The Sources of the Movements in Interest Rates:

An Empirical Investigation

by

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Abstract

The paper investigates the sources of interest rate movements in Italy during the 1970s. In that decade, the country experienced the highest and most variable inflation rate among the industrial nations and was affected both by external shocks and by shocks to the labor and financial markets. The paper shows that movements in inflationary expectations dominated the nominal interest rate variability. The part of variability that was not explained by expected inflation was accounted for by unanticipated changes in the terms of trade; by contrast, domestic monetary and real disturbances did not have a sizeable impact on the variability.

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I. Introduction

A great number of empirical studies have investigated the determinants of the movements in nominal interest rates with a particular emphasis on the role of inflationary expectations.[1] The bulk of the empirical evidence, however, pertains to the U.S. economy and, as far as the post-World War II years are concerned, to sample periods that typically include over 25 years of economic stability and relatively predictable inflation rates (1948-1973). In this paper, I use the time series of a small open industrial country--which is therefore much more vulnerable to external shocks than the U.S. economy--during a period characterized both by extreme uncertainty about the future course of inflation and by recurrent disturbances, in order to re-examine the relation between nominal interest rates and inflation and to assess the relative importance of internal and external shocks for the variability of the domestic real rate of interest.

I find that the Italian economy during the 1970s is an ideal candidate. The inflationary environment in the industrial world, and consequently in Italy, changed markedly at the beginning of that decade. The collapse of the Bretton Woods system, the acceleration in the growth rate of the world money supply that accompanied it and the sharp increase in commodity prices of 1973 created the conditions for the first inflationary wave of the post-war years (Figure 1).[2] As it is well known, high and variable inflation rates persisted for the rest of the 1970s. In the case of Italy, the transition to an inflationary economy was particularly traumatic as the country began to experience the highest and most variable inflation rate among the industrial countries after enjoying the lowest and least variable rate in the 1960s

(Table 1). Survey data indicate that this change in inflationary environment was not fully anticipated by economic agents (Visco, 1984) and, therefore, it is reasonable to infer that the Italian economy was characterized by a higher level of uncertainty in the 1970s than in the previous two decades.

There are other reasons for focussing on the 1970s. First, Italy experienced a "wage shock" between 1969 and 1973, as trade unions successfully increased the labor share in value added by nearly 8 percentage points (Sachs, 1979). Second, since the beginning of the floating exchange rate period, currency fluctuations have added to the variability of the real sector by causing sharp movements in the terms of trade and by triggering wage claims; these "external shocks" were compounded by the two energy crisis of 1973 and 1979 (Figure 2). Third, the steady budgetary policies of the 1950s and 1960s were discontinued around 1972 and the government began to run growing budget deficits (Figure 3). And fourth, monetary aggregates grew at a very uneven pace during the decade, partly because of the need to finance sudden increases in public sector borrowing requirements (Figure 4).

Notwithstanding the uncertainty about the inflation rate and the disturbances that affected the Italian economy in the 1970s, I show that financial markets anticipated the monthly movements in the inflation rate rather accurately so that inflationary expectations accounted for a very large proportion of the short-run changes in nominal interest rates. Part of the interest rate variability, however, was due to movements in the ex-ante real rate. I find that the external shocks, measured by unanticipated changes in the terms of trade, accounted for most of the real rate variability while domestic shocks, such as unanticipated changes in the growth of monetary aggregates, of the government non-monetary debt, and of the level of output, did not have a sizable impact on the variability.

2. Financial Market Versus Time Series Predictions

In each period of time, the nominal rate of interest on a financial asset can be decomposed into the inflation rate expected to prevail during the life of the asset and its anticipated (ex-ante) real rate of return. Let i_t be the nominal interest rate set by the financial market on the last day of t - 1 for an asset that matures at the end of t; π_t^e the prediction of the percentage change in the price level during period t based on information available at the beginning of that period; and r_t the real rate of return on the asset anticipated by the market for period t at the end of t - 1, then:

$$(1 + i_t) = (1 + r_t)(1 + \pi_t^e) \qquad (1)$$

By bringing the nominal interest rate to the right hand side of the equation, the relation between interest rates and inflation can be rewritten as:

$$\pi_t^e = \frac{i_t - r_t}{(1 + r_t)} \qquad (2)$$

which shows that the interest rates set by the market at the beginning of each period contain predictions about the subsequent level of inflation. These predictions will be very accurate if market participants systematically use all available information to price financial assets. As a result, a test of financial market efficiency is a forcasting test, as Fama (1975) pointed out. Specifically, if interest rates are set in an efficient market, the predictions of the inflation rate generated by using equation (2) must be systematically more accurate than those obtained from other available predictors. In what follows, I test this proposition.

In order to make this test operational, which involves the use of equation 2, I need to have--in addition to the nominal interest rate--a time series which measures the ex-ante real rate. Because such a time series cannot be observed, I assume that market participants anticipate the future movements of the real rate of interest on financial assets by looking at the history of the realized (ex-post) real rate of return, that is, the nominal interest rate minus the realized inflation rate. I then construct estimates of the ex-ante real rate, $\hat{r}_t$, by generating one-step-ahead predictions from an ARIMA model for the ex-post real rate of return (R_t),[3] that is:

$$\hat{r}_t = E(R_t | R_{t-1}, R_{t-2}, R_{t-3}, \ldots) = \beta_0 + \beta_1 R_{t-1} + \beta_2 R_{t-2} + \cdots \tag{3}$$

By substituting $\hat{r}_t$ for r_t in equation (2) and using the nominal interest rates observed at the beginning of each period t, in practice the interest rates prevailing on the last day of t - 1, I obtain the "financial market" predictions for the inflation rate (PM_t), that is:

$$PM_t = (i_t - \hat{r}_t)/(1 + \hat{r}_t) \tag{4}$$

It is worthwhile noting that the approach followed here does not impose any constraint on the macroeconomic model underlying the movements in the ex-ante real rate of interest. It is only assumed the real rate can be represented by a stationary process that is known to financial market participants and that this process is used to generate predictions about the future movements in the ex-ante real rate.[4] Finally, more elaborate estimates of the ex-ante real rate could have been used to calculate a time series for PM_t--for example, a multivariate time series estimate of it, as it is done in Section 3. However, because the PM_t's turn out to be already the best available predictions of future inflation rates, more elaborate, and presumably more accurate, estimates of the real rate would have only reinforced the empirical results of this section.

Before presenting the alternative predictors that I will compare to PM, I briefly discuss the time series used, which are shown in Figure 5. The inflation rate is the month-to-month rate of change of the cost-of-living index calculated by the Central Institute for Statistics, that is typically published, at least for the major cities, within two weeks of the end of the month which it refers to. I opt for monthly observations because this is the frequency with which "news" on the inflation rate reach the market. The time series of the nominal interest rate is the average bid-offer 30-day interbank rate quoted by the daily financial newspaper Il Sole-24 Ore on the last day of every month. This interest rate has the advantages of being free from the intervention of the monetary authorities, of being highly publicized, and of being available throughout the 1970s. The rate, which is quoted on an annual basis and in per cent, is converted to a monthly basis according to:

$$i_t = \left[1 + (i_t^A/100)\right]^{1/12} - 1$$

where i_t is the nominal interest rate on a monthly basis and i_t^A is the nominal interest rate on a annual basis and in per cent. Finally, the choice of the initial month of the sample period, which is May 1972, is due to data availability.

