THE FISHER EFFECT AND THE TERM STRUCTURE OF INTEREST RATES

By YOSHINORI SHIMIZU*

Introduction

More than eighty years ago, I. Fisher had already pointed out the possibility of a divergence between the nominal rate of interest (i) and the real rate of interest (r) by the amount of the expected rate of inflation ($\dot{p}^e$). He found the following relationship among these variables; $i=\dot{p}^e+r$.[1] This relationship derived from the fact that, if lenders and borrowers could perfectly foresee future price level movements, lenders would hedge against changes in the real value of their loan principal by adding the percentage change in prices over the life of the loan to the interest charge; borrows, expecting money income to change in proportion to prices, would readily accept the higher rate. Though Fisher's assertion itself was not generally accepted at that time, this relationship has been lighlighted in recent years, and it is already established that the Fisher effect has especially important empirical relevance.

Nevertheless, in Japan there has been little systematic research on the relationship between inflation and interest rates so far. The purpose of this paper is to examine the validity of the Fisher effect in Japan from several points of view. In doing so, we adopted a different approach from the conventional one. Due to the lack of direct observations of both real interest rates and expected rates of inflation, every reseracher on this subject presupposes some specific hypothesis about the formation of expectations. According to the conventional approach, however, the result critically depends on the validity of the assumption about expectations. So, we will examine the relationship without any *a priori* assumption about the formation of expectations.

In Chapter 1 we first examine the effect of lagged monthly actual inflation rates on interest rates and then estimate the formation of expectations based on the result. We also show that the result obtained can be explained consistently by the expectations theory of the term structure of interest rates which was also originated by I. Fisher. In section 2 we analyze the relationship between expected inflation and the term structure of interest rates and compare the empirical validity of our hypothesis with that of the conventional theory. In Chapter 2 we focus our attention on the behavior of the real interest rate. Section 1 surveys the results obtained by Gibson and Fame for the U.S. economy, and Section 2 states our analytical framework i.e., the theory of the efficient market by Fama. Section 3 shows the empirical result for Japan and examines the validity of the Fisher effect in Japan by comparing our with the results obtained for the U.S.

* Lecturer (*Kōshi*) of Money and Banking.

[1] Fisher [9] [10] [11].

Chapter 1. Inflation and the Term Structure of Interest Rates

Section 1. Inflationary Expectations and Interest Rates

1-1. Previous Studies

An ordinary test of the effect of inflationary expectations on interest rates depends on the following two hypothesized relationships:

$$i_t = \dot{p}_t^e + r_t \tag{1}$$

$$\dot{p}_t^e = \sum_{i=t}^{t-n} w_i p_i \tag{2}$$

(1) states that the nominal interest rate (i_t) prevailing at time t for a particular debt instrument is equal to the annual rate of change in prices expected at time t to occur over the life of the instrument ($\dot{p}_t^e$) plus its "real" rate of interest (r_t). (2) is an application of the theory of "adaptive expectations", "error-learning", or alternatively "extrapolative forecasting". Faced with uncertainty about the future, an economic decision-making unit is presumed to base its predictions about future price movements on a weighted average of current and past changes in prices. The weights usually assigned decline monotonically as we go back in time. Substituting equation (2) into equation (1), we obtain the equation usually estimated:

$$i_t = \sum_{i=t}^{t-n} w_i p_i + r_t \qquad (3)^2$$

Employing this equation, however, we cannot identify the validity of each equation separately, for it tests the two relationships (1) and (2) simultaneously. Hence, we adopt a different approach in our analysis.

Using annual and quarterly data for the U.S., Fisher found very long mean lags for the effect of price changes on long- and short-term interest rates. For example, the highest correlation between commercial paper rates and rates of change in the WPI for 1915–27 was obtained when the latter was lagged over 120 quarters (30 years), implying a mean lag of about 40 quarters (10 years). At that time, this kind of result was thought to be unrealistic. This empirical aspect, not the theoretical one was the reason why Fisher's assertion was not generally accepted.

Recently, there has been a considerable revival of interest in the relationship, with the use of additional data and sophisticated estimation techniques. As a result, the effect itself of price changes on interest rates has been recognized, with a variety of opinions about the magnitude of the effect, the length of the lag, and the relationship with the real rate.[3] Most

[2] Though there are several variations of this equation, the following argument holds irrespective of the hypothesized types of the expectational formation.

[3] See Cagan [2], Eisner [4], Friedman [14] [15], Friedman & Schwartz [16], Gibson [18] [19] [20], Hamberger & Silber [23], Lahiri [27], Meiselman [31], Sargent [35] [36], Yohe & Karnosky [40].

of these studies have been based on equation (3), and data intervals have ranged from months to business-cycle phases. Lagged rates of change in various price level indices and even nominal income and a direct survey have been tried as indicators of price expectations. Many studies have found mean lags to be so long that we can almost neglect recent price changes. For example, Friedman and Schwartz [16] found a mean lag of about 10 years for short-term rates and 25–30 years for long-term rates. In contrast, Yohe and Karnosky [40] obtained a result based on the period 1952–1969 which indicates that the lags are very short, with most of the effect of price level changes for both long- and short-term rates occurring within two years. Unfortunately, however, the theoretical explanations given to their results are unsatisfactory.[4] In the follwing section we first analyze this relationship using Japanese monthly data and then give a theoretical explanation which can explain the result obtained.

1–2. Empirical Analysis

The conventional test of the Fisher effect based on (3) puts a very strong restriction on the formation of expectations and does not clarify the relative importance of the relationships (1) and (2). But it is certainly realistic to suppose that expectations about future price level movements would be formed based in some way on past experience. If only we could identify how current rates of interest are affected by past price movements the measurement of the expected rate of inflation itself might be dispensable. Thus we first examine the relationship between market rates and past price movements directly and then attempt to find some implications about the formation of expectations from the result. Instead of the conventional equation (3), we ran regressions according to an equation with the single independent variable $\dot{p}_{t-i}$ based on monthly data. For both the short-term rate and the long-term rate we estimated the effect of the actual inflation rate in each lagged month on the market rate based on equations (1–4) and (1–5) respectively.

$$i_t^S = \alpha_1 + \beta_1 \dot{p}_{t-i} + \varepsilon_t \tag{1–4}$$

$$i_t^L = \alpha_2 + \beta_2 \dot{p}_{t-i} + \varepsilon_t' \tag{1–5}$$

The period analyzed is from January 1966 to March 1976.[5] We used the average call money rate[6] (unconditional, Tokyo) as the short-term rate (i_t^S) and the average final yield of long-term bonds[7] as the long-term rate (i_t^L). For the rate of inflation, we used two indices:[8] the monthly rate of change in the CPI excluding seasonal goods (all Japan), and the rate of change in the CPI from the corresponding month in the preceeding year (all Japan, all commodities).

[4] For example, Friedman & Schwartz [16] attributes their very long lags to "slow and gradual adjustment of anticipations of price changes to the actual behavior of prices". As for the assertion of Yohe & Karnosky [40], see 1–3 of this chapter.

[5] We have obtained better result in this period than in the earlier period. According to Yohe & Karnosky, they have found that the price expectation effect is much larger in the 1961–69 period than in the earlier period, and suggested that some institutional change have occured during 60's. If this is the case also for Japan, the dividing line should be around 1966.

[6] Source; *Economic Statistics Monthly*, Statistics Department, The Bank of Japan.

[7] Source: *Monthly Statistics Report*, Research & Statistics Department. Tokyo Stock Exchange.

[8] Source: *Economic Statistics Monthly*, Statistics Department, The Bank of Japan.

Results

(1): Table 1–1 shows the results of the regressions (1–4) and (1–5), and the series of coefficients obtained for each lag β_1 and β_2 are plotted in Figure 1–1. In spite of the differences in magnitude, the time pattern of β_1 and β_2 are similar, and both the short-term and the long-term rate are most strongly affected by the rate of inflation 3–12 months ago. The coefficients, t-values and $\bar{R}^2$s all decline for both shorter and longer lags and coefficients for the short-term and the long-term rate are not significant at the 5% level for lags more than 23 months and 28 months respectively. Thus the effect of inflation on both interest rates fades away in about two years. Though the timing of the effect of inflation on both

TABLE 1–1

($\dot{P}_{t-i}$: Monthly rate of change in CPI)

(A) $i_t^S = \alpha_1 + \beta_1 \dot{p}_{t-i}$

	β_1	t-value	$\bar{R}^2$	α_1
t	1.37*	4.81	0.15	6.83
$t-1$	1.49*	5.31	0.18	6.75
$t-2$	1.67*	6.13	0.23	6.63
$t-3$	1.82*	6.91	0.28	6.53
$t-4$	1.87*	7.17	0.29	6.49
$t-5$	1.92*	7.47	0.31	6.46
$t-6$	1.94*	7.59	0.32	6.46
$t-7$	1.95*	7.55	0.32	6.47
$t-8$	2.04*	8.01	0.34	6.39
$t-9$	2.01*	8.34	0.36	6.35
$t-10$	2.11*	8.46	0.37	6.34
$t-11$	2.05*	8.15	0.35	6.40
$t-12$	1.92*	7.33	0.30	6.48
$t-13$	1.74*	6.39	0.25	6.61
$t-14$	1.71*	6.24	0.24	6.63
$t-15$	1.61*	5.77	0.21	6.69
$t-16$	1.37*	4.74	0.15	6.84
$t-17$	1.26*	4.31	0.13	6.92
$t-18$	1.32*	4.20	0.12	6.92
$t-19$	1.15*	3.57	0.09	7.04
$t-20$	0.99*	3.05	0.06	7.13
$t-21$	0.86*	2.58	0.04	7.22
$t-22$	0.74*	2.21	0.03	7.30
$t-23$	0.61	1.80	0.02	7.39
$t-24$	0.47	1.37	0.01	7.47

(B) $i_t^L = \alpha_2 + \beta_2 \dot{p}_{t-t}$

	β_2	t-value	$\bar{R}^2$	α_2
t	0.52*	4.41	0.13	7.73
$t-1$	0.58*	5.07	0.17	7.69
$t-2$	0.63*	5.64	0.20	7.65
$t-3$	0.67*	6.08	0.23	7.63
$t-4$	0.66*	5.96	0.22	7.63
$t-5$	0.70*	6.48	0.25	7.60
$t-6$	0.74*	6.94	0.28	7.58
$t-7$	0.74*	6.96	0.28	7.58
$t-8$	0.81*	7.75	0.33	7.54
$t-9$	0.83*	8.10	0.35	7.52
$t-10$	0.84*	8.27	0.36	7.51
$t-11$	0.86*	8.55	0.37	7.50
$t-12$	0.83*	7.94	0.34	7.54
$t-13$	0.76*	7.04	0.29	7.58
$t-14$	0.72*	6.60	0.26	7.60
$t-15$	0.67*	6.02	0.23	7.63
$t-16$	0.63*	5.55	0.20	7.66
$t-17$	0.60*	5.16	0.18	7.68
$t-18$	0.65*	5.25	0.18	7.66
$t-19$	0.62*	4.95	0.16	7.69
$t-20$	0.59*	4.71	0.15	7.71
$t-21$	0.55*	4.24	0.12	7.74
$t-22$	0.51*	3.91	0.11	7.77
$t-23$	0.46*	3.42	0.08	7.80
$t-24$	0.40*	2.91	0.06	7.85
$t-25$	0.34*	2.48	0.04	7.88
$t-26$	0.32*	2.13	0.03	7.90
$t-27$	0.35*	2.11	0.03	7.89
$t-28$	0.34	1.79	0.02	7.90

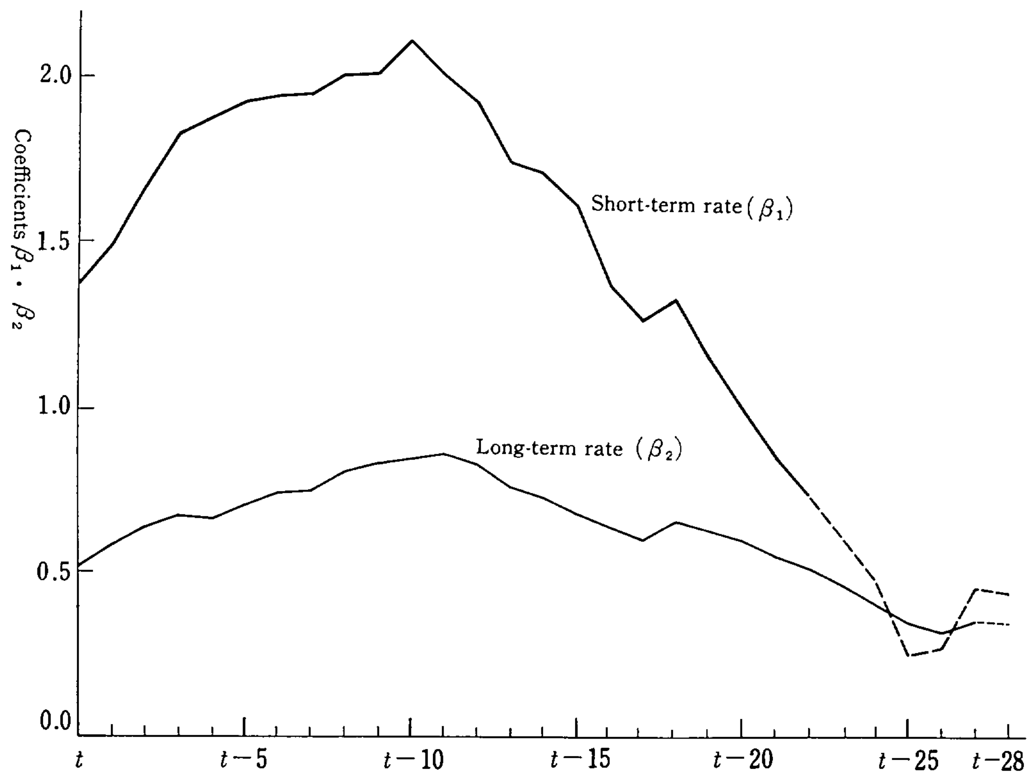
* denotes the significance at the 5% level.

rates is almost the same, the short-term rate is two or three-times more sensitive to the past rate of inflation than the long term rate. But the effect of inflation in each single lagged month is naturally small and explained at most one-third of the variation in both interest rates.

