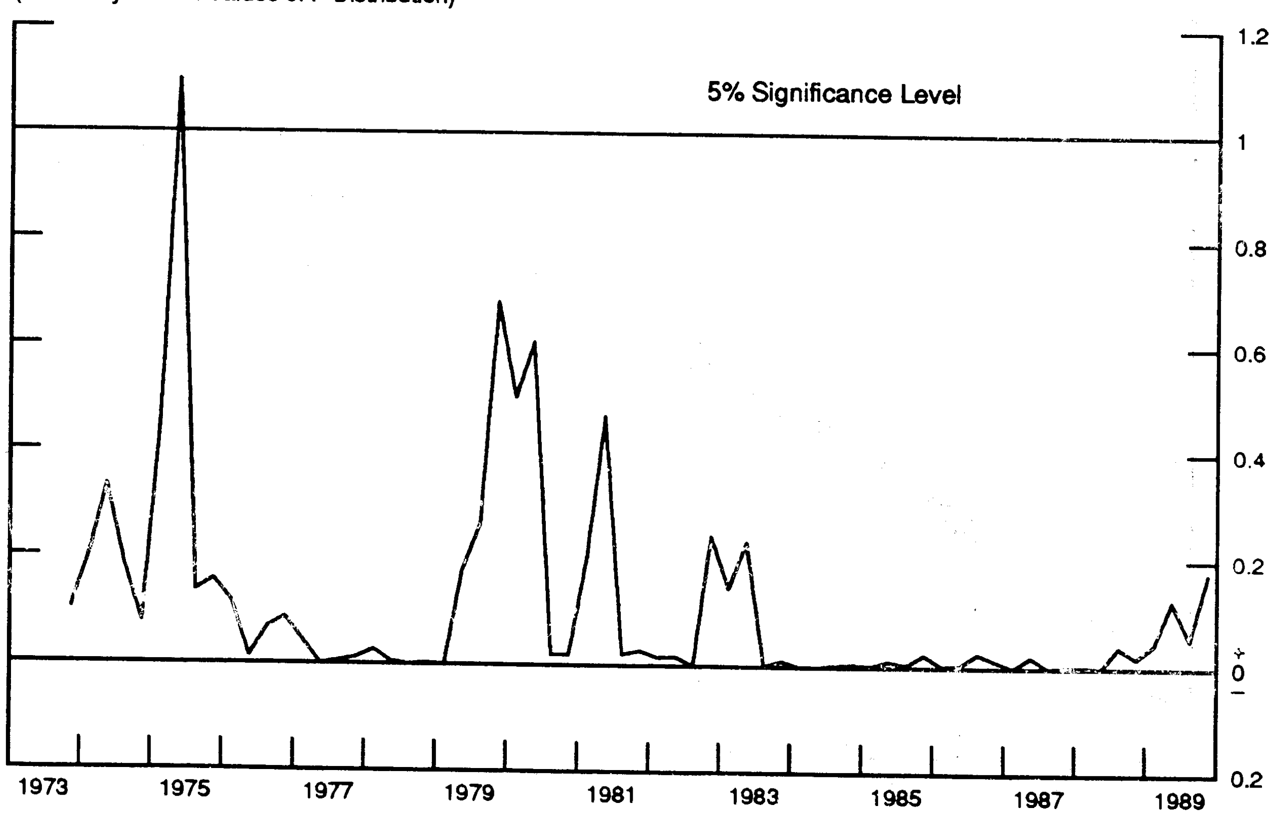