The two alternative predictors that I choose are a univariate ARIMA model for inflation (PA_t) and a distributed lag of past movements in the wholesale price index (PW_t). The first predictor is very popular, is easily computed and has often outperformed structural macroeconometric models in forecasting accuracy. The second predictor is based on the widely-held belief that movements in wholesale prices tend to anticipate movements in consumer prices. It also reproduces the consumer price equation of most structural econometric models in which prices are determined by a distributed lag of domestic production costs and imported inputs. The predictions PW_t were obtained by regressing the rate of change in the cost-of-living index on the rates of change of the wholesale price index of the previous ten months. Summing up, PA_t is a predictor of the inflation rate based on past inflation; PW_t is a predictor based on past movements in the rates of change of the wholesale price index; and PM_t is a predictor based on the most recent nominal interest rate and the history of the ex-post real rate.

Because nominal interest rates at the end of every month cannot contain information beyond that date, I assess the relative accuracy of equation (4) in predicting the future inflation rate by using out-of-sample predictions. Accordingly, I initially estimate an ARIMA model for the inflation rate, an ARIMA model for the ex-post real rate, and a distributed lag model for the wholesale prices by using monthly data from the sample period May 1972 to August 1978. Then, I use the estimated models, together with the end-of-month nominal interest rates, to generate three series of 20 out-of-sample predictions--from September 1978 to April 1980. Table 2 shows the estimated models with data up to August 1978 and with data from the entire sample period; because the estimated parameters are very stable, the results are virtually identical whether the models are re-estimated after each prediction or not.

The lower part of Table 3 reports statistics of the predictive accuracy of PM, PA, and PW. Clearly, the predictions PM, which are based on equation (4), have both lower root mean squared error and mean absolute error than those obtained from the two other predictors of inflation. An additional comparison of the alternative predictors can be obtained by estimating the following regression:

$$\pi_t = (1 - k)Y_t + kPM_t + \varepsilon_t \qquad Y_t = PA_t,\ PW_t \tag{5a}$$

where π_t is the actual inflation rate and Y_t is, alternatively, PA_t or PW_t. Because the competing predictors are highly correlated, it is better to rewrite equation 5a in a different way in order to avoid the problem of multicollinearity in the estimation:

$$(\pi_t - Y_t) = k(PM_t - Y_t) + \varepsilon_t \qquad Y_t = PA_t,\ PW_t \tag{5b}$$

The parameter k is the coefficient of the regression of the prediction errors of PA or PW on the difference between PM and PA or PW; the greater the capability of PM is to account for the prediction errors of PA or PW, the greater its weight in an optimal composite predictor. If the difference between PM and PA or PW explains the prediction errors of PA or PW perfectly, that is k is equal to one, then the predictions based on the nominal interest rates incorporate all information about the future inflation rate that are contained in PA or PW.

There is another interpretation to equation 5. The predictions Y_t are constructed by linearly projecting actual inflation on past inflation rates (measured alternatively by the consumer or the wholesale price index), that is:

$$Y_t = E(\pi_t | \pi_{t-1}, \pi_{t-2}, \ldots) = \alpha_0 + \alpha_1\pi_{t-1} + \alpha_2\pi_{t-2} + \cdot\cdot\cdot \tag{6}$$

while the predictions PM_t are obtained using (3) and (4). If the small term $(1/1 + \hat{r}_t)$ in (4) is neglected, using (3), (4) and (6), equation (5) can be re-written as:

$$\pi_t = \gamma_0 + \gamma_i i_t + \gamma_2 R_{t-1} + \gamma_3\pi_{t-1} + \gamma_4 R_{t-2} + \gamma_5\pi_{t-2} + \cdot\cdot\cdot\cdot + u_t \tag{7}$$

where $\qquad \gamma_0 = (1 - k)\alpha_0 - k\beta_0 \qquad \gamma_3 = (1 - k)\alpha_1$

$$\gamma_1 = k \qquad \gamma_4 = -k\beta_2$$

$$\gamma_2 = -k\beta_1 \qquad \gamma_5 = (1 - k)\alpha_2$$

A test that k is equal to one in (5a) or (5b) is therefore equivalent to testing that the coefficients of past inflation rates are all equal to zero and that γ_1 is equal to one in a distributed lag equation like (7). In other words, notwithstanding the high serial correlation of inflation rates, past inflations do not contain information about π_t once past ex-post real rates and the nominal interest rate prevailing at the end of t - 1 are taken into account. The specification of equation 5b, however, is preferable to that of equation 7, because of the considerably smaller number of parameters to be estimated.

Table 3 shows that the estimated value of k is not significantly different from one when PW is used to estimate (5b). This indicates that past movements in the rates of change of the wholesale price index do not provide market participants with useful information about the future rate of inflation once they know the levels of nominal interest rates and the history of the ex-post real rate. When PA is used in the regression, the table shows that the point estimate of k exceeds the value of two, thus indicating that the ARIMA model produces "bad" predictions of the inflation rate--relative to those based on interest rates. However, because the estimated k is significantly greater than one, it would seem that the relative large prediction errors of the ARIMA model could nonetheless be utilized to improve the accuracy of the inflation predictions. As Granger and Newbold (1977, p. 270) pointed out, a composite forecast of a "good" predictor and a "bad" predictor may outperform the forecast obtained from the "good" predictor alone if the errors of the two individual predictors are sufficiently correlated.[5]

To verify this possibility, which would contradict the market efficient hypothesis, I construct a composite predictor of future inflation rates (CF) by combining PM and PA with time-varying weights along the way suggested in Granger and Newbold (1977), that is:

$$CF_t = \lambda_t PM_t + (1 - \lambda_t) PA_t$$

$$\lambda_t = \frac{(e_{t-1}^{PA})^2}{(e_{t-1}^{PM})^2 + (e_{t-1}^{PA})^2}$$

where e_t^{PA} and e_t^{PM} are the prediction errors of PA and PM, respectively. The estimated λ_t's range from .13 to .92 with an average value of .51.[6]

In Table 3, I show that CF does not outperform PM, the efficient market forecast, even though it is clearly more accurate than both PA and PW. The mean absolute error of CF is 1.7 percent larger than that of PM and the root mean square error is 5.3 percent larger. There is strong evidence, therefore, that, even for periods as short as a month, nominal interest rates are very accurate predictors of subsequent inflation rates so that changes in inflationary expectations are a main determinant of interest rate movements.

Figure 5, in which the nominal interest rate and the monthly inflation rate are plotted together, may raise some doubts about the robustness of the preceding econometric evidence because the figure shows that the smooth behavior of the nominal interest rate contrasts sharply with the wide and frequent fluctuations of the monthly inflation rate. Shiller (1980) and Fama (1977) have suggested the possibility that the correlation between inflation and nominal interest rates is mostly "spurious," that is, it is not the result of the economic relationship existing between the two variables, but is a statistical artifact. The correlation would then stem from the presence of unit roots in the process followed by inflation and nominal interest rates.[7]

In order to test the "spurious" correlation hypothesis and to gain further insight into the relationship between inflation and interest rates, I first take the first differences of both variables and then compute their squared coherency. By computing the squared coherency, I measure the part of the variability of one time series that is "accounted for" by the other series at each frequency or cycle. If the correlation between inflation and interest rates is mainly due to unit roots in their processes, the squared coherency of the first differences of the two series should not be significantly different from zero at all frequencies.

In the lower part of Figure 6 and in Table 4, I show instead that the estimated squared coherency is very high at low frequencies and becomes not significantly different from zero at the .10 significance level only when the cycle period is shorter than nine months.[8] Thus, long- and medium-run cycles of inflation are highly correlated with the long- and medium-run cycles of the nominal interest but this correlation declines as the cycle period of the inflation rate become shorter. These findings provide evidence that the correlation between interest rates and inflation is not spurious. They also indicate that the short-run relationship between interest rates and inflation is not stable in the short run, perhaps due to random structural shifts in the process for the real rate.[9] Notwithstanding these shifts, however, equation (4) remains the best available predictor of future inflation even in the short run, as the previous evidence clearly indicates.