Figure 1–1 shows that in Japan the effect of inflation on interest rates does not decline monotonically with the passage of time. Contrary to the results for the U.S.,[9] the weights are relatively light for recent experience and become heavier for older experience within a year. The weights are heaviest for 10–11 month lags and rapidly decline for lags more than 12 months. This might suggest the necessity of different hypothesis on the formation of expectations in Japan, since in the U.S. most studies obtained monotonically declining weights for older experiences.

(2): From Analysis (1) we found that past inflation rates within a year have a dominant effect on both the long-term and the short-term interest rate. Based on this result, we ran different regressions according to the same equations (1–4) and (1–5) using the rate of change in the CPI from 12 months earlier. These results are shown in Table 1–2 and Figure 1–2.

FIGURE 1–1



Solid line denotes the significance at the 5% level.

[9] e.g. Yohe & Karnosky [40].

TABLE 1-2

($\dot{P}_{t-i}$: Rate of change in CPI from corresponding month in preceeding year)

(A) $i_t^S = \alpha_3 + \beta_3 \dot{p}_{t-i}$

	β_3	t-value	$\bar{R}^2$	α_3
t	0.33*	17.09	0.70	4.93
$t-1$	0.33*	18.81	0.71	4.86
$t-2$	0.34*	19.84	0.76	4.82
$t-3$	0.34*	20.06	0.77	4.82
$t-4$	0.34*	19.43	0.76	4.85
$t-5$	0.33*	17.83	0.72	4.92
$t-6$	0.32*	16.08	0.77	5.01
$t-7$	0.31*	14.32	0.63	5.12
$t-8$	0.29*	12.59	0.56	5.26
$t-9$	0.27*	10.89	0.31	5.45
$t-10$	0.25*	9.39	0.42	5.61
$t-11$	0.23*	8.06	0.34	5.81
$t-12$	0.21*	6.81	0.27	6.02
$t-13$	0.18*	5.77	0.21	6.23
$t-14$	0.16*	4.79	0.15	6.45
$t-15$	0.13*	3.92	0.11	6.66
$t-16$	0.12*	3.26	0.01	6.83
$t-17$	0.10*	2.74	0.05	6.95
$t-18$	0.09*	2.26	0.03	7.07
$t-19$	0.07	1.72	0.02	7.22
$t-20$	0.05	1.21	0.00	7.36
$t-21$	0.03	0.75	0.00	7.45
$t-22$	0.02	0.38	0.01	7.62
$t-23$	0.00	0.03	0.01	7.73
$t-24$	-0.02	-0.29	0.01	7.85

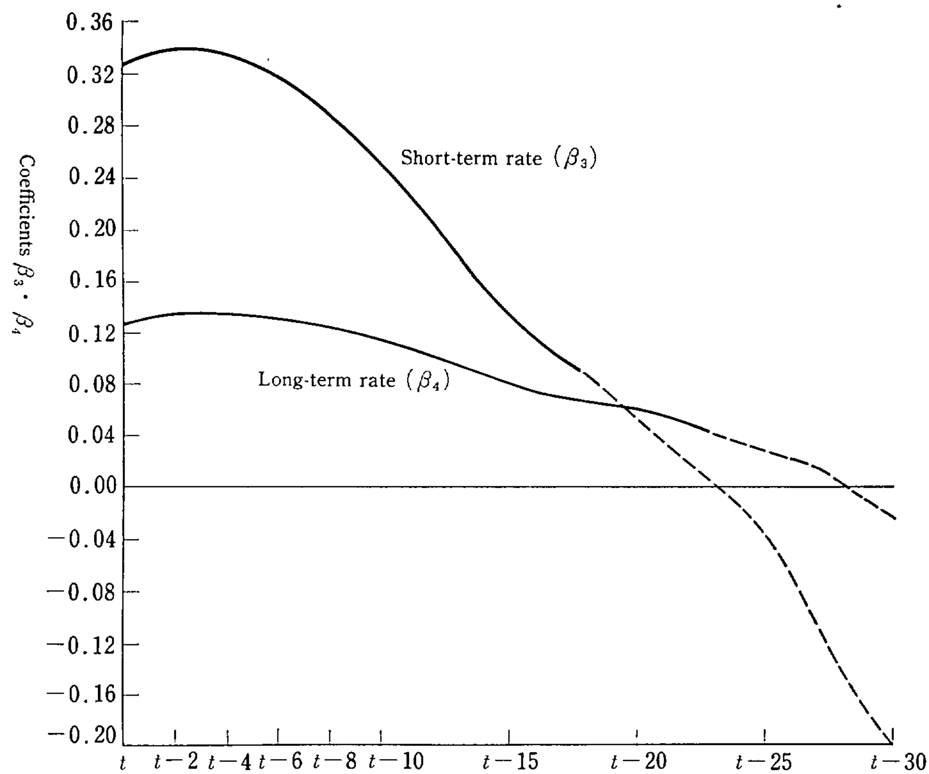
(B) $i_t^L = \alpha_4 + \beta_4 \dot{p}_{t-i}$

	β_4	t-value	$\bar{R}^2$	α_4
t	0.125*	14.63	0.64	6.99
$t-1$	0.130*	16.31	0.69	6.95
$t-2$	0.132*	17.38	0.71	6.93
$t-3$	0.133*	17.88	0.73	6.93
$t-4$	0.132*	17.81	0.72	6.93
$t-5$	0.132*	17.42	0.71	6.94
$t-6$	0.130*	16.43	0.69	6.96
$t-7$	0.127*	15.35	0.66	6.99
$t-8$	0.124*	14.28	0.63	7.02
$t-9$	0.120*	13.02	0.58	7.06
$t-10$	0.114*	11.55	0.52	7.11
$t-11$	0.108*	10.38	0.47	7.16
$t-12$	0.101*	9.15	0.41	7.23
$t-13$	0.094*	8.00	0.34	7.29
$t-14$	0.086*	7.05	0.28	7.37
$t-15$	0.079*	6.20	0.24	7.43
$t-16$	0.074*	5.59	0.20	7.48
$t-17$	0.071*	5.03	0.17	7.52
$t-18$	0.068*	4.53	0.14	7.55
$t-19$	0.063*	4.03	0.11	7.60
$t-20$	0.059*	3.53	0.09	7.64
$t-21$	0.054*	3.03	0.06	7.69
$t-22$	0.048*	2.54	0.04	7.74
$t-23$	0.041*	2.00	0.02	7.80
$t-24$	0.034	1.51	0.01	7.85

* denotes the singnificance at the 5% level.

In this case, the short term rate is most strongly affected by the annual rate of inflation 1–4 months before. Though the coefficients (β_3), t-values, $\bar{R}^2$s all decline gradually for longer lags, the annual rate of change in the CPI explains more than one-half of the total variation in the interest rate for lags less than 8 months. Since the coefficients for lags more than 19 months are not significant at the 5% level, the effect of the annual rate of change in the CPI on the short-term rate fades away in 18 months. The long-term rate is most strongly affected by the annual rate of inflation 1–6 months before. Similar to what was found with the short-term rate, the coefficients (β_4), t-values, R^2s all decline gradually for longer lags, and the 12-month rate of change in the CPI explains more than one-half of the total variation in the interest rate for lags less than 10 months. Since the coefficients for lags more than 24 months are not significant at the 5% level, the effect of the annual rate of change in the CPI for each month on the long term rate fades away in 23 months. A 1% change in the annual

FIGURE 1-2



Solid line denotes the significance at the 5% level.

rate of change in the CPI of preceding several months affects the long term rate by 0.12–0.13% and the short-term rate by 0.30–0.33%. Thus, in this case too the short-term rate is two and a half-times more sensitive than the long-term rate. The rate of decline in the coefficients is larger for the short-term rate than the long-term rate, implying that experiences in the more distant past have relatively larger effects on longer term expectations.

(3): Figure 1-3 shows the coefficients of regressions (1–6) and (1–7) which are similar to the equation used by Yohe and Karnosky [40].

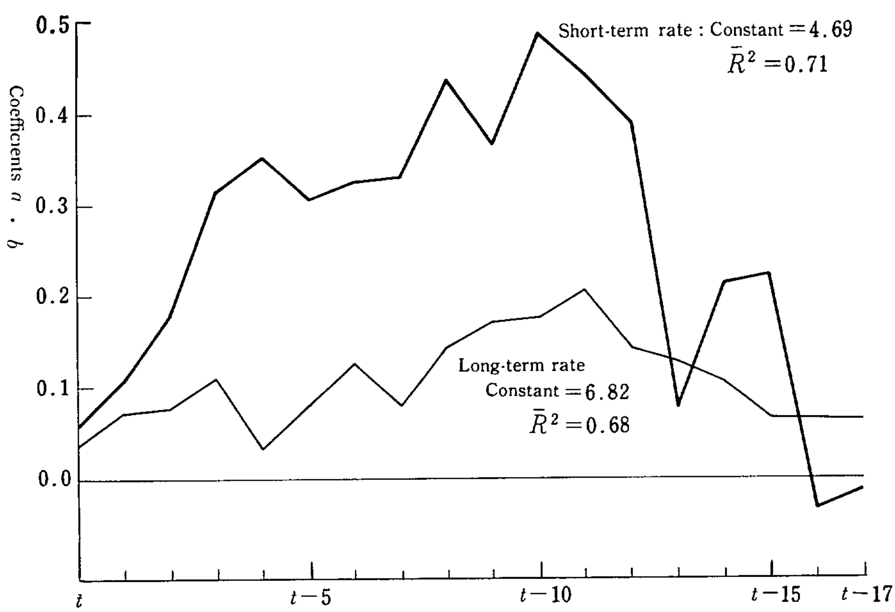
$$i_t^S = a_0 + a_1\dot{p}_t + a_2\dot{p}_{t-1} + \cdots a_{18}\dot{p}_{t-17} + \mu_t \tag{1-6}$$

$$i_t^L = b_0 + b_1\dot{p}_t + b_2\dot{p}_{t-1} + \cdots b_{18}\dot{p}_{t-17} + \mu_t' \tag{1-7}$$

In this case, the estimated coefficients are not necessarily reliable due to the high serial correlation among the independent variables, i.e., the existence of the multicollinearity. Nevertheless, the behavior of these coefficients is consistent with the results of Analyses (1) and (2), and it might confirm our estimates in these analyses. What is interesting here is the magnitude of the constant term estimated for (1–6) and (1–7) i.e., 4.69 for the short-term rate and 6.82 for the long-term rate. We will discussed the importance of the constant term in the next chapter.

FIGURE 1-3

$$i_t^S = a_0 + a_1 \dot{p}_t + a_2 \dot{p}_{t-1} + \cdots\cdots + a_{18} \dot{p}_{t-17}$$
$$i_t^L = b_0 + b_1 \dot{p}_t + b_2 \dot{p}_{t-1} + \cdots\cdots + b_{18} \dot{p}_{t-17}$$
($\dot{p}_{t-i}$: Monthly rate of change in CPI)



1-3. Implication of the Findings

The main results from the above analyses are summarized as follows. 1) The effect of inflation on the market interest rates works within two years. 2) The time pattern of the effect of past inflation on both the short-term and the long-term rate is almost the same, but the short-term rate is much more sensitive than the long-term rate. 3) In forming price expectations the weights attached to past inflation do not decline monotonically with the length of the lag. Except for point 3, these findings are a most completely consistent with the results of Yohe and Karnosky [40]. In contrast, as noted above, earlier studies yielded much longer lags. In order to reconcile this difference. Yohe and Karnosky suggested the following three hypotheses:

(1) The "true" lags of interest rates behind price changes are short, so that biases arise in aggregating the interest-rate and price-change series over longer observation periods which lead to systematic overestimates of the length of the lags.

(2) The forms of the lags estimated in other studies, in contrast to the more flexible class of lags estimated in this study, are biased toward yielding longer average lags.

(3) Institutional changes have occurred over time in financial and real markets, with the result that price-level changes have come to have prompter and larger effects on interest rates. To put it differently, there has been considerable thinning of the "molasses (long-lag) world", particularly in the past decade. ([40], pp. 26–27)

Since our analysis also uses relatively recent monthly data and employs the most flexible method in estimating lags without any *a priori* assumption about the formation of expectations, our result reinforces the three hypotheses of Yohe and Karnosky.