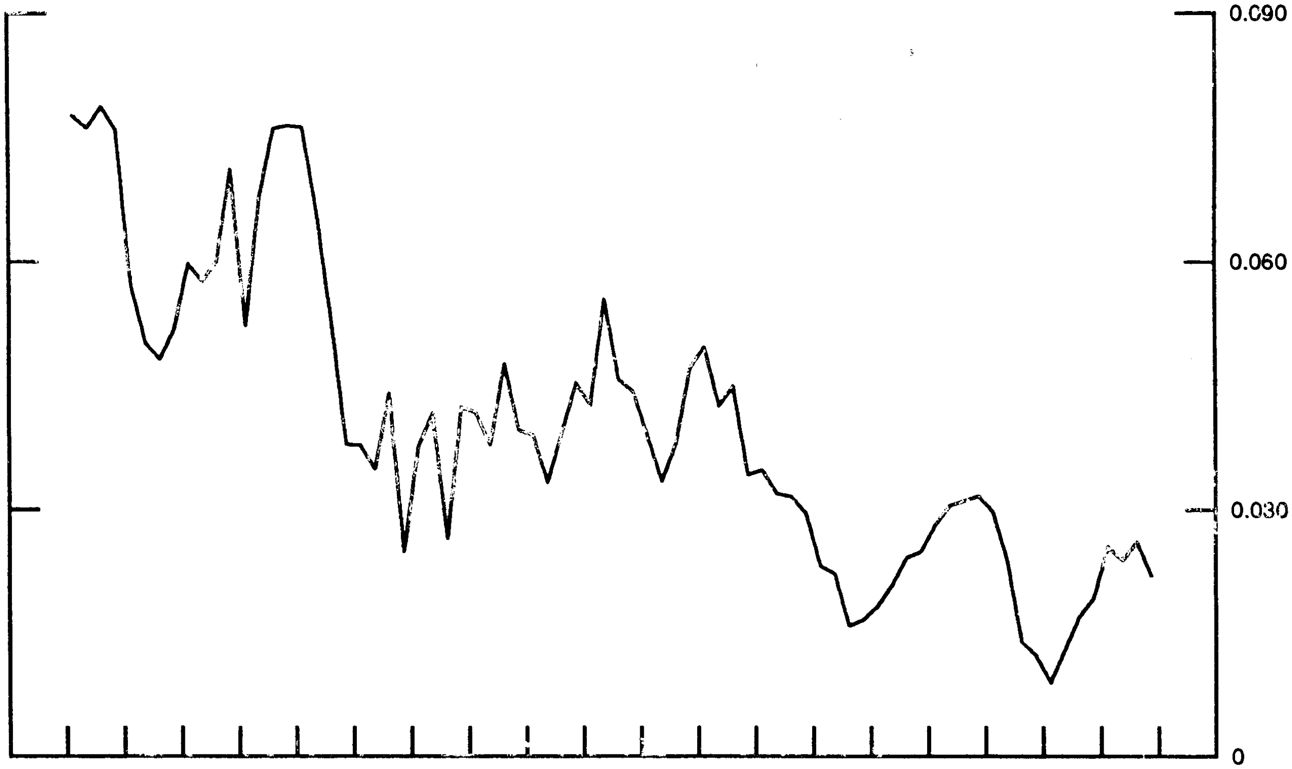