3. The Variability of the Ex-Ante Real Rate

In addition to the expected inflation rate, changes in the nominal interest rate reflect movements in the ex-ante real rate. In this section, I use the method of vector autoregressions (VAR) proposed by Sims (1980) to assess the relative importance of domestic and external shocks as sources of the real rate variability. Although structural inference cannot be solely based on VAR models, the models provide useful information about the dynamic relationship among economic variables that structural models must ultimately account for.

Calling $\underline{x}_t$ the vector of all the variables entering the VAR model, I estimate

$$\underline{x}_t = \sum_{i=1}^{K} A_i \underline{x}_{t-i} + \underline{v}_t \qquad E(\underline{v}_t \underline{v}_t') = V \tag{8}$$

where the dash denotes a vector; A_i is a suitable matrix; K is the number of lags which I arbitrarily choose to be six; and $\underline{v}_t$ is the vector of disturbance terms which are cross-correlated but serially uncorrelated. Assuming that the A_i matrices satisfy certain conditions, it is possible to obtain the moving average representation of (8)

$$\underline{x}_t = \sum_{i=0}^{\infty} D_i \underline{v}_{t-i} \tag{9}$$

Because the $\underline{v}_t$s are cross correlated, it is useful to orthogonalize them by using a suitable linear transformation,

$$\underline{v}_t = L\underline{e}_t \qquad E(\underline{e}_t \underline{e}_t') = I \tag{10}$$

where I is the identity matrix. The model is now ready to be transformed so that each variable is expressed as a linear combination of all the innovations of the system; the innovations are uncorrelated and have unit variances. In terms of the previous notation, (9) can be expressed as:

$$\underline{x}_t = \sum_{i=0}^{\infty} H_i \underline{e}_{t-i} \tag{11}$$

where $H_i = D_iL$. In order to make the method operational, I have to neglect the lags beyond a certain order in (11). I call this maximum order g; in practice, I set g equal to 48.

If n is the number of variables in the system and h^1_{ij} is the i^{th} row and j^{th} column element of the matrix H_1, then the k^{th} variable of the model can be expressed as

$$x_{kt} = \sum_{j=1}^{n} \sum_{i=1}^{g} h^i_{k,j} \, e_{j,t-i+1} \tag{12}$$

The ex-ante real rate does not appear in the $\underline{x}_t$ vector because it is not observed. However, it can be expressed as a function of the variables of the system, that is, the xs. This can be done in the following way. Assume that the nominal interest rate is the kth variable of the system and the inflation rate the sth variable. I first express the nominal interest rate and the inflation rate as in (12). Then, I delete the contemporaneous disturbance from the expression for the inflation rate in order to get the prediction of the inflation from the vector autoregression. Finally, I obtain the ex-ante real rate (RR_t) by subtracting this prediction of the inflation rate from the nominal interest rate; RR_t is then equal to:

$$RR_t = (x_{kt} - \hat{x}_{st}) = \sum_{j=1}^{n} h^1_{kj} \, e_{j,t} + \sum_{j=1}^{n} \sum_{i=2}^{g} (h^i_{kj} - h^i_{sj}) e_{j,t-i+1}$$

where ^ means predicted value. Because the e_ts are uncorrelated, the variance of RR_t, $V(RR_t)$, is a linear combination of the variances of the innovations of the model:

$$V(RR_t) = \sum_{j=1}^{n} \left[(h^1_{kj})^2 + \sum_{i=2}^{g} (h^i_{kj} - h^i_{sj})^2 \right] \sigma^2_{ej} \tag{13}$$

Expression (13) gives the contribution of a unitary increase in the variance of each of the j innovations of the system to the variance of the ex-ante real rate.

I now turn to the selection of the variables entering the model. In the introduction I argued that several shocks hit the Italian economy during the 1970s and, presumably, affected the real rate of interest. I take the monthly rate of change in the terms of trade as a measure of the external shock; the monthly rate of change in the index of industrial production as a measure of the real shock; the monthly rate of change of money-M2 and of the monetary base as measures of the monetary shock; and the monthly rate of change of non-monetary government debt as a measure of the fiscal shock. These four variables, together with the inflation rate and the nominal interest rate, are alternatively used in the VAR models.[10]

Because the matrix in L in (10), and consequently the final results, depend on the order in which the variables are taken to orthogonalize the vector of innovations, I consider several models. In particular, Cooley and LeRoy (1982) show that the economic interpretation of VAR models is correct only if the orthogonalization procedure in (10) restricts the variance-covariance matrix to be equal to the matrix of the true--but unobserved--structural model. Clearly, the economic interpretation of these models is less likely to be affected by misspecification if the results are insensitive to the order in which variables are orthogonalized.

The first two models analyzed (Models A and C) typify textbook models of a small open economy; in Model A, the terms of trade are given to the country; the stock of money is fixed by the monetary authorities; movements in interest rates transmit monetary impulses to the real sector; and the inflation rate is determined by an excess demand in the goods market. Model C typifies instead the Mundell-Fleming model in which there is a one-to-one correspondence between movements in the nominal exchange rate and the terms of trade. The variance decomposition of the ex-ante real rate is shown in Table 5. The two models produce the same results, indicating that the terms of trade is the most important determinant of the real rate variability during the sample period. The second most important factor is the uncertainty about the inflation rate. The large contribution of inflation innovations to the variance of the ex-ante real rate of interest indicates that the uncertainty about future changes in the price level has a substantial effect on the anticipated rate of return on real assets.[11] By contrast, the innovations of the industrial production of the money stock and of the nominal interest rate contribute very little to the variability of the real rate. Thus, it appears that external shocks were the main source of disturbances to the real sector in the 1970s. The results do not change if I substitute the monetary base for M2 as a measure of the monetary shock in the two previous models (Models B and D).

I also try three other specifications. First, I assume that M2 is not controlled by the monetary authorities, but that is instead endogeneously determined by the level of economic activity and the inflation rate; M2 is thus entered as the last variable in the model (Model E). Second, I assume that monetary authorities adjust the monetary base by using a feedback rule with respect to economic activity which fluctuates mainly due to real exogenous shocks (Model F). Thirdly, I add a fiscal shock to the VAR model by adding the changes in non-monetary government debt to the VAR as the first variable (Model G). Notwithstanding these numerous alternative specifications, I obtain the striking result that the terms of trade innovations dominate the real rate variability regardless of the domestic policy variables chosen and of the order in which the variables enter the VAR models.

4. Conclusions

In an economy in which the inflation rate fluctuates, movements in observed nominal interest rates reflect movements both in the ex-ante real rate and in the expected inflation rate. This paper shows that, in Italy, the 30-day interbank rate predicted the future inflation rates more accurately than two elaborate quantitative predictors during the 1970s. This result suggests that inflationary expectations were the major source of the monthly movements in nominal interest rates. This finding is somewhat surprising given the dramatic change in the inflationary environment that the country experienced at the beginning of the 1970s.

Movements in nominal interest rates did not simply mirror the movements in the expected rate of inflation because the unprecedented shocks that hit the Italian economy during the 1970s caused the ex-ante real rate, and thus the nominal interest rate, to vary a great deal. External shocks appear to have had the largest impact on the variability of the ex-ante real rate, which, by contrast, does not seem to be accounted for by domestic monetary and fiscal shocks.

Footnotes

*I thank Arnold Zellner, Rusdu Saracoglu and two anonymous referees for their very helpful comments. I remain responsible for remaining errors.

[1]Chapter 10 in Friedman and Schwartz (1982), for example, contains a critical review of the literature.

[2]The data used for Figures 1 through 4 are from IMF, International Financial Statistics.

[3]Nelson and Schwert (1977) and Hess and Bicksler (1975) make the same assumption about the ex-ante real rate of interest.