These three hypotheses cannot be regarded as a theoretical explanation for our results, however. What is important in our result is the fact that the time patterns of the effect of past inflation on both the long-term and the short-term rate are almost the same, although the short-term rate is more sensitive than the long-term rate. The earlier studies discussed in 1–1 implicitly assumed that (1–3) holds for the short-term rate and (1–4) for the long-term rate independently and that the difference in the response of both rates stems from the difference in the length of the lags and the weights in forming expectations. Put differently, the long term expectation has been assumed to have longer distributed lags than the short term expectations, and indeed research on this problem has been focused on the sophistication of the hypothesis about the formation of expectations.

In order to explain our result, we should not analyze the effect on the short-term and the long-term rate separately. We must take account of the relationship between the long-term and the short-term rate i.e., the term structure of interest rates. In the next section, we will show that the effect of inflation on market interest rates should be understood in connection with the term structure and that we can explain the result of our analysis consistently by taking it explictly into account.

1–4. Inflationary Expectations and the Expectations Hypothesis of the Term Structure of Interest Rates

A theory which is very important in explaining our result is the expectations hypothesis of the term structure of interest rates. It is interesting to note that this theory as well as equation (1–1) was advocated first by I. Fisher. The expectations hypothesis about the term structure of interest rates follows from the assumptions that short-term and long-term securities can be treated as if they were perfect substitutes and that transactors, indifferent to uncertainty and having similar expectations, equate the forward rates in the market to the expected rates.

Let ${}_ir_j$ represent the forward one-period short-term rate of j-th period expected at period i ; ${}_iR_j$, the actual rate prevailing in the market at period i for a security of j periods to maturity; ${}_i\dot{p}_j^e$, the rate of inflation in the j-th period expected at the i-th period. Let us consider the case in which recent rates of inflation have been rising. In this case a series of ${}_i\dot{p}_j^e$s are represented by the curve in Figure 1–4. Under the prevalence of this expectation, forward one-period short-term rates ${}_ir_j$ rise by the expected rate of inflation in the corresponding period ${}_i\dot{p}_j^e$. For simplicity suppose that real one-period short-term rates remain constant for all periods. Then, take the ordinate of ${}_ir_j$ in such a manner that ${}_ir_j$ which is equal to the real rate plus ${}_i\dot{p}_j^e$ is represented by the same curve with ${}_i\dot{p}_j^e$. Thus, the curve labeled ${}_i\dot{p}_j^e$ in Figure 1–5 shows the pattern of forward one-period short term rates that have taken the inflationary expectations into account.

According to the expectations hypothesis, the current actual rate on the security which is t periods to maturity (${}_1R_t$) would be determined in such a manner that areas A and B in Figure 1-5 would be equal through an arbitrage between ${}_1R_t$ and the implied forward rates on one-period securities, i.e.,

$$\sum_{j=1}^{k} ({}_1r_j - {}_1R_t) = \sum_{j=k+1}^{t} ({}_1R_t - {}_1r_j) \tag{1-8}$$

This implies the following equation:

$$\begin{aligned}(1+{}_1R_t) &= \sqrt[t]{(1+{}_1r_1)(1+{}_1r_2)\cdots(1+{}_1r_t)} \\ &= \sqrt[t]{\prod_{j=1}^{t}(1+{}_1r_j)} \end{aligned} \tag{1-9}$$

FIGURE 1-4

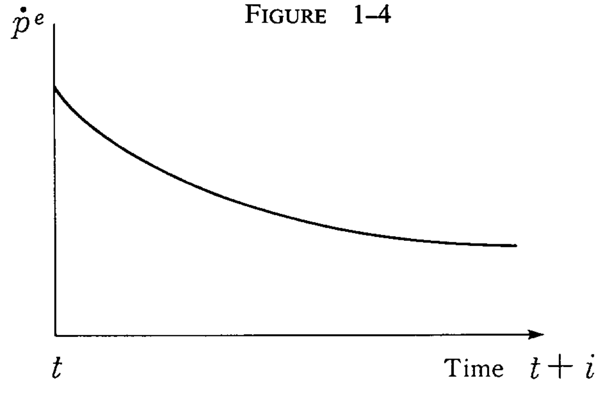
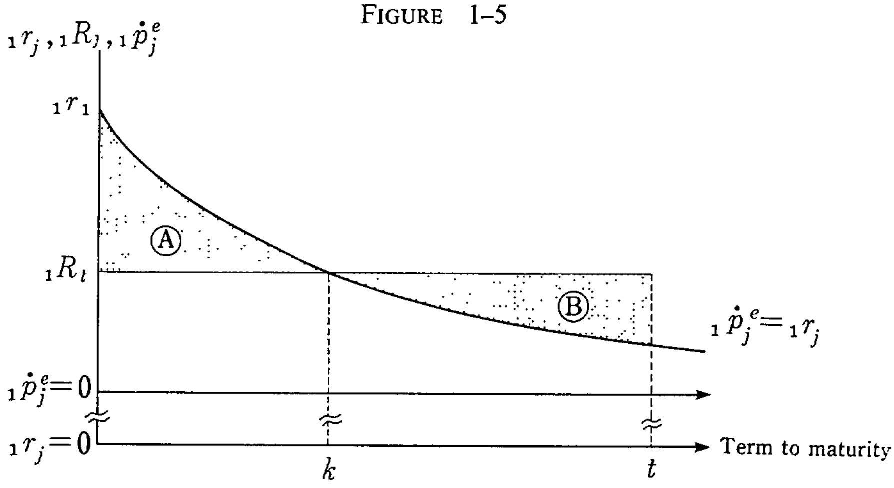


FIGURE 1-5

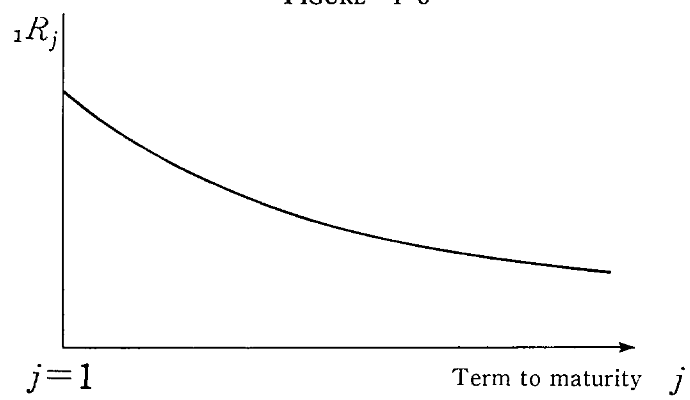


The relationship between a given term structure of rates and the implied forward short term rates is analogous to the well-known relationship between average and marginal quantities. Given curve ${}_1r_j$ in Figure 1–5, the longer the periods to maturity of long term securities, the larger the difference between the short-term rate ${}_1r_1$ and the long term rates ${}_1R_t$. The yield curve is downward sloping as in Figure 1–6.

Now, let's suppose that after several month's rise in the rate of inflation the short-term inflationary expectation has risen while the long-term expectation has been only negligibly affected due to the smaller weights given to more recent experiences. In Figure 1–7, curve ${}_1\dot{p}_j{}^e$ would rise only partially as shown by the dotted curve, keeping ${}_1\dot{p}_j{}^e$ for $j>k$ unchanged. By assumption curve ${}_1r_j$ shows the same pattern. Then, the actual long-term rate at period 1 (${}_1R_t$) would also rise from ${}_1R_t$ to ${}_1R_t'$. According to equation (1–9), ${}_1R_t$ rises simultaneously by the rise in ${}_1r_1$, even if ${}_1r_2 \cdots {}_1r_t$ all remain unchanged. The rise in ${}_1R_t$ is, however, smaller than the rise in ${}_1r_1$. This rise in ${}_1R_t$ occurred not because of a rise in the long term inflationary expectations, but because of a rise only in the short-term inflationary expectation.

Thus, we have explained the observed facts completely; why both the long-term and the

FIGURE 1–6



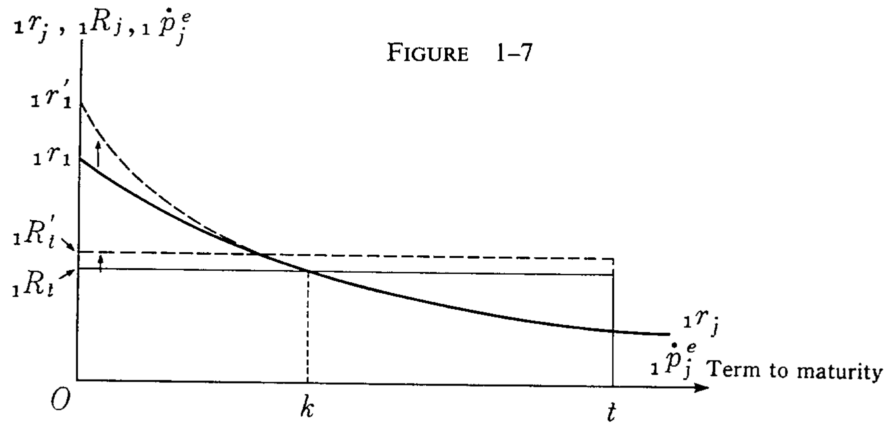


FIGURE 1–7

short-term rate are affected by past price changes with identical time lags and why the effect is larger on the short-term rate than on the long-term rate. The advantage of our analysis is that the whole effect of price changes on both the long- and the short-term rate can be explained not by the unobservable inflationary expectations but by the theory of the term structure of interest rates. Hence, it is obviously important and fruitful to consider the Fisher effect in connection with the term structure of interest rates. In the next section we will focus on this relationship directly.

Section 2. The Structure of Interest Rates

2–1. The Term Structure and the Inflationary Expectations

The importance of the divergence between nominal and real rates is commonly recognized in almost all relevant areas of economic theory. Strange as it is, in the analysis of the term structure of interest rates itself, no explicit attention has been paid so far to the distinction between nominal and real rates. According to economic theory, changes in "real" rates should reflect both shifts in the equilibrium relationship between real saving and investment and current capital market disequilibrium. Thus, it is such "real" rate series that should be employed in studies of the term structure of interest rates.

The discussion in the previous section implies that even if the yield curve of real rates is horizontal, observed yield curves could be downward or upward sloping, corresponding to price level movements in the recent past. Judged only from the observable market interest rates, we would erroneously find a positive risk premium when the rate of inflation has been falling for a considerable period and a negative risk premium when it has been rising. Therefore, in the analysis of the term structure, we have to carefully eliminate the effect of inflationary expectations in order to find the term structure of real rates. In the above analysis, the constant terms of equations (1–4) and (1–5) in Section 1 represent the fraction of interest rates which is not affected by the price level movement. If we could assume that real rates are not affected by past price changes (or inflationary expectations), the constant terms would be identified with real rates. Unfortunately, it is not appropriate to assume so, because there are famous arguments (e.g., Mundell [32] and Tobin [39]) that assert real rates fall in response to inflationary expectations, and this fact is thought to be empirical evidence for the assertion that market interest rates do not change in proportion to the actual rate of inflation. As will be shown in Chapter 2 which focuses on this problem, the real rate in Japan has been affected by inflationary expectations. Thus, we cannot find the level of the real rates based only on the constant terms of Table 1–2 and Figure 1–3.

Nevertheless, in both Tables 1–1 and 1–2 the values of the constant terms for the long term rate (α_2, α_4) are consistently higher than those for the short term rate (α_1, α_3). Also, in Figure 1–3 (the result of the regressions based on equations (1–6) and (1–7)), the constant term for the long term rate (6.82) is higher by about 30% than that for the short term rate (4.69). These findings might suggest a positive risk premium[10] in the term structure of real

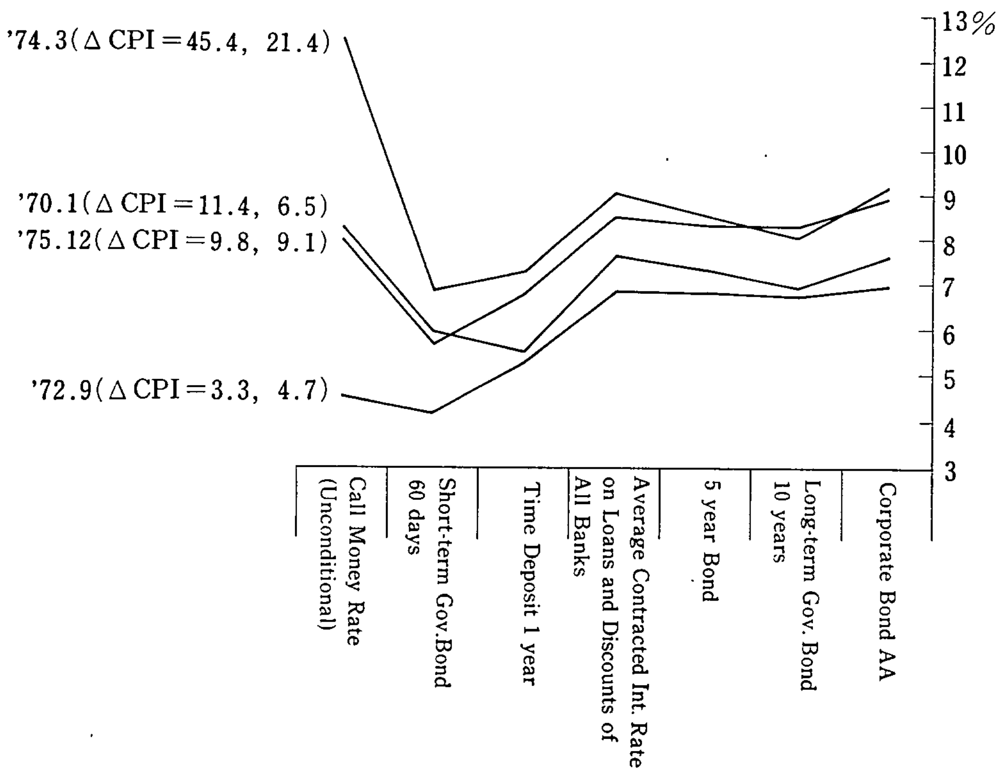
[10] A positive risk premium in the term structure of real rates is found also in the U.S. See Fama [7] and footnote (12) below.

rates. But from the analysis in the previous section only we cannot tell the exact magnitude of the risk premium, which leads us to the analysis below.