Chart 13

GERMAN INFLATION AND THE PRICE GAP

4 Quarter Change in GNP Deflator (in logs)



P* and GNP Deflator (in logs)

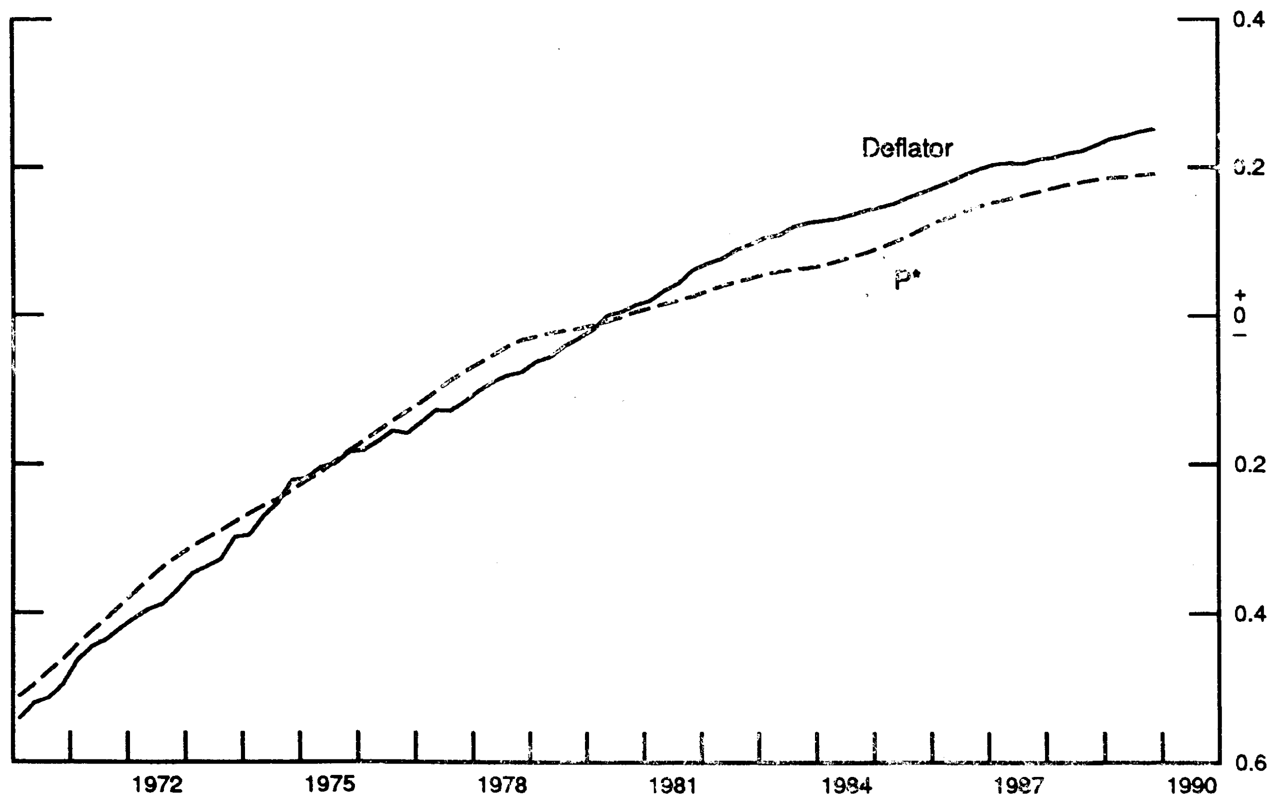


Chart 14

COMPONENTS OF THE GERMAN PRICE GAP