[4]For example, if there is a Mundell-Tobin effect, the nominal interest rate at the beginning of each period will be set by estimating the impact of the expected change in inflation on the equilibrium real rate of interest and then by adding the best available prediction of inflation to it. It is assumed here that the parameters of the univariate process for the ex-post real rate subsume those of the underlying structural model that is used by market participants to assess the Mundell-Tobin effect. (See, Zellner and Palmer, 1974).

[5]A combined forecast of PA and PM with a negative weight given to PA will be more accurate than PM alone if $\sigma_2/\sigma_1 < \rho$, where σ_1 denotes the standard error of the PA prediction errors, σ_2 that of PM prediction errors and ρ their correlation coefficient. (Granger and Newbold, 1977, p. 270).

[6]Granger and Newbold (1977, p. 272-76) show that more elaborate λ_ts than the one used here do not produce significantly better results.

[7]Granger and Newbold (1974) showed that the estimated correlation between two time series that are independently generated and nonstationary in the mean can be very high.

[8]The squared coherency was estimated by using a Parzen window, 24 lags as the truncation point and 96 observations.

[9]As McCallum (1984) pointed out, the lack of correlation between inflation and interest rates at high frequencies cannot be interpreted as evidence that inflation is unpredictable in the short-run.

[10]I also experimented with the rate of change of both unit labor costs and real wages in order to capture the effect of the real shock; the results, however, were not affected.

[11]Mishkin (1981) found that the inflation rate had a strong impact on the ex-ante real rate of interest, in the U.S. during the post-World War II years.

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Table 1: Inflation Rates in Industrial Countries Based on Consumer Price Indices

(In per cent at monthly rate)

	Jan. 1962/ Dec. 1971	Jan. 1972/ June 1980	Jan. 1962/ Dec. 1971	Jan. 1972/ June 1980
	Variability		Average	
Italy	0.18	1.31	0.35	1.10
United States	0.17	0.70	0.26	0.64
Canada	0.18	0.72	0.25	0.67
Japan	0.33	0.85	0.45	0.74
Austria	0.38	0.53	0.31	0.54
Belgium	0.21	0.59	0.29	0.61
Denmark	0.32	0.90	0.53	0.80
France	0.21	0.84	0.35	0.78
Germany	0.21	0.43	0.24	0.42
Netherlands	0.36	0.61	0.42	0.60
Norway	0.28	0.71	0.40	0.71
Sweden	0.26	0.84	0.39	0.74
Switzerland	0.22	0.42	0.32	0.38
United Kingdom	0.22	1.23	0.38	1.11

Source: IMF, International Financial Statistics.

Note: I arbitrarily use the mean absolute change of the inflation rate as the measure of its variability.

Table 2: Predictors of the Inflation Rate

a) ARIMA model for the inflation rate (π_t):

(1) May 1972 - April 1980

$$(1 - B)\,\pi_t = \hat{\eta}_t - \underset{(.09)}{.42}\,\hat{\eta}_{t-1} - \underset{(.09)}{.35}\,\hat{\eta}_{t-2} \qquad \sigma^2_{\hat{\eta}} = .000036 \qquad Q(10)=3.01$$

(2) May 1972 - August 1978

$$(1 - B)\,\pi_t = \hat{g}_t - \underset{(.11)}{.40}\,\hat{g}_{t-1} - \underset{(.11)}{.32}\,\hat{g}_{t-2} \qquad \sigma^2_{\hat{g}} = .000035 \qquad Q(10)=6.79$$

b) Distributed lag of the inflation rate on past rates of change of the wholesale price index (WP_t):

(3) May 1972 - April 1980

$$\pi_t = \underset{(.0009)}{.0068} + \underset{(.06)}{.43} \sum_{i=1}^{10} WP_{t-i} + \hat{v}_t \qquad \bar{\pi} = .0126 \quad SE = .0052 \qquad \hat{\rho} = .17 \quad \sigma^2_{\hat{\rho}} = .10$$

(4) May 1972 - August 1978

$$\pi_t = \underset{(.0010)}{.0060} + \underset{(.06)}{.43} \sum_{i=1}^{10} WP_{t-i} + \hat{u}_t \qquad \bar{\pi} = .0120 \quad SE = .0043 \qquad \hat{\rho} = .16 \quad \sigma^2_{\hat{\rho}} = .10$$

c) ARIMA model for the ex post real rate of interest (R_t):

(5) May 1972 - April 1980

$$R_t = -.0032 + \hat{z}_t + \underset{(.09)}{.44}\,\hat{z}_{t-1} \qquad \sigma^2_{\hat{z}} = .000029 \qquad Q(11)=1.68$$

(6) May 1972 - August 1978

$$R_t = -.0029 + \hat{d}_t + \underset{(.10)}{.44}\,\hat{d}_{t-1} \qquad \sigma^2_{\hat{d}} = .000029 \qquad Q(11)=4.10$$

Note: Standard errors are in parentheses; $\bar{\pi}$ is the sample mean of the inflation rate; SE is the standard error of the regression; $\hat{\rho}$ is the estimate first order autocorrelation of the residuals; $\sigma^2_{\hat{x}}$ is the estimated variance of variable $\hat{x}$; Q is the modified Box-Pierce statistics to test the hypothesis that the first p sample autocorrelations are zero and it is equal to

$$Q = n(n + 2) \sum_{k=1}^{p} (n - k)^{-1}\, \hat{\Gamma}_k^2$$

where $\hat{\Gamma}_k$ is the k^{th} sample autocorrelation and n is the sample size. The Q statistics is distributed χ^2 with the degrees of freedom given in parenthesis. The lag function of the rates of change of the wholesale price index was approximated by a fourth order Almon polynomial which was constrained to be zero at the far right.

Table 3: A Comparison of Alternative Predictors of Inflation: Out-of-Sample Predictions

(September 1978 - April 1980)

	R^2	$\hat{\rho}_1$	$\hat{\rho}_2$	$\hat{\rho}_3$	$\sigma_{\hat{\rho}}$
(1) $(\pi_t - PA_t) = \underset{(2.27)^*}{2.16}\ (PM_t - PA_t) + \hat{e}_t$	.31	.07	-.03	.03	.22
(2) $(\pi_t - PW_t) = \underset{(2.09)^*}{0.95}\ (PM_t - PW_t) + \hat{e}_t$	.26	.09	-.26	.01	.22

	PM	PA	PW	CF
Mean error	.00163	.00143	.00337	.00143
Mean absolute error	.00419	.00449	.00427	.00426
Root mean squared error	.00579	.00629	.00624	.00610

Note: The t statistics are in parentheses; the asterisk indicates that the coefficient is significantly different from zero at the .05 significance level; $\hat{\rho}_i$ is the i^{th} sample autocorrelation of the fitted residuals and $\sigma_{\hat{\rho}}$ is their standard error.

PM_t = predictions of the financial market equal to $(i_t - \hat{r}_t)/(1 + \hat{r}_t)$; where $\hat{r}_t$ is the proxy for the ex-ante real rate of interest described in the text.

PA_t = predictions obtained from the ARIMA model.

PW_t = predictions obtained from a distributed lag regression on past rates of change in wholesale price index.

CF_t = Predictions obtained by combining the forecasts PM_t and PA_t with time-varying weights as described in the text.

Table 4: The 90 Per Cent Confidence Interval Of the Estimated Square Coherency

Square Coherency	Upper Bound	Lower Bound
.10	.40	.00
.20	.51	.00
.30	.60	.03
.40	.68	.09
.50	.74	.18
.60	.80	.29
.70	.86	.43
.80	.91	.59
.90	.95	.78

Note: The confidence interval was calculated following the suggestions of Jenkins and Watts (1968).