2-2. *The Empirical Analysis*

Figure 1-8 shows yield curves and the corresponding rates of inflation observed in Japan during the last decade. The first numbers inside the parentheses are the rates of inflation during the preceding 3 months (including the month in which the yeild curve is observed) shown as an annual rate and the second are the averages of the ratios of the CPI in the preceding 5 months to the corresponding month of the previous year. This figure clearly shows the tendency for yield curves to be more upward sloping the lower the rate of inflation in the preceding several months. This finding supports our hypothesis of the preceding section. Even if the real rate is affected by inflationary expectations, it should be possible to measure the difference between the long-term and the short-term interest rates, i.e., the risk premium, since the both rates are affected by inflationary expectations with identical time lags. In order to clarify the relationship between past inflation and the risk premium, we ran regressions according to the following equations for the same period analyzed in Section 1.

FIGURE 1-8 YIELD CURVES IN JAPAN



$$\left(\frac{i_t^S - i_t^L}{i_t^L}\right)\% = \alpha + \beta \dot{p}_{t-i} + \lambda_t \qquad (1\text{–}10)$$

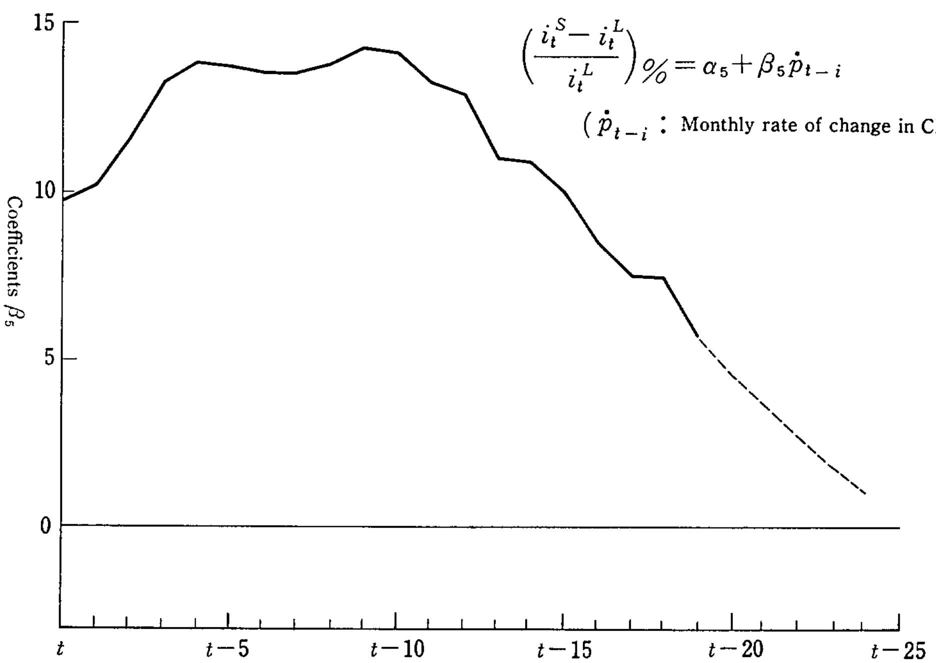
$$\left(\frac{i_t^S - i_t^L}{i_t^L}\right)\% = c_0 + c_1 P_t + c_2 P_{t-1} + \cdots c_{18} P_{t-17} + \lambda_t' \qquad (1\text{–}11)$$

Results

(1): Figure 1–9 illustrates the series of estimated coefficients for equation (1–10) based on the ratio of month-to-month changes in the CPI. This figure clearly shows that the difference between the short-term and the long-term rate is also affected by past inflation with a lag pattern almost identical to the effect on the interest rates themselves. The length of the lags with significant coefficients is 19 months which is almost the same as for the short-term rate.

(2): Table 1–3 shows the results of the regressions according to the same equation (1–10) but based on the annual rate of inflation for each month. Again in this case, the risk premium is most strongly affected by the annual rate of inflation 1–4 months earlier. Though the coefficients (β_6), t-values, $\bar{R}^2$s all decline gradually for longer lags, the rate of change in the CPI in each month relative to the corresponding month of the previous year explains more than half of the total variation in the risk premium for lags less than 7 months. Since the coefficients for lags more than 16 months are not significant at the 5% level, the effect of the annual rate of change in the CPI on the risk premium fades away in 15 months. The value of the constant term (α_6) is a measure of the risk premium which is not affected by past

FIGURE 1–9



Solid line denotes the singificance at the 5% level.

TABLE 1–3

$$\left(\frac{i_t{}^S - i_t{}^L}{i_t{}^L}\right)\% = \alpha_6 + \beta_6 \dot{p}_{t-i}$$

($\dot{P}_{t-i}$: Rate of change in CPI from corresponding month in preceeding year)

	β_6	t-value	$\bar{R}^2$	α_6
t	2.28*	13.94	0.62	−25.48
$t-1$	2.32*	14.73	0.64	−25.75
$t-2$	2.24*	15.04	0.65	−25.90
$t-3$	2.33*	14.93	0.65	−25.80
$t-4$	2.31*	14.53	0.63	−25.59
$t-5$	2.23*	13.26	0.59	−24.90
$t-6$	2.14*	12.13	0.55	−24.12
$t-7$	2.03*	10.87	0.59	−23.13
$t-8$	1.89*	9.52	0.43	−21.90
$t-9$	1.70*	7.60	0.32	−19.26
$t-10$	1.57*	7.06	0.29	−19.03
$t-11$	1.39*	5.94	0.22	−17.41
$t-12$	1.20*	4.93	0.16	−15.75
$t-13$	1.01*	4.03	0.11	−14.15
$t-14$	0.82*	3.17	0.07	−12.46
$t-15$	0.63*	2.40	0.04	−10.90
$t-16$	0.47	1.73	0.02	−9.53
$t-17$	0.37	1.29	0.01	−8.60
$t-18$	0.24	0.80	0.00	−7.55

* indicates the significance at the 5% level.

inflation and ranges around −25% for the most recent several months which have high $\bar{R}^2$s. This means that the long-term rate is about 25% higher than the short-term rate without the effect of inflation.

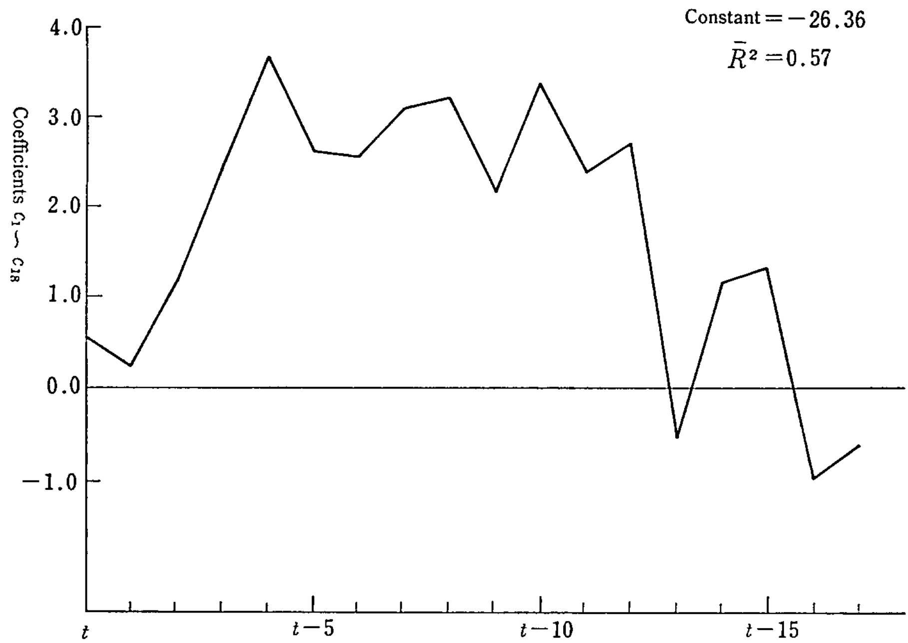
(3): Figure 1–10 shows the results of the estimation of equation (1–11) based on the monthly change in the CPI. Though we have the problem of multicollinearity in this case, the estimated coefficients are consistent with the results of the Analyses (1) and (2). That is, past rates of inflation within a year have a dominant effect on the risk premium. Here, we also obtain the value of −26% as the constant term for equation (1–11) which is perfectly consistent with the results of Anaysis (2).

From the both Analyses (2) and (3), we have obtained a positive risk premium of 25~26% of the long-term rate in the term structure of real interest rates. This result is also consistent with Analyses (2) and (3) of Section 1. Table 1–2 shows the differences between the short-term and the long-term rate of about 25~30% for lags with high $\bar{R}^2$s and in Figure 1–3 we have estimated constant terms of 6.82 for the long-term rate and 4.69 for the short-term rate, which indicates the existence of a positive risk premium of about 30% of the long-term rate. Since the difference between the long-term and the short-term rate which is estimated by a

FIGURE 1-10

$$\left(\frac{i_t^S - i_t^L}{i_t^L}\right)\% = c_0 + c_1\dot{p}_t + c_2\dot{p}_{t-1} + \cdots\cdots + c_{18}\dot{p}_{t-17}$$

($\dot{p}_{t-i}$: Monthly rate of change in CPI)



different method and using different data shows an almost identical value, we should conclude that a positive risk premium of 25∼30% of the long-term rate exists in the term structure of interest rates without the effect of inflationary expectations.

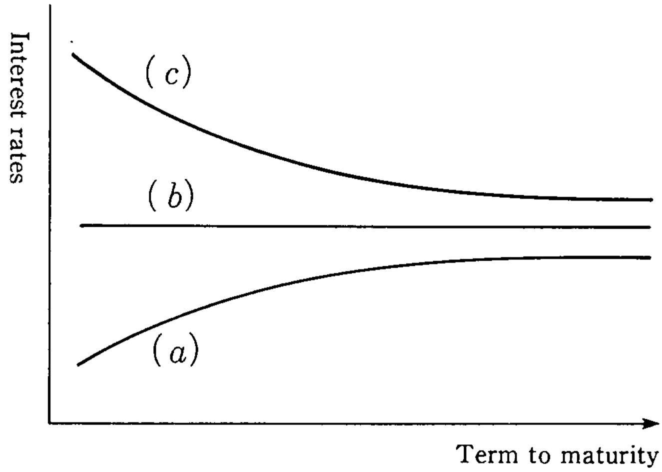
2-3. *Trade Cycle and Inflationary Expectations*

In the foregoing analysis, we have found a close relationship between the term structure and inflationary expectations. Another point that should be examined in this chapter is the advantage or explanatory power of our hypothesis relative to the conventional explanation of the term structure.

According to the conventional explanation of the term structure, the shapes of the yield curve are explained in connection with trade cycle.[11] In Figure 1-11 the yield curve (a) is said to be observed during recessions, (c) during booms and (b) during interim periods. Suppose, for example, that during a boom the level of interest rates is already high and thus people expect it to fall in the near future. In this situation lenders will shift their funds to longer term markets, since the longer the terms of lending the greater their expected capital

[11] e. g. Goode & Birnbaum [21], Smith [38]

FIGURE 1–11



gain. On the other hand, borrowers who also expect a future fall in the interest rate try to borrow short and the demand for funds shifts to shorter-term markets. As a result of these shifts of demand and supply, the short-term interest rate rises relative to the long-term rate, which leads to the yield curve of type (c) in Figure 1–11. During a period of recession an opposite movement leads the yield curve to take a shape like (a). In this manner, the conventional explanation of the cyclical movements of yield curves rests on changes in demand-supply conditions which reflect only changes in real factors in the long term and the short term bond markets. This conventional explanation presupposes an expectation of a rise or fall in the interest rate which is formed based on the divergence of the current interest rate from the level that is regarded as normal. For the interest rate to be expected to rise or fall in the near future, the current level must already be lower or higher than the normal level. According to this theory we can certainly explain the shape of the yield curve at peaks and troughts of the trade cycle, but we cannot explain the reason why the yield curve shifts from (a) to (c) during the process of recovery or from (c) to (a) during the process of contraction.

Suppose, for example, at a trough the business situation just turned to a recovery after a process of contraction. During the preceding period of contraction the level of the interest rate has already been low and people expected it to rise in the near future (yield curve (a)). In order to explain the shifts of yield curve from (a) to (c), we have to suppose instead that people revise their expectations at this point in time and expect the interest rate, which is already low, to fall further during the coming period of recovery. Otherwise, we cannot explain this shift of the yield curve from (a) to (c). But this kind of assumption hardly seems to be realistic.