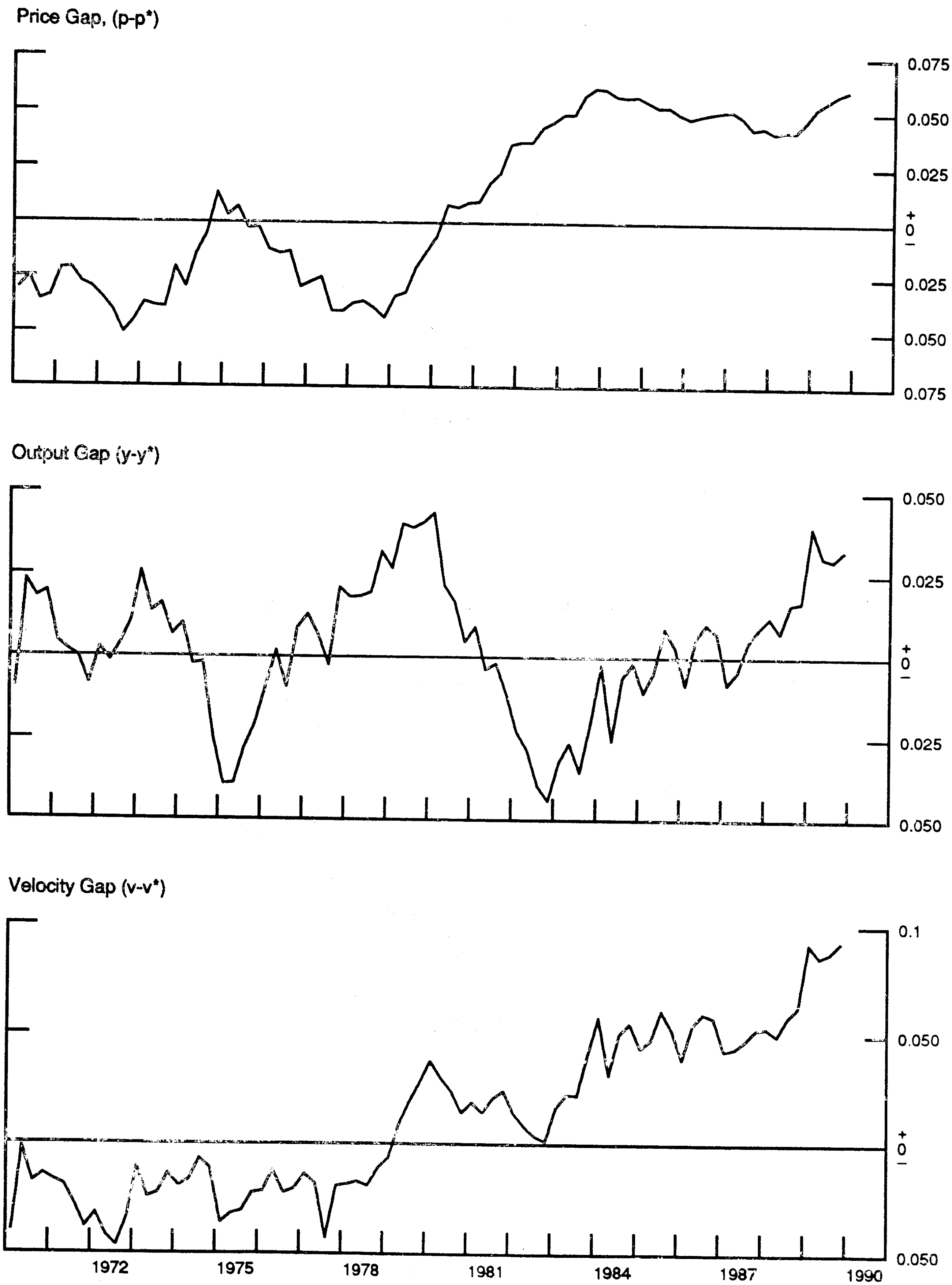
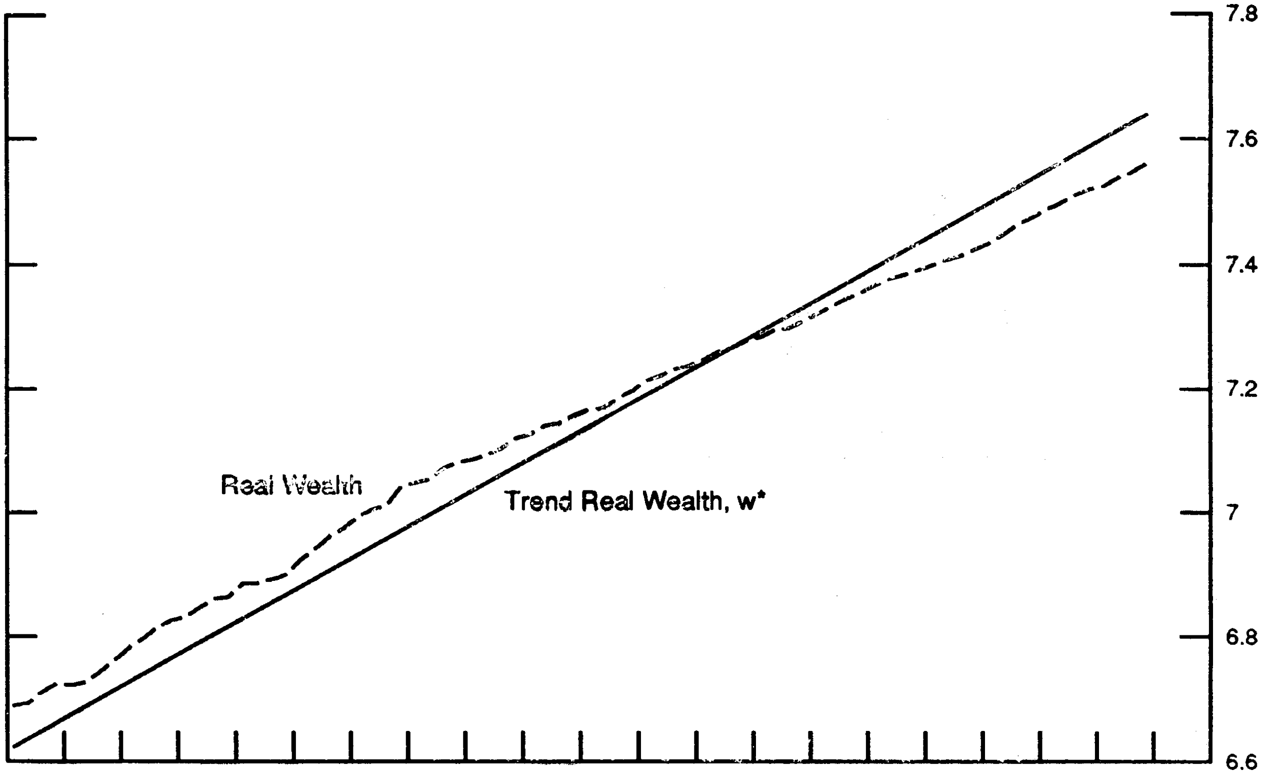