Table 5: Variance Decomposition of the Ex-Ante Real Rate of Interest

Proportion of the variance of the ex-ante real rate accounted for by the orthogonolized innovations of:

	Terms of Trade (TT)	Consumer Price (CP)	Industrial Production (IP)	Monetary Base (MB)	Money--M2 (M)	Nominal Interest Rate (IN)	Non-Monetary Government Debt (GD)	Total
Model A	38	35	7	-	11	9	-	100
Model B	37	32	9	11	-	11	-	100
Model C	38	36	7	-	9	10	-	100
Model D	39	32	8	10	-	11	-	100
Model E	38	28	9	-	10	15	-	100
Model F	39	32	8	10	-	11	-	100
Model G	33	30	9	12	-	10	6	100

Note: All the variables are measured on a comparable scale: the nominal interest rate is the 30-day interbank rate on a monthly basis; the terms of trade, consumer price index, industrial production, money, monetary base and non-monetary government debt are seasonally adjusted monthly rates of change.

The orthogonalization order of the variables is the following:

Model A, TT, M, IN, IP, CP ; Model B, TT, BM, IN, IP, CP ;
Model C, M, IN, IP, CP, TT ; Model D, BM, IN, IP, CP, TT ;
Model E, TT, IP, CP, IN, M ; Model F, IP, BM, IN, CP, TT ;
Model G, GD, BM, IN, IP, CP, TT.

Figure 1. Inflation Rate

(In per cent)

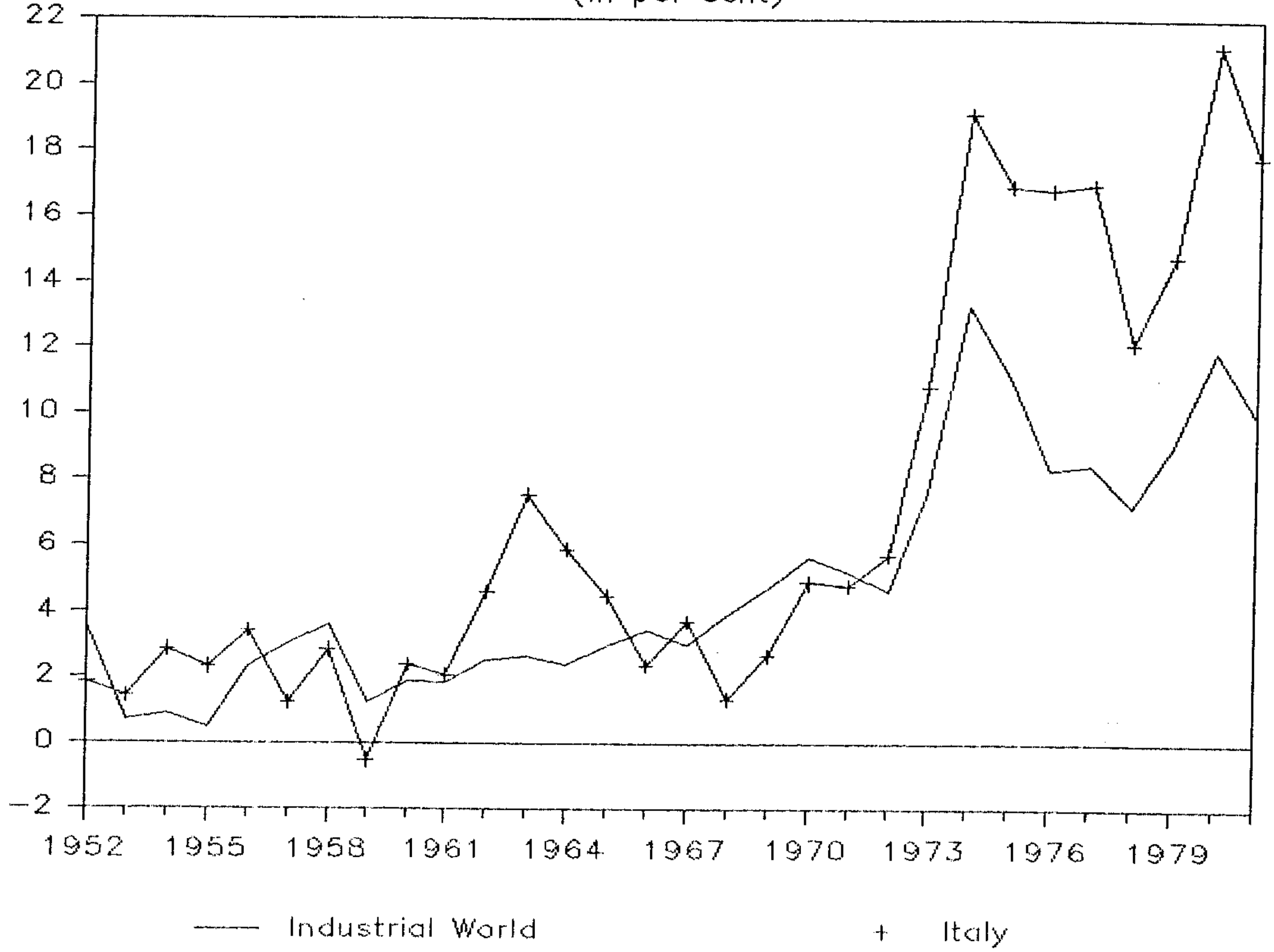


Figure 2. Real Exchange Rate.
(Percentage Changes; 1980=100)