In contrast, our hypothesis can explain the process of the shift of the yield curve consistently because the rate of inflation tends to move pro-cyclically. Suppose that at the beginning of the recovery the actual rate of inflation also begins to rise and thus people expect the rate of inflation to increase in the near future. According to our hypothesis, this increase in the expected rate of inflation is sufficient to explain the rise of the short-term rate relative to the long-term rate, which causes the shift of yield curve from (a) to (c). Hence, our hy-

FIGURE 1-12 (A) YIELD CURVES FOR THE U.S. GOV. BONDS

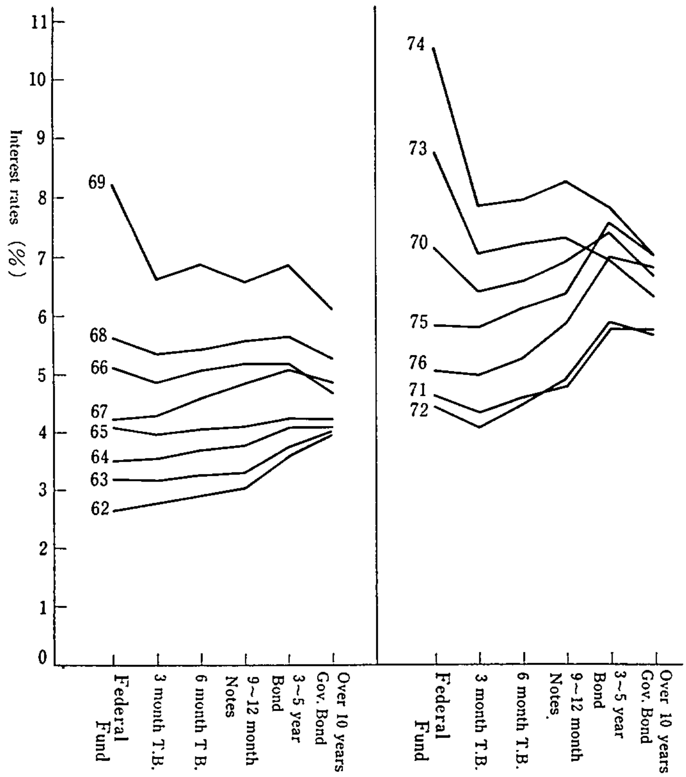
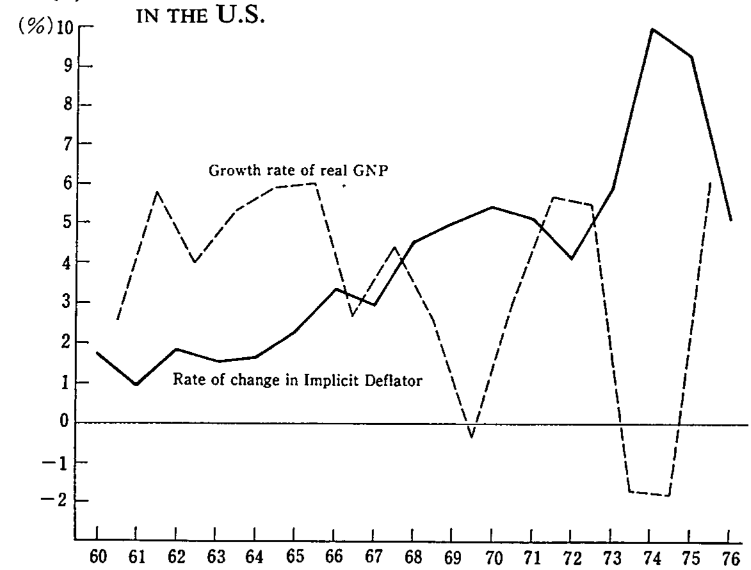


FIGURE 1-12 (B) ANNUAL RATE OF INFLATION AND RATE OF GROWTH IN REAL GNP IN THE U.S.



pothesis can explain not only the shape of the yield curve at peaks or troughs of the trade cycle but also its cyclical movement, while the conventional theory which does not take explicit account of inflationary expectations cannot fully explain the cyclical movement of the yield curve.

Further, it should be useful to compare the empirical validity of our hypothesis based on inflationary expectations and the conventional explanation based on trade cycle. For this purpose, Figure 1–12 (A) shows the yield curves for U.S. government bonds with various maturities observed during the last 15 years. Figure 1–12 (B) shows the annual rate of iflation and the rate of growth in real U.S. GNP for the corresponding period. We take the growth rate of real GNP as a representative index that shows the movement of business conditions. Though there could be other proxies, their behaviors are almost identical to that of real GNP. It is evident from these two Figure 1–12 (A) and (B) that the shapes and the movements of yield curves are more closely related with the rate of inflation than with the growth rate of real GNP. Hence, the term structure of interest rates should not be explained in connection only with the trade cycle. The expected rate of inflation is a much more important factor in explaining the term structure.

Section 3. Summary

In this chapter, we examined the Fisher effect in Japan and its relation to the term structure of interest rates. In Section 1 we analysed the effect of past inflation on long-term and short-term interest rates using the data from the last decade to obtain the following results:

a: Both the long-term and the short-term rate are afiected most strongly by the monthly rate of inflation of 3–12 months ago, but the short-term rate is 2 or 3 times more sensitive than the long-term rate.

b: Both the long-term rate and the short-term rate are affected dominantly by the past inflation during the preceding 12 months and the effect of past inflation fades away within two years.

c: The longer the horizon of expectations, the heavier the relative weights of the experience in the more distant past.

d: Inflationary expectations in Japan are not formed with distributed lags whose weights decline monotonically with the length of the lag. The weights placed on the most recent experience are relatively light. Then up to 11 months the weights become heavier the longer the lags, with the heaviest weights for lags of 10∼11 months, and they rapidly decline for lags longer than 12 months. Hence, it might be necessary to construct a hypthesis on the formation of expectations, more appropriate to Japan, taking this pattern of distributed lags into account.

Further in this chapter, we have shown that the above results of our analysis can be explained consistently by combining the inflationary expectations and the expectations hypothesis on the term structure. This argument clearly shows the importance of the Fisher effect in the analysis of the term structure. This led us to the ahalysis in Section 2 that focused on the relationships between the term structure and inflationary expectations to yield the following results:

e: The difference between the short-term and the long-term rate, i.e., the risk premium is affected by past inflation with a lag pattern almost identical to the effect on both interest rates themselves.

f: In the term structure of real interest rates without the effect of inflation, there exists a positive risk premium of about 25~30% of the long term rate.

In this section, we also pointed out that our hypothesis can explain the changes in the yield curve better than the conventional theory that explains these changes in connection with trade cycle. We also showed both the theoretical and the empirical superiority of our hypothesis to the conventional theory in explaining the changes in the yield curve, thus demonstrating the importance of inflationary expectations in the analysis of the term structure of interest rates.

Chapter 2. Inflation and The Real Interest Rate

Section 1. Previous Studies

1–1. Inflationary Expectations and the Real Interest Rate

In the previous chapter, we examined the relationship between inflation and market interest rates but did not directly test the validity of the Fisher equation (1–1). That is, we have not yet examined to what degree the nominal rate reflects inflationary expectations. It might be suspect whether the exact relationship is represented by equation (1–1) and the relationship between the real rate and inflation is as yet not clear. In this chapter we examine the validity of the Fisher effect by considering the cause of the difference between the empirical results for Japan and the U.S.

The world-wide inflation that revived interest in the Fisher effect at the same time aroused suspicions about the effect. During the period of rapid inflation in the early 70s the rate of increase in nominal rates was much smaller than that of actual inflation, which might imply negative real interest rates. In order to defend the Fisher effect, we have to interpret this situation as being caused by either imperfect expectations or a fall in real rates, or both. There are many theoretical analyses of this problem. Mundell [32] asserted that the real rate falls under inflationary expectations due to the existence of the wealth-saving relationship and Tobin [39] drew the same conclusion based on the argument that inflationary expectations raise capital intensity and lead to a fall in the real rate of return. Recently Feldstein [8], Gandolfi [17] and others pointed out the importance of the tax burden on this relationship.

On the empirical side, as noted already, there are numerous analyses based on various assumptions on the formation of expectation. In order to examine the validity of the Fisher effect without any specific assumption on the formation of expectation, W. Gibson [20] adopted a unique approach that utilized direct data on expected inflation. Since this analysis is closely related with our analysis in the previous chapter, we first examine his study.

1–2. Gibson's Study

In testing the validity of the Fisher effect, W. Gibson [20] did not presuppose the impossibility of measuring inflationary expectations. Instead, he used data on expectations obtained by direct surveys of important market participants. The data come from Joseph Livingston's survey of a group of business, government, labor, and academic econimists concerning their expectations of future values of selected aggregate economic variables which included the consumer price index, six and twelve-months hence. Using this data he measured the effect of price expectations on interest rates in the U.S. by estimating the following equation:

$$i = a_0 + a_1 \dot{p}^* \qquad (2\text{–}1)$$

where $\dot{p}^*$ is the expected rate of inflation obtained by a direct survey. Theoretically a_1 is expected to be between zero and one. $a_1=1$ is consistent with a situation in which the real rate of interest is unaffected by inflationary expectations and nominal interest rates fully adjust to these expectations. It is, however, also consistent with a situation in which positive (negative) effects on the real rate are exactly matched by underadjustment (overadjustment) of nominal rates. Similarly $a_1=0$ is consistent with an unchanged real rate to expectations as well as with other combinations of real and nominal rate adjustment. Though it is certainly possible that the latter outcomes might exist, there are no theoretical arguments that such relationship should always hold. There is, he said, as yet no theoretical consensus on the relationship between the real rate and the expected rate of inflation. As a result, he assumed in his analysis that variations in price expectations do not affect the real rate. Thus the estimated coefficient a_1 shows directly the degree of the response of nominal rates to inflationary expectations. Another point that should be noted is that the measures used here should affect interest rates differently depending on the term to maturity of the security. Since the data used here for $\dot{p}^*$ are price expectations for the coming six and twelve months, the influence of these expectations should diminish as terms to maturity increases beyond six and twelve months.

The result of Gibson's estimation based on these assumptions and using the yields on U.S. Treasury securities is shown in Table 2–1. Five different maturity categories were used ranging from 3-month bills to 10-year and longer-term to maturity bonds and the period analyzed is 1952–1970. The table shows that over half the coefficients are quite close to 1.0 and that all coefficients are significant at the 5% level or better. The 6- and 12-month expected rates of inflation have their largest effects on the 6-month and 9-12 month bill rates and the coefficients range from 0.911 to 1.096, which means perfect adjustment of nominal rates and no change in the real rate. The coefficients decline with term to maturity as hypothesized. The nominal interest rate on 10-year and longer bonds is increased by 0.450 of the change in the rate expected for the following 6 months and by 0.675 of the rate expected for the following 12 months. For the period 1959–1970, the coefficients are even closer to 1.0 as shown in Table 2–2. Since this estimation is based on the data with 6 month interval, the adjustment period of nominal rates to the inflationary expectations could be considered to be less than 6 months. This finding is perfectly consistent with our result in the preceding chapter[12]. Gibson himself

[12] Though Gibson himself did not mention this aspect, the values of the constant terms in Table 2–1 and 2–2 are worthy of note in connection with our preceding analysis. Since the real interest rates are assumed to be constant in this case, the constant terms can be identified with the real rates. These values are the higher the longer the term to maturity, which is perfectly consistent with our finding of positive risk premiums. Moreover, the magnitudes of the risk premiums are about 30% of the long-term rates which is very close to our result.

TABLE 2–1 GIBSON'S RESULT ON INTERTST RATES AND INFLATIONARY EXPECTATIONS

	Constant	$\dot{P}^*_{t+n}$	R^2	$S.E.$
$n=6$				
3-month bills	2.6643[a] (0.177)	0.6616[a] (0.0764)	0.667	0.9535
6-month bills	2.359[a] (0.293)	0.9358[a] (0.1140)	0.751	0.7578
9–12-month bills	2.496[a] (0.273)	0.911C[a] (0.1062)	0.767	0.7061
3–5-year notes	3.371[a] (0.164)	0.6113[a] (0.0707)	0.666	0.8822
10-year and longer bonds	3.58C[a] (0.125)	0.4503[a] (0.0540)	0.649	0.6742
$n=12$				
3-month bills	2.207[a] (0.170)	0.9300[a] (0.0854)	0.761	0.8076
6-motn bills	2.045[a] (0.304)	1.0958[a] (0.1236)	0.779	0.7174
9–12-month bills	2.192[a] (0.282)	1.06548[a] (0.1149)	0.794	0.6637
3–5-year notes	2.921[a] (0.133)	0.8959 (0.0668)	0.829	0.6317
10-year and longer bonds	3.23C[a] (0.094)	0.675C[a] (0.0472)	0.847	0.4460

a: Singificant at the 5% level.
(*Source:* W.E. Gibson [20] p. 856)

noted that his result "is consistent with other findings (Gibson, Sargent, Yohe and Karnosky) which suggest that long-term expectations are based heavily (but not to the same extent) on the same factors determining shorter term expectations". This reinforces our hypothesis in the preceding chapter. Based on these results, consequently, Gibson concluded that nominal rates fully adjust to inflationary expectations within 6 month, that the real rate is not affected by the change in the inflationary expectations, and that the Fisher equation holds almost perfectly in the U.S. for the period 1952–1970.

1–3. *Fama's Study*

Now, in Japan how does the real rate behave and how valid is the Fisher equation? One way to examine these problems is of course to follow Gibson's approach. Unfortunately, however, we cannot obtain such direct survey data on price expectations as used by Gibson. The alternative method we have adopted is the approach used by Fama [5].