Chart 15

GERMAN REAL WEALTH (IN LOGS)

Case A - Trend Wealth Estimated from 1960Q1 to 1989Q4



Case B - Trend Wealth Estimated from 1970Q1 to 1989Q4

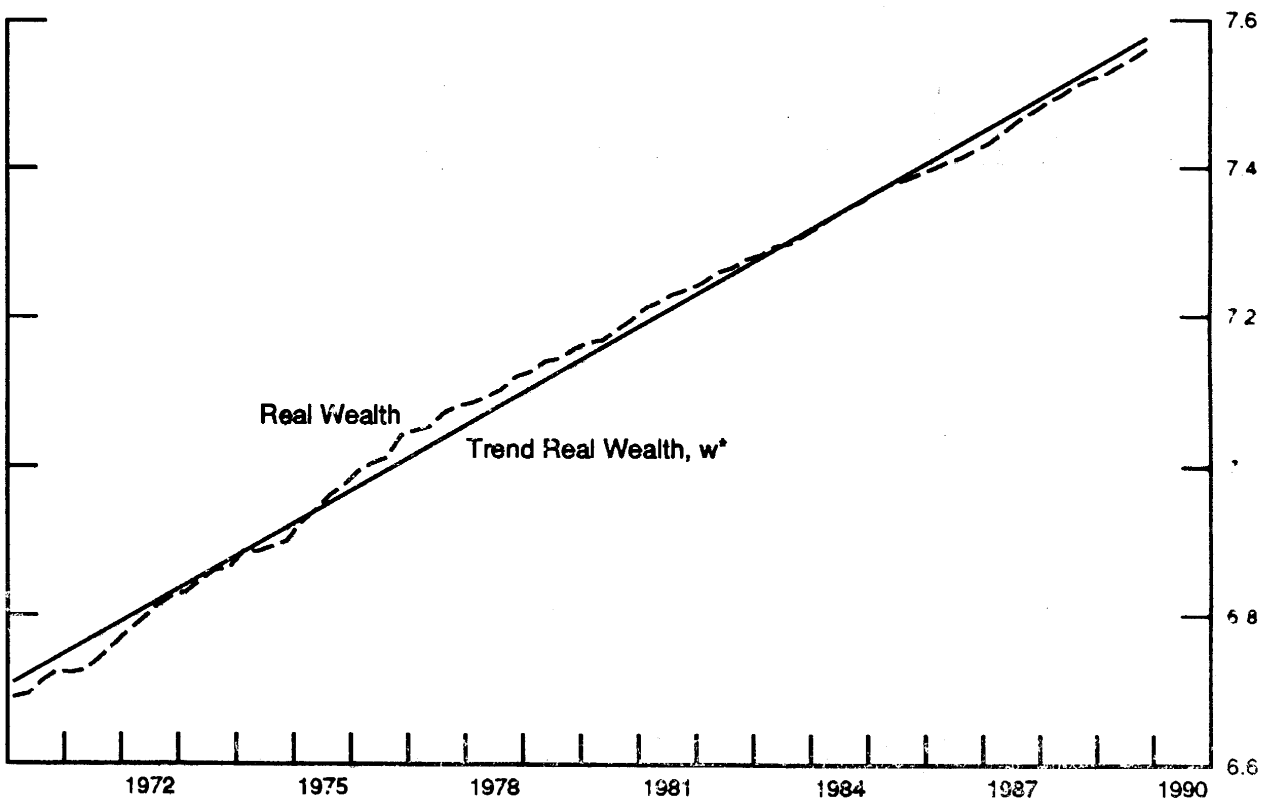
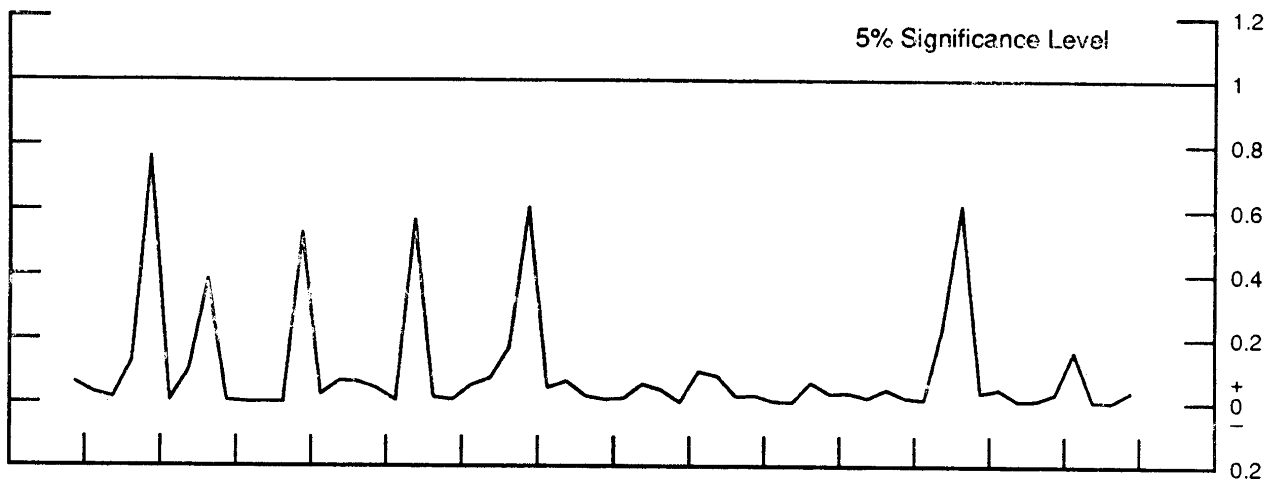