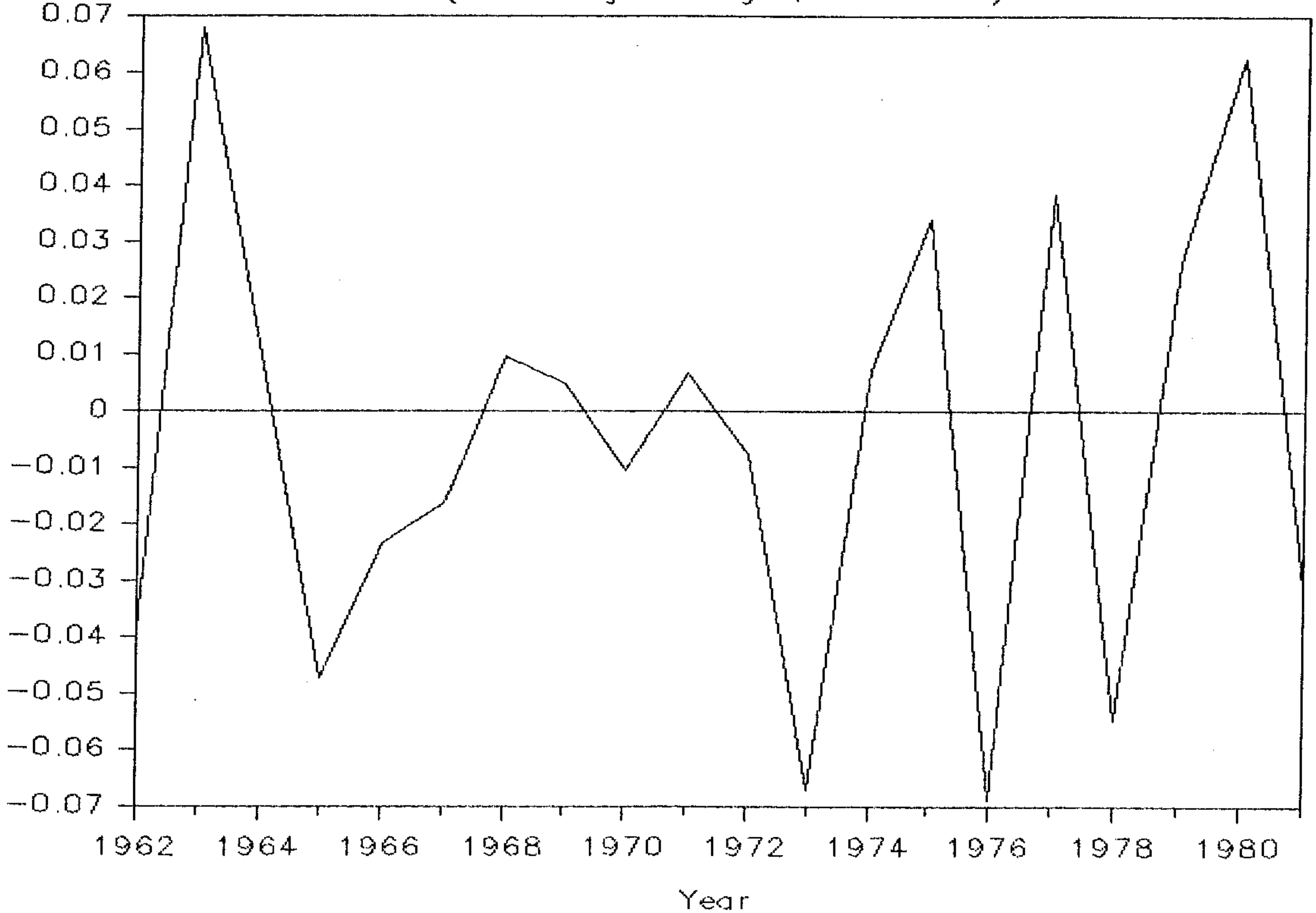


Figure 3. Budget Deficit/GDP

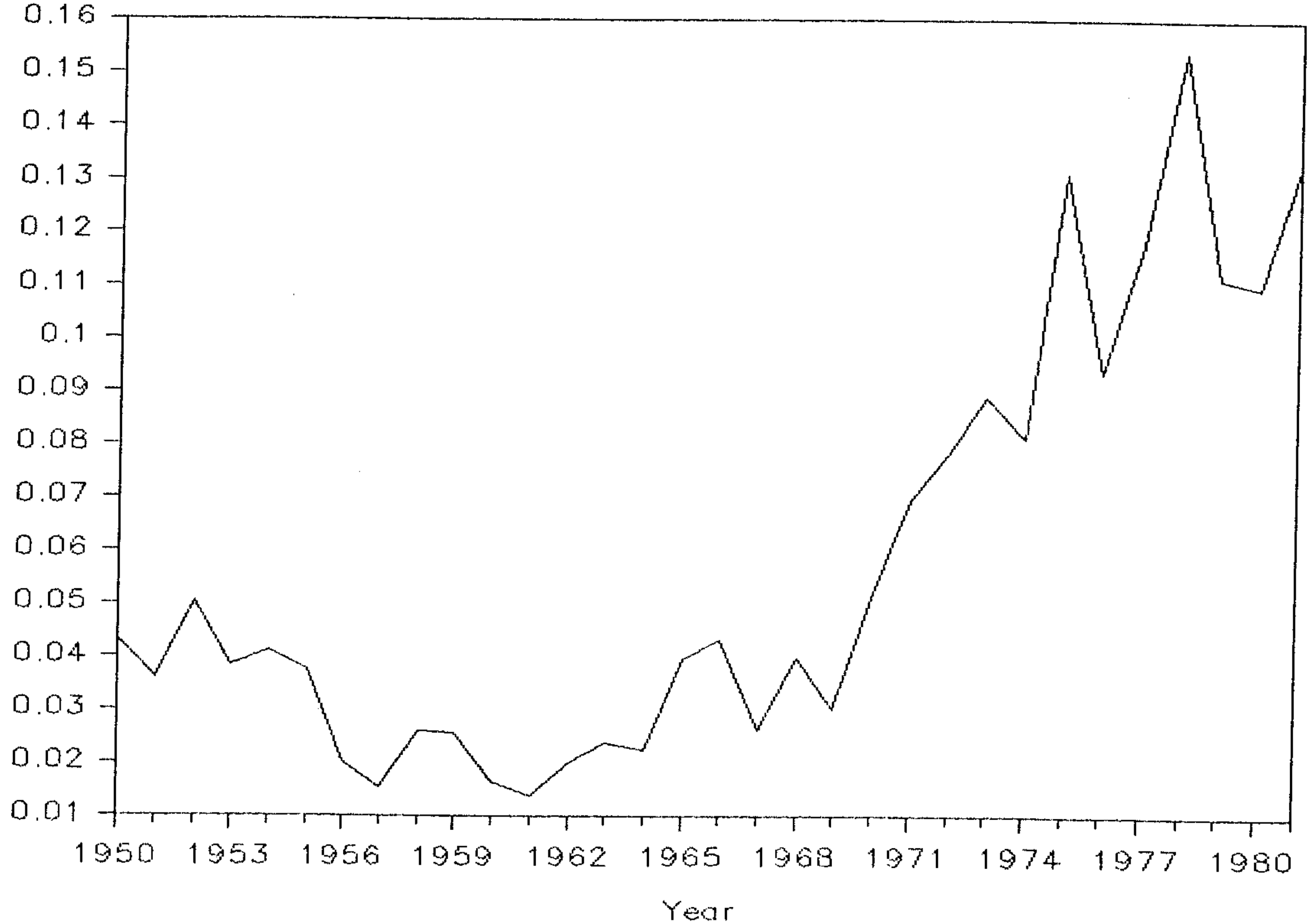
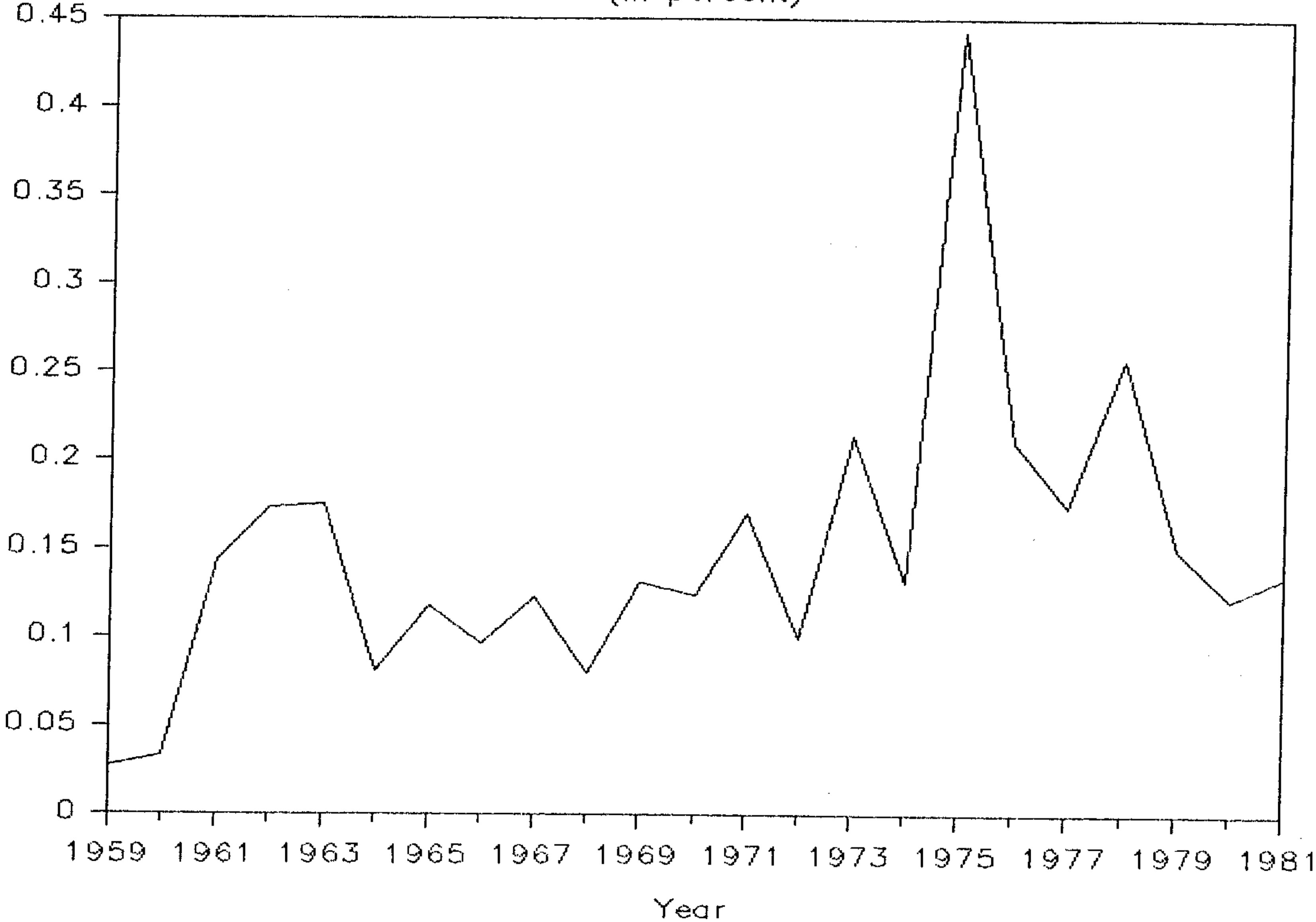