In a world of uncertainty, the Fisher equation can be thought to assert that the nominal rate is equal to the equilibrium expected real return plus the market's assessment of the expected rate of inflation. Fama's analysis is based on the theory of efficient markets and uses data on U.S. treasury bills. His conclusion is that expected real returns on treasury bills seem to be constant during 1953–1971 and that the bill market seems to be efficient in the sense that nominal interest rates summarize all the information about future inflation rates

TABLE 2–2 GIBSON'S RESULT ON INTEREST RATES AND INFLATIONARY EXPECTATIONS

	Constant	$\dot{P}^*_{t+n}$	R^2	*S.E.*
$n=6$				
3-month bills	2.204[a] (0.302)	0.9266[a] (0.1172)	0.737	0.7790
6-month bills	2.359[a] (0.293)	0.9358[a] (0.1140)	0.751	0.7578
9–12-month bills	2.496[a] (0.273)	0.911C[a] (0.1062)	0.767	0.7061
3–5-year securities	3.087[a] (0.238)	0.8312[a] (0.0924)	0.784	0.6140
10-year and longer bonds	3.437[a] (0.177)	0.6012[a] (0.0689)	0.774	0.4579
$n=12$				
3-month bills	1.889[a] (0.312)	1.0869[a] (0.1269)	0.767	0.7330
6-month bills	2.045[a] (0.304)	1.0957 (0.1237)	0.779	0.7141
9–12-mothn bills	2.192[a] (0.282)	1.0658[a] (0.1149)	0.794	0.6637
3–5-year securities	2.771[a] (0.222)	0.9903[a] (0.0905)	0.844	0.5226
10-year and longer bounds	3.170[a] (0.144)	0.7342[a] (0.5847)	0.877	0.3377

a: Singificant at the 5% level.
(*Soruce*[7] W.E. Gibson [20] p. 858)

that is in time-series of past inflation rates. This conclusion is, as will be shown below, based on the fact that the autocorrelations among real returns on treasury bills are near zero and that substantial variation in nominal bill rates seems to be due entirely to variation in inflationary expectations.

In the following section, we analyze the behavior of the real rate in Japan according to Fama's approach and examine the validity of the Fisher equation.

Section 2 The Analytical Framework

2-1. The Theory of Efficient Markets

Here we just restate Fama's analytical framework on which our analysis in the next section is based. First we define the variables used below. i_t is the nominal return on a security from the end of period $t-1$ to the end of period t, that is $i_t=(v_t-v_{t-1})/v_{t-1}$ where v_t, v_{t-1} are the prices of the security at t and $t-1$ respectively. For a security with a known v and one period to maturity at $t-1$, once v_{t-1} is set, i_t is known and can be identified with the one-period nominal rate of interest set in the market at $t-1$ and realized at t. Let P_t be the price level at t, then the price of money in terms of goods is shown as $\pi_t=1/P_t$.
The real return from $t-1$ to t on a one-period security is

$$\tilde{r}_t = (v_t\tilde{\pi}_t - v_{t-1}\pi_{t-1})/v_{t-1}\pi_{t-1} \tag{2-3}$$

$$= i_t + \tilde{\Delta}_t + i_t\tilde{\Delta}_t \tag{2-4}$$

where tildes(~) indicate random variables and $\tilde{\Delta}_t = (\pi_t - \pi_{t-1})/\pi_{t-1}$ (2–5) is the rate of change in purchasing power from $t-1$ to t. Since in monthly data i_t and Δ_t are close to zero, we can safely use

$$\tilde{r}_t = i_t + \tilde{\Delta}_t \tag{2-6}$$

as an approximation. Thus, the real return from the end of period $t-1$ to the end of period t on a security with one period to maturity at $t-1$ is the nominal return plus the rate of change in purshasing power from $t-1$ to t.

In the following analysis the word "market efficiency" means that in setting the price of a one-period security at $t-1$, the market correctly uses all available information to assess the distribution of $\tilde{\Delta}_t$. Fomally, in an efficient market,

$$f_m(\Delta_t|\phi_{t-1}{}^m) = f(\Delta_t|\phi_{t-1}) \tag{2-7}$$

where ϕ_{t-1} is the set of information avaialble at $t-1$, $\phi_{t-1}{}^m$ is the set of information used by the market, $f_m(\Delta_t|\phi_{t-1}{}^m)$ is the market-assessed density function for $\tilde{\Delta}_t$, and $f(\Delta_t|\phi_{t-1})$ is the true density function implied by ϕ_{t-1}.

When the market sets the equilibrium price of a one-period security at $t-1$, it is also set. Given the relationship among $\tilde{r}_t$, i_t and $\tilde{\Delta}_t$ in (2–6), the market's assessed distribution for r_t is implied by i_t and its assessed distribution for $\tilde{\Delta}_t$. If (2–7) holds, then the market's assessed distribution for $\tilde{r}_t$ is the true distribution

$$f_m(r_t|\phi_{t-1}{}^m, i_t) = f(r_t|\phi_{t-1}, i_t) \tag{2-8}$$

In short, if the market is efficient, then in setting the nominal price of a one-period security at $t-1$, it correctly uses all available information to assess the distribution of $\tilde{\Delta}_t$. In this sense v_{t-1} fully reflects all available information about $\tilde{\Delta}_t$. Since an equilibrium value of v_{t-1} implies an equilibrium value of i_t, the one-period nominal rate of interest set in the market at $t-1$ likewise fully reflects all available information about $\tilde{\Delta}_t$. Finally, when an efficient market sets i_t, the distribution of the real return $\tilde{r}_t$ perceives is the true distribution.

2–2. *A Model of Market Equilibrium*

Since neither $f_m(\Delta_t|\phi_{t-1}{}^m)$ nor $f(\Delta_t|\phi_{t-1})$ are directly observable we cannot know whether they are equal or not, and thus we cannot evaluate the efficiency of the market. In empirical analysis, therefore, it is necessary to specify the relationship between $f_m(\Delta_t|\phi_{t-1}{}^m)$ and v_{t-1} in more detail. For this purpose Fama presents two alternative models of market efficiency. The common assumption of these models is that the primary concern of investors is the distribution of the real return on a security.

(1): Analysis of the Autocorrelations of r_t

Fama's first model supposes the following relationship about the characteristics of the market assessed distribution $f_m(r_t|\phi_{t-1}{}^m, i_t)$ that results from an equilibrium price v_{t-1} at $t-1$.

$$E_m(\tilde{r}_t|\phi_{t-1}{}^m, i_t) = E(\tilde{r}) \tag{2-9}$$

In words, equation 2–9 means the equilibrium expected real return on a one-period security is assumed to be constant through time. In an efficient market

$$E_m(\tilde{r}_t|\phi_{t-1}{}^m, i_t) = E(\tilde{r}_t|\phi_{t-1}, i_t) \tag{2-10}$$

holds and (2–9) and (2–10) jointly imply

$$E(\tilde{r}_t|\phi_{t-1}, i_t) = E(\tilde{r}) \tag{2-11}$$

Thus at any time the market sets the price of a one-period security so that its assessment of the expected real return is the constant $E(\tilde{r})$. Since an efficient market correctly uses all available information, $E(\tilde{r})$ is also the true expected real return on the security.

On these assumptions we can test that there is no way to use ϕ_{t-1}, the set of information available at $t-1$, or any subset of ϕ_{t-1}, as the basis of a correct assessment of the expected real return on a one-period security which is other than $E(\tilde{r})$. One subset of ϕ_{t-1} is the time-series of past real returns. If (2–11) holds,

$$E(\tilde{r}_t|r_{t-1}, r_{t-2}, \cdots) = E(\tilde{r}) \tag{2-12}$$

that is, there is no way to use the time-series of past real returns as the basis of a correct assessment of the expected real return which is other than $E(\tilde{r})$. If (2–12) holds, the autocorrelations of r_t for all lags are equal to zero, so that sample autocorrelations provide tests of (2–12).

But the autocorrelations are joint tests of market efficiency and of the model for the equilibrium expected real return. Thus zero autocorrelations of $\tilde{r}_t$ give support to both hypotheses of market efficiency and constant expected real return, but non-zero autocorrelations do not show which hypotheses is inappropriate. Put differently, non-zero autocorrelations are consistent with the situation where the equilibrium expected real return is constant and the market is inefficient, and also with the situation where the market is efficient and equilibrium expected real returns change over time. Therefore Fama presented an alternative test that covers this deficiency.

(2): Regression Analysis

Suppose that at any time $t-1$ the market always sets the price of a one-period security so that it perceives the expected real return to be

$$E(\tilde{r}_t|\phi_{t-1}{}^m, i_t) = \alpha_0 + \gamma i_t \tag{2-13}$$

If the market is also efficient,

$$E(\tilde{r}_t|\phi_{t-1}, i_t) = \alpha_0 + \gamma i_t \tag{2-14}$$

With (2–6), (2–13) and (2–14) imply that

$$E_m(\tilde{\Delta}_t|\phi_{t-1}{}^m) = \alpha_0 + \alpha_1 i_t \qquad \alpha_1 = \gamma - 1 \tag{2-15}$$

$$E(\tilde{\Delta}_t|\phi_{t-1}) = \alpha_0 + \alpha_1 i_t \qquad \alpha_1 = \gamma - 1 \tag{2-16}$$

In this model, γ is the proportion of the change in the nominal rate from one period to the next that reflects a change in the equilibrium expected real return, and $-\alpha_1 = 1 - \gamma$ is the proportion of the change in i_t that reflects a change in the expected value of $\tilde{\Delta}_t$.

Estimates of α_0 and α_1 in (2–16) can be obtained by applying least squares to

$$\tilde{\Delta}_t = \alpha_0 + \alpha_1 i_t + \tilde{\varepsilon}_t \tag{2–17}$$

If the coefficient estimates are inconsistent with the hypothesis that

$$\alpha_0 = E(\tilde{r}) \quad \text{and} \quad \alpha_1 = -1 \tag{2–18}$$

the model of a constant equilibrium expected real return is rejected.

In order to test of market efficiency we can use another result of (2–17). Since in an efficient market i_t summarizes all the information about the expected value of $\tilde{\Delta}_t$ which is in ϕ_{t-1}, the sequence of past values of the disturbance $\tilde{\varepsilon}_t$ in (2–17) should be of no additional help in assessing the expected value of $\tilde{\Delta}_t$ which implies that the autocorrelations of the disturbance should be zero for all lags.

Another way to examine market efficiency is to run the following regression.

$$\tilde{\Delta}_t = \alpha_0 + \alpha_1 i_t + \alpha_2 \Delta_{t-1} + \tilde{\varepsilon}_t \tag{2–19}$$

One item of information available at $t-1$ is Δ_{t-1}. If the information in Δ_{t-1} is not correctly used by the market in setting i_t, then $\alpha_2=0$. On the other hand, if the market is efficient i_t summarizes any information included in Δ_{t-1} and past values of $\tilde{\varepsilon}_t$. Thus $\alpha_2=0$ and the autocorrelations of the disturbance $\tilde{\varepsilon}_t$ in (2–19) should be zero for all lags.

Section 3. Empirical Analysis

3–1. Results of the Analysis

We analyze here the Japanese call money market based on the same monthly data used in the analysis in the previous chapter. Since seasonal factors affect the results in this case, we used three CPI series—(I) the CPI including all commodities (Cities with 50 thous. inhabitants and over), (II) the CPI excluding seasonal goods (all Japan), and (III) the seasonally adjusted CPI including all commodities (all Japan). The period analyzed is from January 1966 to March 1976.

(1): Analysis of the Autocorrelations of Δ_t

Table (2–3) shows sample autocorrelation $\hat{\rho}_\tau$ of Δ_t for lags τ of from one to twelve months. It also shows sample means $\bar{\Delta}$ and standard deviations $S(\Delta)$, and the approximate standard error of $\hat{\rho}_1$ under the hypothesis that the true autocorrelation is zero $\sigma(\hat{\rho}_1)$[13]. The market efficiency hypothesis to be tested below is that the one-period nominal interest rate i_t set in the market at the end of period $t-1$ is based on correct utilization of all the information about the expected value of $\tilde{\Delta}_t$, which is in the time-series of passt values $\Delta_{t-1}, \Delta_{t-2}, \cdots$. For this hypothesis to be meaningful past rates of change in purchasing power do indeed have information about the expected future rate of change.

[13] $\sigma(\hat{\rho}_\tau)$ is calculated according to the following equation

$$\sigma(\hat{\rho}_\tau) = \frac{1}{\sqrt{T-\tau}}$$

where T is the number of observations used to compute $\sigma(\hat{\rho}_\tau)$.