Chart 16

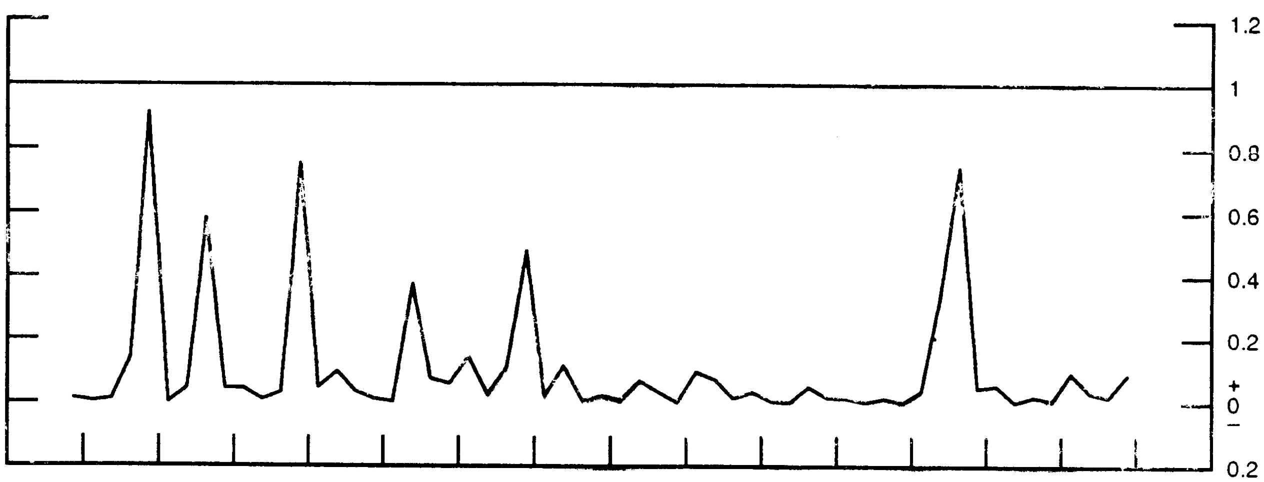
ONE-STEP-AHEAD CHOW TESTS FOR INFLATION INDICATORS

(Scaled by Critical Values of F-Distribution)

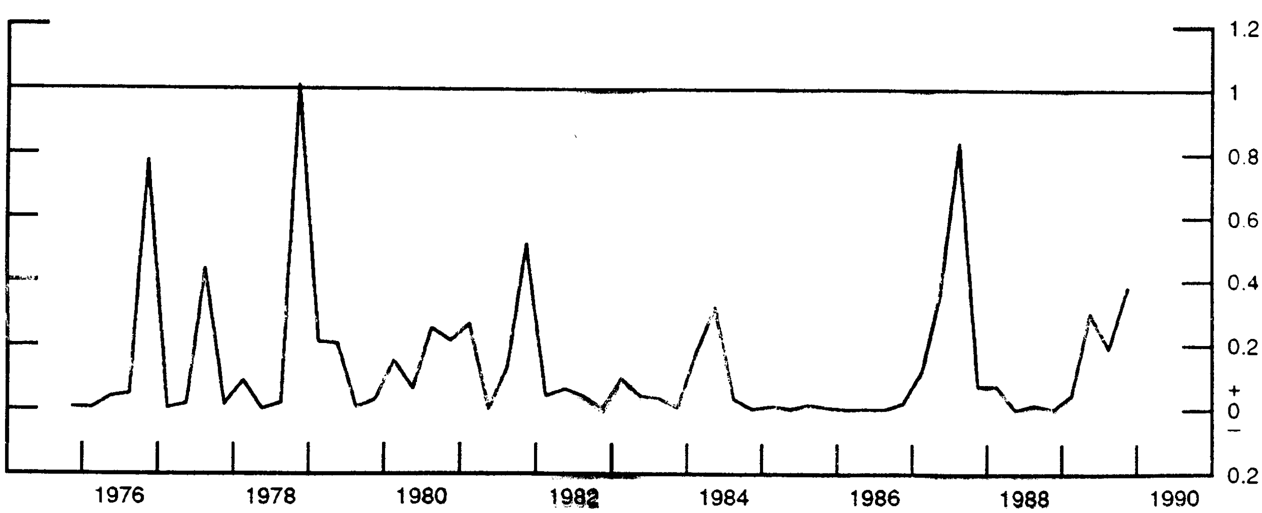
Price Gap, Trend Wealth - Case A



Price Gap, Trend Wealth - Case B



Output Gap



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