Figure 4. Growth Rate of Base Money

(In percent)



MONTHLY RATE OF CHANGE OF THE COST OF LIVING INDEX AND 30-DAY INTERBANK RATE

(APRIL 1972 - APRIL 1980)

(At annual rate)

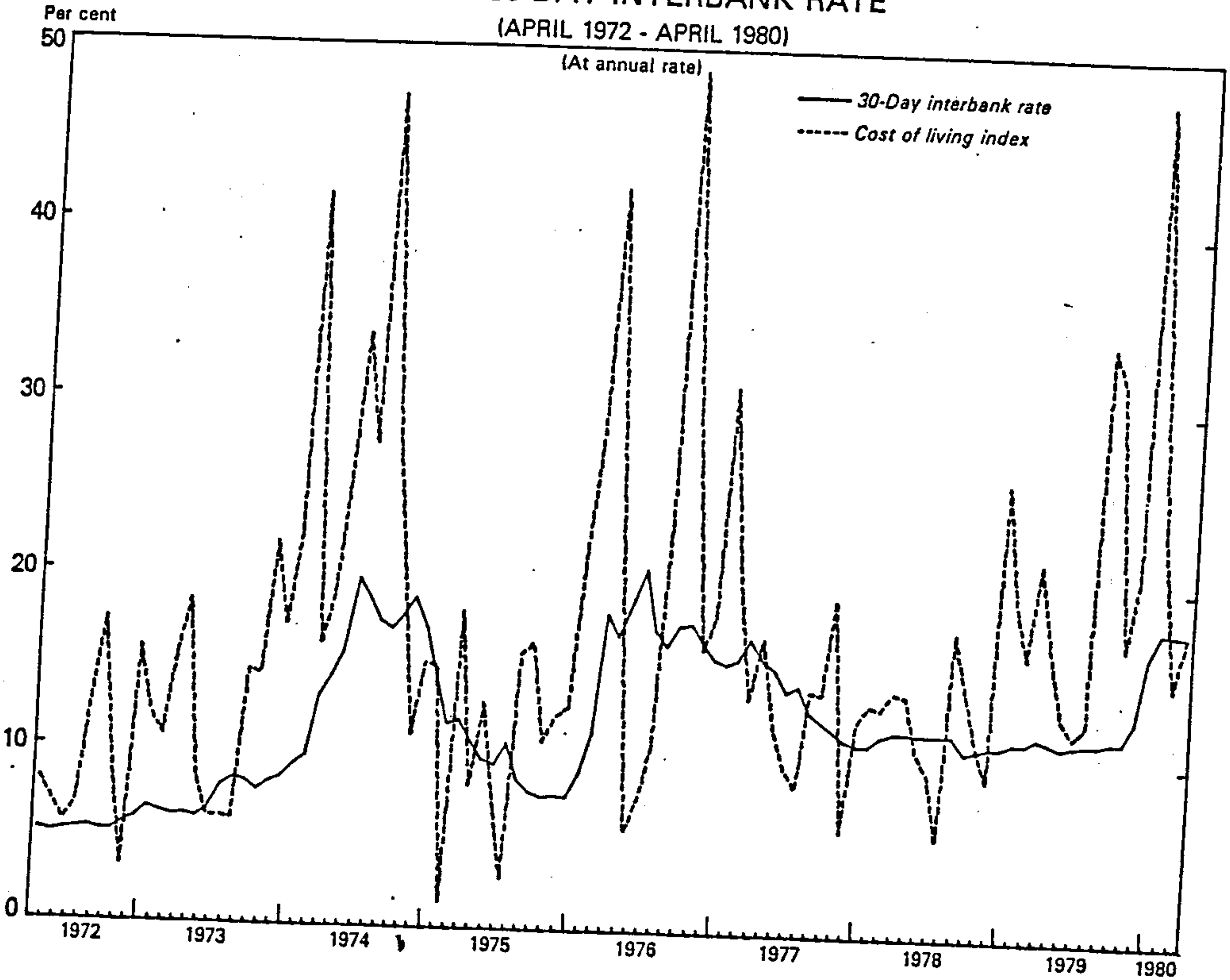
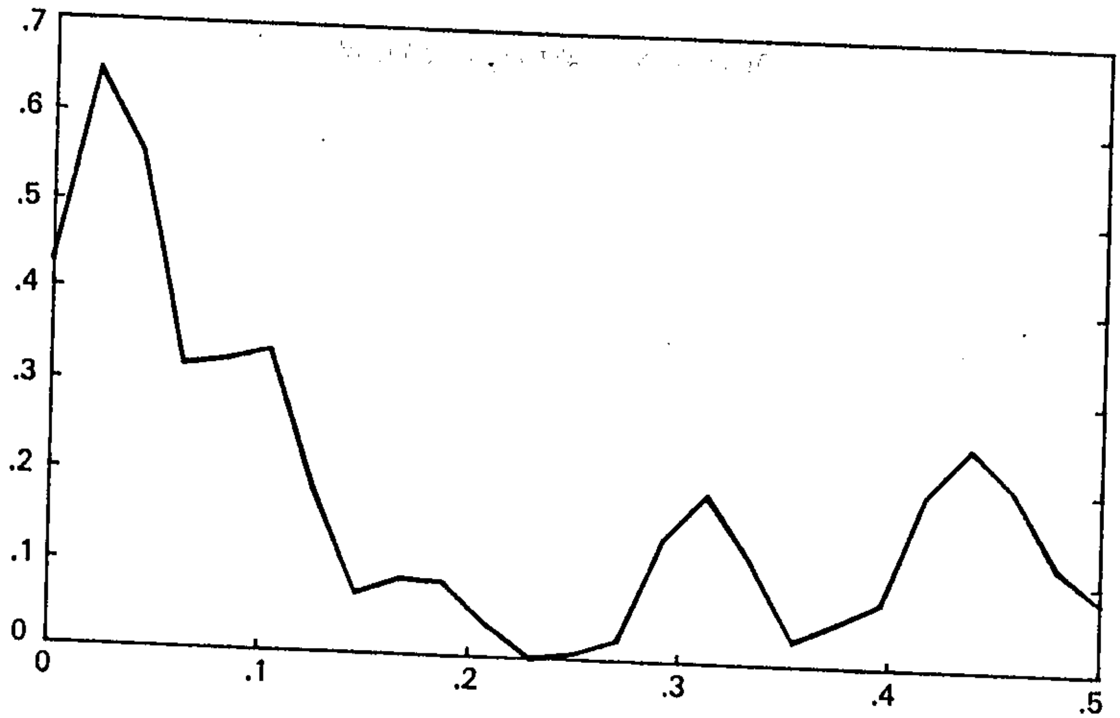