TABLE 2–3 AUTOCORRELATION OF Δ_t
(Jan. 1966~Mar. 1976)

CPI	(I) All commodities	(II) Excl. seasonal goods	(III) Seasonally adjusted
$\hat{\rho}_1$	0.27	0.57	0.48
$\hat{\rho}_2$	0.02	0.34	0.34
$\hat{\rho}_3$	0.13	0.31	0.25
$\hat{\rho}_4$	0.05	0.26	0.21
$\hat{\rho}_5$	0.12	0.40	0.25
$\hat{\rho}_6$	0.24	0.43	0.28
$\hat{\rho}_7$	0.13	0.30	0.29
$\hat{\rho}_8$	0.14	0.26	0.32
$\hat{\rho}_9$	0.24	0.34	0.40
$\hat{\rho}_{10}$	0.05	0.29	0.34
$\hat{\rho}_{11}$	0.19	0.27	0.28
$\hat{\rho}_{12}$	0.27	0.27	−0.08
$\sigma(\hat{\rho}_1)$	0.09	0.09	0.09
$\bar{\Delta}$	−0.00697	−0.00667	−0.00688
$s(\Delta)$	0.00974	0.00686	0.00749
T	123	123	123

The table indicates positive autocorrelations for all three CPI series, but their magnitudes differ depending on the particular CPI series. Though 3/4 of the sample autocorrelations of CPI(I) are larger than $\sigma(\hat{\rho}_1)$, most of them are less than 0.2. On the other hand, the sample autocorrelations of CPI (II) and CPI(III) are much larger than those of CPI(I) for almost all lags, which means that these series of past CPIs do include more useful information about the expected future rate of change than CPI(I) does.

(2): Analysis of the Autocorrelations of r_t

Table 2–4 shows the same statistics as Table 2–3 for the real rates which are computed using the three CPI series. If the equilibrium expected real return is constant over time and if the market is efficient at the same time, the autocorrelations of r_t should be zero for all lags. Though the autocorrelations in column (I) are relatively low, this might be interpreted as the result of the small autocorrelations of CPI(I). In columns (II) and (III), however, positive autocorrelations are observed for all lags, with magnitudes of about 0.2~0.4. Though the autocorrelations of r_t are a little smaller than those of Δ_t, still remaining substantial positive autocorrelations of r_t mean that i_t does not fully summarize the available information which is in the time series of past values, Δ_{t-1}, Δ_{t-2}, $\cdots$. This means in turn that the market is inefficient and does not use the available information completely, or that the equilibrium expected real returns are not constant, or both. Thus, we have to turn to Tables 2–5 and 2–6 in order to identify the cause of the substantial autocorrelations of r_t.

(3): Market Efficiency

Tables 2–5 and 2–6 show the results of the regressions estimated according to the equation

TABLE 2–4 AUTOCORRELATION OF r_t
(Jan. 1966～Mar. 1976)

CPI	(I) All commodities	(II) Excl. seasinal goods	(III) seasonally adjusted
$\hat{\rho}_1$	0.21	0.48	0.41
$\hat{\rho}_2$	−0.06	0.21	0.24
$\hat{\rho}_3$	0.07	0.15	0.15
$\hat{\rho}_4$	−0.02	0.12	0.11
$\hat{\rho}_5$	0.06	0.31	0.16
$\hat{\rho}_6$	0.19	0.37	0.19
$\hat{\rho}_7$	0.08	0.22	0.21
$\hat{\rho}_8$	0.10	0.19	0.26
$\hat{\rho}_9$	0.20	0.30	0.34
$\hat{\rho}_{10}$	−0.00	0.24	0.29
$\hat{\rho}_{11}$	0.16	0.27	0.22
$\hat{\rho}_{12}$	0.23	0.27	−0.18
$\sigma(\hat{\rho}_1)$	0.09	0.09	0.09
$\bar{r}$	−0.00049	−0.00042	−0.00022
$s(r)$	0.00907	0.00701	0.00633
$s(\bar{r})$	0.00082	0.00063	0.00057
T	123	123	123

TABLE 2–5 REGRESSION ANALYSIS
$\Delta_t = a_0 + a_1 i_t + e_t$
(Jan. 1966～Mar. 1976)

CPI	(I) All commodities	(II) Ecl. seasonal goods	(III) Seasonally adjusted
a_0	0.00259	0.00241	0.00213
a_1	−1.48	−1.41	−1.39
$s(a_0)$	0.00292	0.00197	0.00219
$s(a_1)$	0.43	0.29	0.32
Coefficient of determination	0.08	0.15	0.13
$s(e)$	0.00938	0.00633	0.00703
$\hat{\rho}_1(e)$	0.21	0.48	0.40
$\hat{\rho}_2(e)$	−0.05	0.21	0.23
$\hat{\rho}_3(e)$	0.07	0.17	0.14
$\hat{\rho}_4(e)$	−0.02	0.12	0.09

TABLE 2–6 REGRESSION ANALYSIS

$\Delta_t = a_0 + a_1 i_t + a_2 \Delta_{t-1} + e_t$

(Jan. 1966–Mar. 1976)

CPI	(I) All commodities	(II) Excl. seasonal goods	(III) Seasonally adjusted
a_0	0.00205	0.00087	0.00067
a_1	−1.18	−0.66	−0.74
a_2	0.20	0.49	0.40
$s(a_0)$	0.00288	0.00176	0.00205
$s(a_1)$	0.45	0.29	0.33
$s(a_2)$	0.09	0.08	0.09
Coefficient of determination	0.11	0.34	0.25
$s(e)$	0.00923	0.00559	0.00650
$\hat{\rho}_1(e)$	0.04	0.02	−0.01
$\hat{\rho}_2(e)$	−0.11	−0.05	0.07
$\hat{\rho}_3(e)$	0.09	0.07	0.02
$\hat{\rho}_4(e)$	−0.05	−0.09	−0.01

(2–17) and (2–19). Namely, the tables show the estimated coefficients a_0, a_1, and a_2, the the standard errors $s(a_0)$, $s(a_1)$, $s(a_2)$, the coefficients of determination adjusted for the degrees of freedom, the standard deviations of the disturbance $s(e)$ and the autocorrelations of the disturbances for the first four monthly lags $\hat{\rho}_\tau(e)$.

If the market is efficient, a_2 the coefficient of Δ_{t-1} in (2–19) has to be zero, and the autocorrelations of the disturbances of both regressions also have to be zero for all lags. According to Table 2–5, however, the autocorrelations of the disturbances are large especially in columns (II) and (III). This implies that i_t does not summarize the available information completely. In Table 2–6 the estimated coefficients of $\Delta_{t-1}(a_2)$ are far greater than the standard errors from zero in all columns, and the hypothesis that $a_2=0$ is clearly rejected. Moreover, irrespective to the CPI series the coefficients of determination rise when Δ_{t-1} is included among the explanatory variables. In particular, the coefficients of determination in columns (II) and (III) of Table 2–6 are twice as large as those in Table 2–5. And in Table 2–6, the autocorrelations of the disturbances of (2–19) are near zero for all columns. This result is also consistent with the finding of especially large autocorrelations for lags of one month in Table 2–5. Now, it is clear that the market does not fully utilize the available information included in Δ_{t-1} and thus the market is not efficient. Therefore we can conclude that one reason for the large autocorrelations of real rates is market inefficiency. This result is also consistent with our finding in the previous chapter of the relatively low weights given to most recent experiences.

(4): The Real Rate of Return

We have next to examine the hypothesis of the constant equilibrium expected rate of return. If the expected value of r_t is constant over time, the constant terms a_0 in (2–17) and (2–19) represent the constant expected real return $E(r)$ and the coefficients on i_t (a_1) must be

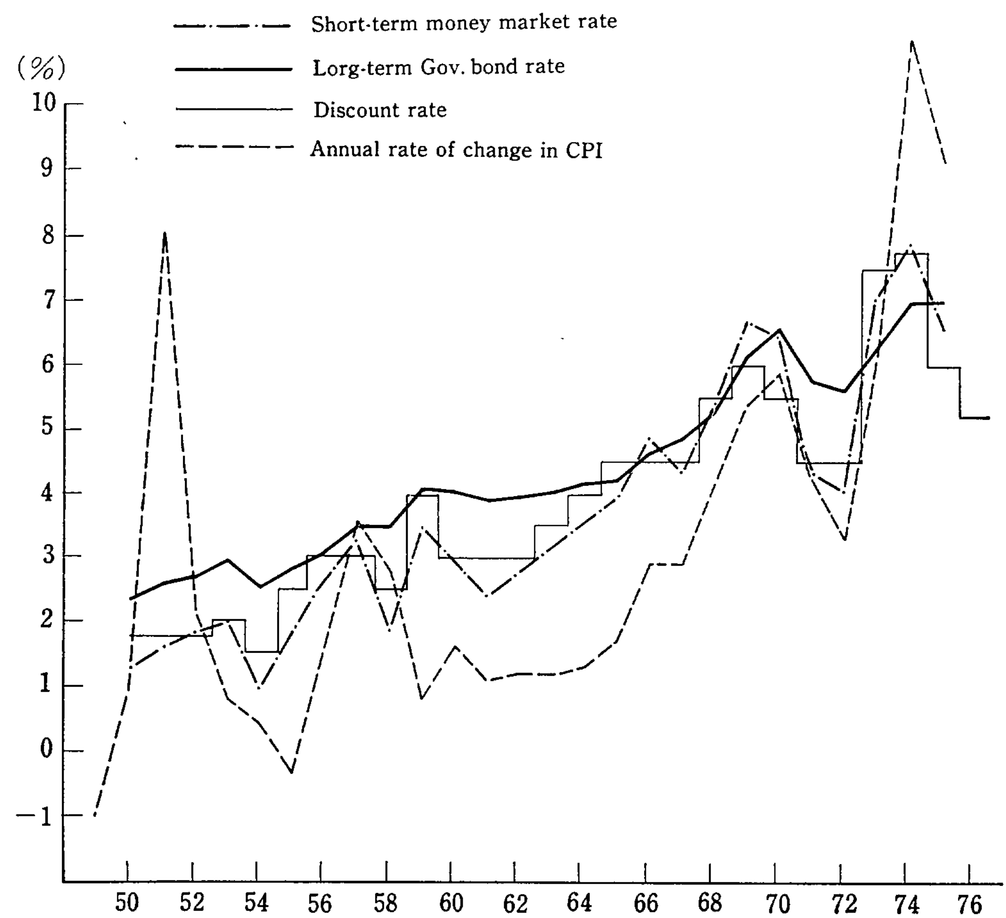
-1.0. According to Table 2–5 and 2–6, however, the estimated coefficients (a_1) are much different from -1.0 since all the estimated values of a_1, except in column (I) of Table 2–6, are more than two standard errors away from -1.0. This means that expected real returns have changed over time.

Thus in Japan we cannot say either that real interest rates were constant over time or that the market was efficient. Obviously this result is very much different from the findings of Gibson and Fama for the U.S. In order to examine the validity of the Fisher effect, then, it is necessary to explain the reasons for the different results obtained for the two countries.

3–2. The Validity of the Fisher Effect

Before we compare our result, for Japan with those for the U.S., it might be useful to see the actual movements of the price level and market interest rates in both countries. Figure

FIGURE 2–1 INFLATION AND INTEREST RATES IN THE U.S.



2-1 shows the movements of the rate of change in the price level and several interest rates in the U.S. during the last quarter century. These interest rates apparently show a steadily rising trend from the 1950's to the 70's accompanied by cyclical fluctuations. Since according to Gibson and Fama real interest rates in the U.S. can be regarded to have been constant between 1952 and 1971, this rising trend of market interest rates should be explained by the change in the expected rate of inflation. The rate of change in the CPI depicted in the same

FIGURE 2-2 ESTIMATED REAL INTEREST RATES IN THE U.S.

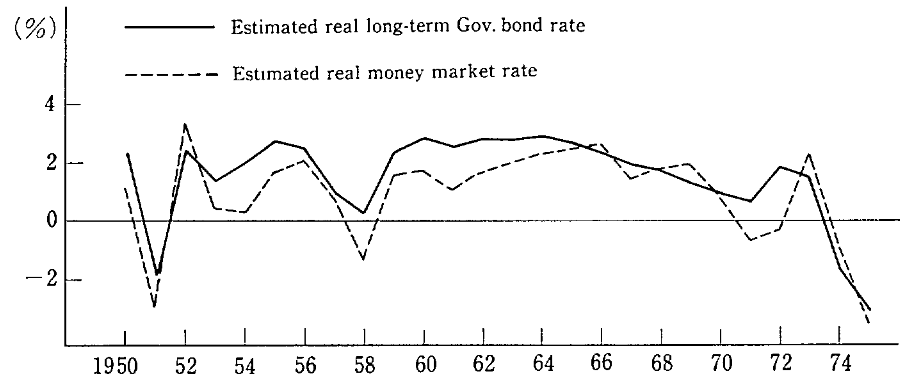


FIGURE 2-3 INFLATION AND INTEREST RATES IN JAPAN

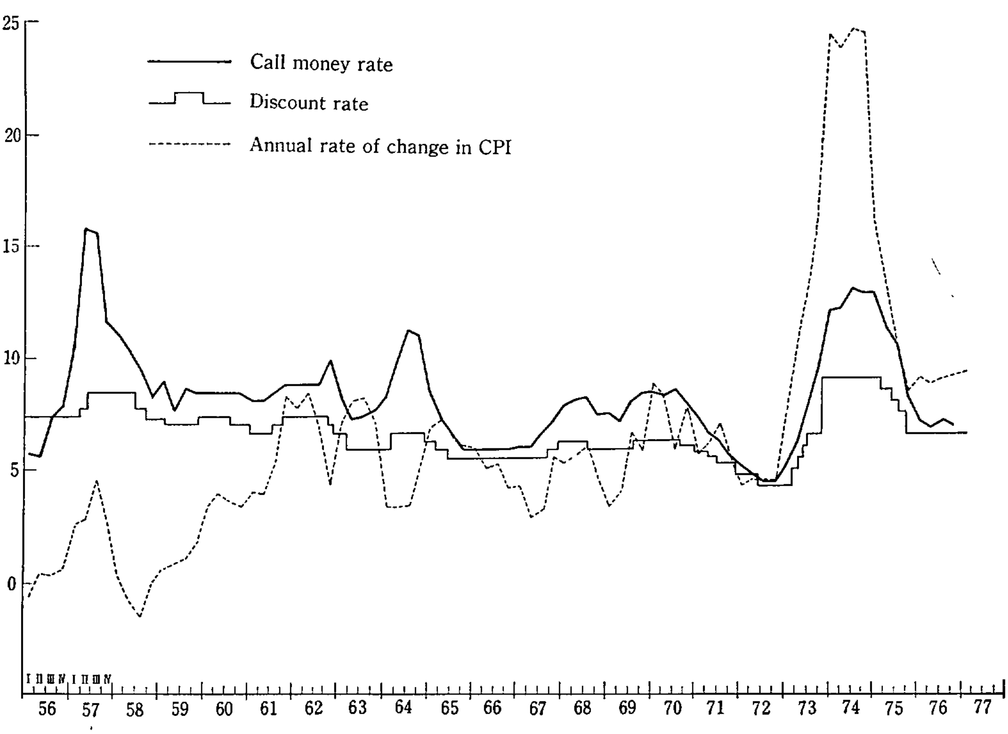
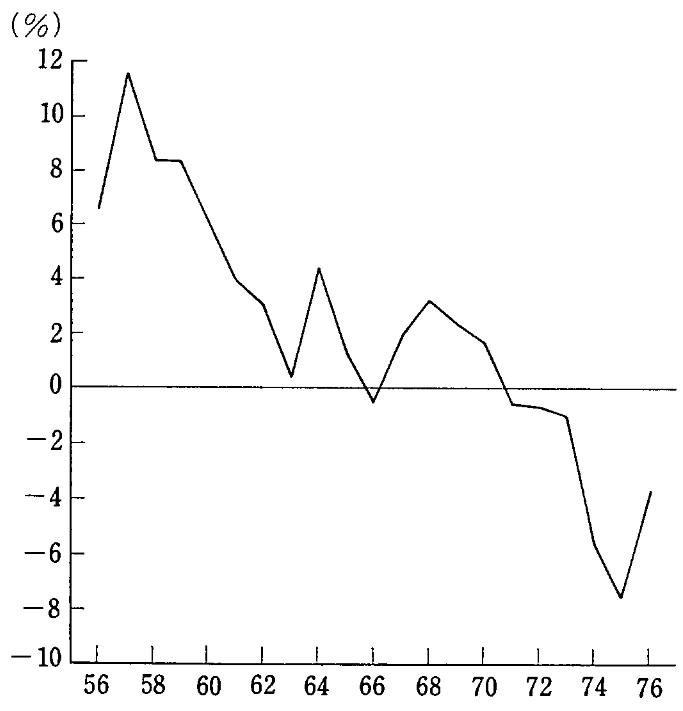


figure also shows a similar rising trend and thus gives support to the conclusion of Gibson and Fama. Though we cannot tell the exact length of the lags, this figure also gives intuitive support to the results of Yohe and Karnosky [40] and Gibson [20] that the effect of past inflation on interest rates works out within two years and that the adjustment of nominal interest rates to inflationary expectations occurs within six months.

Figure 2–2 shows the estimated real interest rates for long term U.S. government bonds and the real money market rate. These series are calculated by subtracting the moving average of the rates of inflation in the preceding two years from the average market rate in the current year. The estimated rates seem to be roughly constant during the period 1952–1970, and thus they seem to be consistent with the results of Gibson and Fama. These rates are especially stable at the level of 2~3% during the 1960s. This pattern is perfectly consistent with the fact that Gibson's estimates of a_1 were closest to -1.0 for the period 1959–1970 and the estimated constant terms ranged around 2~3%.

Now for Japan, Figure 2–3 shows the discount rate, the call money rate and the rate of change in CPI during 1956–1976. The movement of the real call money rate, which was estimated by the same method as used for Figure 2–2, is shown in Figure 2–4. Figure 2–3 indicates no clear trend in the movement of the rate of change in the CPI before 1970 and the occurrence of sudden and very rapid inflation after 1973. Since this period of sudden and very rapid inflation occupies a significant part of our analysis of interest rates in Japan, the difference in the types of inflation could be a cause of the difference between the results obtained

FIGURE 2–4 ESTIMATD REAL CALL MONEY RATE IN JAPAN



for Japan and the U.S. The unstable movement of the estimated real call money rate in Figure 2–4 is consistent with our result in the previous section that the real return in the Japanese call money market has changed over time.

In order to examine the difference between Japan and the U.S., it is best to compare our results for Japan with a study for the U.S. which also included a more inflationary period. Fama [6] extended his analysis to cover the period 1971–1974 during which the U.S. also suffered from a sudden and rapid inflation. Tables 2–7 and 2–8 contrast his results for the period 1953–1971 which we surveyed in section 1–3 and for the period 1959–1974 which includes the period of the rapid inflation. In the former period, as noted above, the market was efficient and the real rate of return was constant over time. However, according to Fama's analysis of the latter period, the same market was inefficient and the real interest rate changed over time. It can be seen in the tables that during this period the autocorrelations of r_t are larger and the coefficient a_1 cannot be identified with -1.0. Nevertheless when we contrast this result with ours (Tables 2–3 to 2–6), we see that the autocorrelation of r_t and the disturbances for the U.S. are smaller than for Japan, despite the larger autocorrelations observed in the U.S. by including the strong inflationary period. Moreover in Table 2–8, the addition of Δ_{t-1} to the explanatory variable does not affect the coefficient of determination. Therefore we can conclude that the treasury bill market in the U.S. seems to be far more efficient than the call money market in Japan. As for the real rate of return, however, we can find little difference between Japan and the U.S. only from the comparison of these results. Consequently, both in Japan and the U.S. the real interest rates seem to have changed in response

TABLE 2–7 FAMA'S RESULTS

	Autocorrelation of Δ_t		Autocorrelation of r_t	
	1/53–7/71	3/59–6/74	1/53–7/71	3/59–6/74
$\hat{\rho}_1$	0.36	0.50	0.09	0.09
$\hat{\rho}_2$	0.37	0.56	0.13	0.21
$\hat{\rho}_3$	0.27	0.50	−0.02	0.11
$\hat{\rho}_4$	0.30	0.52	−0.01	0.08
$\hat{\rho}_5$	0.29	0.58	−0.02	0.15
$\hat{\rho}_6$	0.29	0.62	−0.02	0.23
$\hat{\rho}_7$	0.25	0.51	−0.07	0.07
$\hat{\rho}_8$	0.34	0.50	0.04	0.07
$\hat{\rho}_9$	0.36	0.55	0.11	0.20
$\hat{\rho}_{10}$	0.34	0.52	0.10	0.18
$\hat{\rho}_{11}$	0.27	0.53	0.13	0.15
$\hat{\rho}_{12}$	0.37	0.60	0.19	0.31
$\sigma(\hat{\rho}_1)$	0.07	0.07	0.07	0.07
$\bar{\Delta}$	−0.00188	−0.00283	0.00074	0.00061
$s(\Delta)$	0.00234	0.00277	0.00197	0.00210
T	223	184	223	184

(*Source*: Fama [6] ch. 6)

TABLE 2–8 FAMA'S RESULTS

	$\Delta_t=a_0+a_1i_t+e_t$		$\Delta_t=a_1+a_1i_t+a_2i_{t-1}+e_t$	
	1/53–7/71	3/59–6/74	1/53–7/71	3/59–6/74
a_0	0.00070	0.00212	0.00059	0.00205
a_1	−0.98	−1.44	−0.89	−1.40
a_2			0.11	0.03
$s(a_0)$	0.00030	0.00041	0.00030	0.00044
$s(a_1)$	0.10	0.11	0.12	0.15
$s(a_2)$			0.07	0.08
Coefficient of determination	0.29	0.48	0.30	0.48
$s(e)$	0.00196	0.00203	0.00195	0.00204
$\hat{\rho}_1(e)$	0.09	−0.00	−0.05	−0.03
$\hat{\rho}_2(e)$	0.13	0.14	0.13	0.14
$\hat{\rho}_3(e)$	−0.02	0.02	−0.04	0.02
T	223	184	223	184

(*Source*: Fama [6] ch. 6)

to the sudden and rapid inflation,[14] but the Japanese call money market seems to be far less efficient than the U.S. treasury bill market in the sense that it does not fully utilize the available information.

This conclusion sounds quite natural because there are several reasons to believe that the Japanese call money market is not efficient. The attitude of the monetary authority in Japan is usually said to have a significant effect on the determination of the call money rate. This is so first because the demand for call money is closely related to the ease (availability) of borrowing from the central bank which has a lower interest rate and second because the monetary authority has direct control over the amount of lending of commercial banks and thus affects the demand for call money. Third, because there exist extended regulations on most Japanese financial markets, some kind of distortions might be introduced into the call money market even though it is not regulated directly. Thus, it is already a common sense that the Japanese call money market is more or less regulated by the monetary authority. Moreover, it is believed that the demanders who are always bigger banks tend to dominate the suppliers who are always smaller banks in the determination of the market interest rate. These pecuriarities of the Japanese call money market might have produced the situation in which the available information about the future real rates is not fully utilized in the determination of the interest rate.

In the previous chapter we obtained smaller weights for more recent experiences than for more distant experiences. This might imply the length of adjustment time necessary for the market to use the available information. In this sense our finding in Chapter 1 is consistent with the finding in this chapter that the Japanese market is not efficient.

There is still another explanation for the difference in the validity of the Fisher effect

[14] As the cause of the inefficiency and the changing real rate in the extended analysis for 1959–1974, Fama himself pointed out the unreliability of CPI data due to the price control exercised during Aug. 1971–June 1974.

in Japan and the U.S. It comes from an international consideration that the nominal interest rate all countries tend to harmonize due to the international mobility of capital. If we assume perfect capital mobility,[15] in a country with a higher (lower) rate of inflation than the world average the nominal rate necessarily underadjusts (overadjusts) to domestic inflationary expectations. From this point of view, the fall of the real rate in Japan estimated in Figure 2–4 could be understood as the result of the higher rate of inflation in Japan than the world average. On the other hand, the almost perfect validity of the Fisher equation for the U.S. during the period 1952–1971 can be explained by noting that the rate of inflation in the U.S. during that period was almost equal to the average rate of inflation world-wide.

The above arguments about the validity of the Fisher effect can be summarized as follows. The real interest rate in Japan has changed over time due both to market inefficiency and to a sudden and very rapid rise in the rate of inflation. The latter might have caused the change in the real rate both through the difficulty of forming accurate expectations and through the underadjustment of the nominal rate to inflationary expectations and through the underadjustment of the nominal rate to inflationary expectations due to the international harmonization of the nominal rate. Thus, we might be tempted to conclude that the Fisher effect has not fully worked in Japan and the real interest rate has changed during the last decade.

Fisher himself, however, clearly distinguished between "full equilibrium" and "the transition period" or "disequilibrium" and stressed that the nominal rate adjusts by the amount of expected inflation and keeps the real rate constant only during a period of "full equilibrium".[16] The transition period, he argued, would be characterized by an increase in the nominal rate and a decrease in the real rate which is the major determinant of the trade cycle:

> "Yet, in actual practice, for the very lack of this perfect theoretical adjustment, the appreciation or depreciation of the monetary standard does produce a real effect on the rate of interest, and that a most vicious one. This effect, in times of great changes in the purchasing power of money, is by far the greatest of all effects on the rate of interest". ([11], p. 493)

Since he clearly asserted that the major influence of inflation was on the real interest rate during the transition period, it is not appropriate to say that the Fisher equation is not valid in Japan. With respect evaluating to the validity of the Fisher effect, it is more important whether the period analysed is better approximated by "full equilibrium" or a "transition period". In this sense, we should conclude that the U.S. was close to "full equilibrium" during the period 1952–1971, while Japan was just in a "transition period" from 1966 to 1976.

Section 4. Summary

Our purpose in this chapter has been to examine the validity of the Fisher effect empirically. In Section 1 we surveyed previous studies for the U.S. which asserted the validity of the Fisher equation and the hypothesis of constant real interest rates. In Section 2 we

[15] For this point I am indebted to the discussion with Professor H. Inagaki of Tokyo Municipal University.

[16] See Fisher [9] [10] [11], Rutledge [34].

restated the analytical framework of Fama's theory of efficient markets on which our analysis of Section 3 is based. According to the framework we analyzed the Japanese call money market in Section 3 to find that market has been inefficient during the last decade in the sense that it has not fully utilized the available information and that the real interest rate has changed over time. Contrasting our results with findings for the U.S., we determined that the Japanese call money market is far less efficient than the U.S. treasury bill market, but real interest rates have changed over time in both countries during periods of sudden and rapid inflation.

International capital mobility also has an important effect on the empirical validity of the Fisher effect, since it leads to the international harmonization of interest rates. From this point of view the apparent underadjustment of the nominal rate in Japan can be understood as the result of the higher rate of inflation than the world average.

Therefore, in Japan the Fisher effect has not worked in such a manner that the real rate remained unchanged over time. Fisher himself, however, stressed that the real interest rate changes during "the transition period." Thus we should conclude that Japan has been in a "transition period" during the last decade, while the U.S. was close to "full equilibrium" from 1952 to 1971.*

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