FIGURE 5

Nominal Interest Rate and Inflation: Sample Squared Coherency[1]

(May 1972 - April 1980)



[1]A 90 percent confidence interval for the squared coherency is reported in Table 4

FIGURE 6

Table 2: Predictors of the Inflation Rate

a) ARIMA model for the inflation rate (π_t):

(1) May 1972 - April 1980 $\quad (1 - B)\,\pi_t = \hat{\eta}_t - \underset{(.09)}{.42}\,\hat{\eta}_{t-1} - \underset{(.09)}{.35}\,\hat{\eta}_{t-2} \quad \sigma^2_{\hat{\eta}} = .000036 \quad Q(10)=3.01$

(2) May 1972 - August 1978 $\quad (1 - B)\,\pi_t = \hat{g}_t - \underset{(.11)}{.40}\,\hat{g}_{t-1} - \underset{(.11)}{.32}\,\hat{g}_{t-2} \quad \sigma^2_{\hat{g}} = .000035 \quad Q(10)=6.79$

b) Distributed lag of the inflation rate on past rates of change of the wholesale price index (WP_t):

(3) May 1972 - April 1980 $\quad \pi_t = \underset{(.0009)}{.0068} + \underset{(.06)}{.43}\sum_{i=1}^{10} WP_{t-i} + \hat{v}_t \quad \bar{\pi} = .0126,\ SE = .0052 \quad \hat{\rho} = .17,\ \sigma^2_{\hat{\rho}} = .10$

(4) May 1972 - August 1978 $\quad \pi_t = \underset{(.0010)}{.0060} + \underset{(.06)}{.43}\sum_{i=1}^{10} WP_{t-i} + \hat{u}_t \quad \bar{\pi} = .0120,\ SE = .0043 \quad \hat{\rho} = .16,\ \sigma^2_{\hat{\rho}} = .10$

c) ARIMA model for the ex post real rate of interest (R_t):

(5) May 1972 - April 1980 $\quad R_t = -.0032 + \hat{z}_t + \underset{(.09)}{.44}\,\hat{z}_{t-1} \quad \sigma^2_{\hat{z}} = .000029 \quad Q(11)=1.68$

(6) May 1972 - August 1978 $\quad R_t = -.0029 + \hat{d}_t + \underset{(.10)}{.44}\,\hat{d}_{t-1} \quad \sigma^2_{\hat{d}} = .000029 \quad Q(11)=4.10$

Note: Standard errors are in parentheses; $\bar{\pi}$ is the sample mean of the inflation rate; SE is the standard error of the regression; $\hat{\rho}$ is the estimate first order autocorrelation of the residuals; $\sigma^2_{\hat{x}}$ is the estimated variance of variable $\hat{x}$; Q is the modified Box-Pierce statistics to test the hypothesis that the first p sample autocorrelations are zero and it is equal to

$$Q = n(n + 2) \sum_{k=1}^{p} (n - k)^{-1} \hat{\Gamma}_k^2$$

where $\hat{\Gamma}_k$ is the k^{th} sample autocorrelation and n is the sample size. The Q statistics is distributed χ^2 with the degrees of freedom given in parenthesis. The lag function of the rates of change of the wholesale price index was approximated by a fourth order Almon polynomial which was constrained to be zero at the far right.

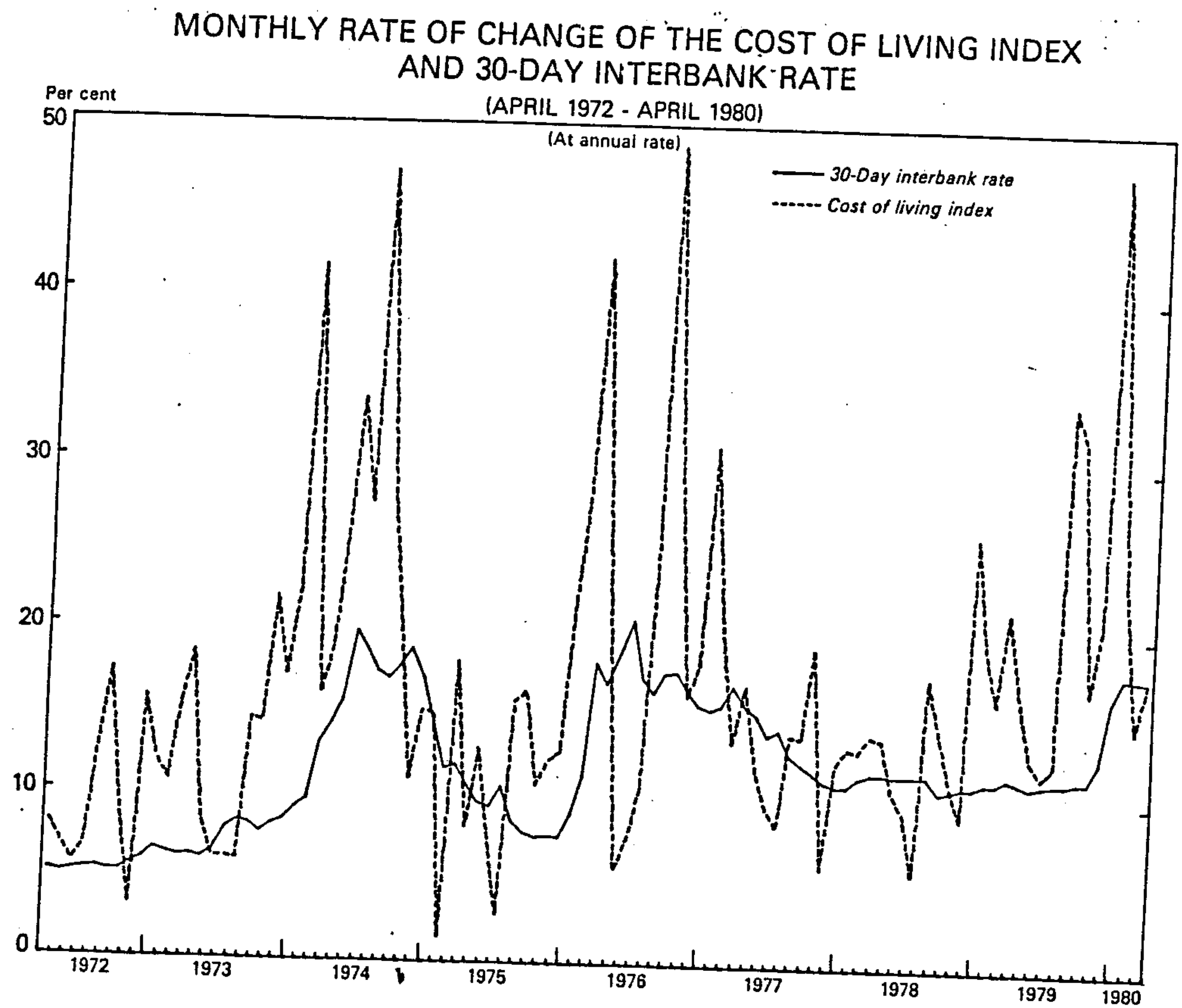

MONTHLY RATE OF CHANGE OF THE COST OF LIVING INDEX
AND 30-DAY INTERBANK RATE
(APRIL 1972 - APRIL 1980)
(At annual rate)
Per cent
50
40
30
20
10
0
30-Day interbank rate
Cost of living index
1972
1973
1974
1975
1976
1977
1978
1979
1980

FIGURE 5