Philippine Review of Economics and Business
Volume XXVII, No. 1, June 1990

THE RELATIONSHIP BETWEEN THE VARIABILITY OF OUTPUT AND THE VARIABILITY OF NOMINAL SHOCKS: JAPAN 1970-1985

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This study attempts to relate the level of output to the variability of nominal and real shocks. Expectations are assumed to be formed rationally and Mishkin's (1983) empirical methodology is used. The results obtained using Japanese data for the January 1970 to December 1974 period show that higher and more variable rates of inflation are associated with a rise in unemployment and a reduction in real output. The converse case holds for the April 1981 to June 1985 period in which the tradeoff between output growth and inflation became more favorable. The results indicate that the variability of real or aggregate supply shocks, rather than the variability of nominal shocks, was significant in explaining the response of output.

1. Introduction

A major proposition derived from business cycle theories (Lucas, 1973) is that as the variance of nominal shocks increases, real output becomes less responsive to general price movements induced by these shocks. This study attempts to test whether the variability of output and the variability of nominal shocks are negatively correlated. Specifically, this study tests Friedman's view (1977) that more variable rates of inflation, which tend to be associated with higher rates of inflation, tend to reduce the efficiency of the price system and bring about an increase in unemployment and a reduction in the rate of change of real output. In the case of Japan, as shown in Appendix A, the Phillips curve has not only exhibited different slopes over time, but more importantly, the curve seems to have shifted between the '60s and '70s.[1] A related question to be addressed is whether there are qualitative differences in the way the Japanese economy was affected by the oil shocks in 1973 and 1979.

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[1]Ideally, one would have wanted to use data for the '60s and '70s, but comparable data were not available for the '60s.

Hamada and Kurosaka (1984) studied the applicability of Okun's Law to the postwar Japanese economy. They found that compared to the U.S., the output gap is extremely responsive to the rate of unemployment in Japan. However, there appear to be conceptual problems with the definition of the unemployment rate for Japan which lead to a measure of unemployment with very little variability.[2] When the unemployment statistics in Japan were redefined to conform with those in the U.S., the degree of responsiveness of the output gap to the unemployment rate was halved. One is left with the conclusion that the Okun coefficient is large and unstable, depending on how the rate of unemployment is defined. In order to avoid these pitfalls, it is necessary to use data other than the unemployment rate to verify the changes in the slope of or shifts in the Phillips curve. Hamada and Kurosaka contend that,

> "Before the (first) oil crisis, during the period of preconditioning for the high rate of growth as well as during the period of very fast growth, a 1% decline in the unemployment rate in Japan was associated with an increase in output of 18-32%. After the oil crisis, on the other hand, the Okun coefficient declined to 13, narrowing its divergence from the American value but still being a very high value. Thus unemployment has begun to respond slightly more rapidly to changes in output" (Hamada and Hurosaka, 1984, p. 77.)

What factors can account for this purported change in the response of the economy?

Froyen and Waud (1985) suggest these possible explanations in addition to Friedman's view:

(1) Following Lucas (1973), the new classical view of the output-inflation tradeoff posits that the response of real output is inversely related to the variability of inflation and aggregate demand.

(2) *Ceteris paribus,* supply-side shocks, such as the oil shock, could cause an increase in the rate of inflation and a decrease in the rate of change of real output.

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[2]See, for example, the discussions in Hamada, Koichi and Yoshio Kurosaka, "The Relationship Between Production and Unemployment in Japan," *European Economic Review* 25 (1984), pp. 72-76; and Grossman, Herschel I. and William S. Haraf, "Shunto, Rational Expectations, and Output Growth in Japan," NBER Working Paper No. 1144, June 1983, p. 6.

Most of the studies testing Lucas' hypothesis regarding the negative correlation between the variance of the nominal disturbances and the response of real output have used cross-section data of several countries. Alberro (1981), Hercowitz (1983), Lucas (1973), Attfield and Duck (1983) have found support for the hypothesis. Hercowitz (1983) postulates that as transaction frequency increases with expected inflation, the flow of price information across markets is accelerated. This results in less misperception of money shocks, and hence, lower sensitivity of real output to nominal shocks. In Froyen and Waud's study of the relationship between the rate of inflation and the growth of real output in the U.S. over the past three decades, only Friedman's view and (2) were supported by the empirical evidence.

In the case of Japan, the correlation coefficient between output growth and the rate of inflation (using the wholesale price index) is negative in the period January 1970 to December 1974 (–0.05), positive in the periods January 1975 to March 1981 (0.008) and April 1981 to June 1985 (0.118). A negative correlation coefficient between output growth and the rate of inflation is consistent with a positive relationship between the rate of inflation and the rate of unemployment, i.e., a positively-sloped Phillips curve if as is normally assumed, unemployment and output growth are negatively related. The converse also holds.

The variability of the overall rate of inflation in the January 1970 to December 1975 period of 0.016 percent is larger than that for the January 1975 to March 1981 period of 0.003 percent and the April 1981 to June 1985 period of 0.005 percent. Hence, there is casual empirical evidence that seems to be consistent with the Lucas hypothesis. When the variability of inflation increases, as in the first period, there is a deterioration in the tradeoff between output growth and inflation. Demand variability, proxied for by the variance of unanticipated monetary growth, averaged 0.11 percent in the January 1970 to December 1974 period, 0.05 percent in the January 1975 to March 1981 period as well as in the April to June 1985 period.[3] Aggregate supply variability, proxied by the variance of the detrended log of the oil price index, averaged 2.5 percent in the January 1970 to December 1974 period, 1.5 percent in the January 1975 to March 1981 period, and 1.0 percent in the April 1981 to June 1985 period. In general, therefore, both proxy measures indicate greater variability in the earlier period and is consistent with the greater variability of the overall rate of inflation in this period as well.

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[3]This proxy for demand variability ignores velocity shocks.

This study is divided into the following: the second part discusses the empirical methodology, the third discusses the results, and the summary and conclusions are presented in the fourth.

2. Empirical Methodology

The methodology used follows that in Mishkin (1983). Assume that the money supply process may be characterized by:

(1) $M_t = Z_{t-1}\tau + u_t$

where:

M_t : rate of growth of money

Z_{t-1} : a vector of variables used to forecast M_t available at time $t-1$

τ : a vector of coefficients

u_t : an error term which is assumed to be serially uncorrelated and uncorrelated with Z_{t-1}

Explanatory variables in the money forecasting equation were chosen on the basis of F-tests which tested the joint significance of the particular variable under consideration and its lags. The lagged values of the variable in question were tested in an equation which included other explanatory variables. Following Mishkin (1983, p. 114), the lagged values of each variable were retained only if they were jointly significant at the 5 percent level using an F-test.

The industrial production equation is specified as:[4]

(2) $IP_t = C_o + \sum_{i=o}^{..} \beta_i (M_{t-i} - M^e_{t-i}) + \sum_{i=o}^{..} \delta_i (M^e_{t-i})$
$+ \alpha.TREND + V1.VARM_t + S1.VARO_t + et$

where:

IP_t : the log of the industrial production index

[4]There are various ways of inducing stationarity in the output data. One way is to use the growth rate of the industrial production index as the dependent variable. This assumes that the time series for industrial production is best described by a stochastic trend. This method of inducing stationarity in the output data was inferior to the one used here in which the time series for output is characterized as stationary fluctuations around a deterministic trend. See Gochoco (1986).

C_o : constant

M_t^e : optimal linear forecast of monetary growth based on information available at time t–1

$TREND$: a linear time trend

$VARM_t$: variability of the aggregate demand shock measured as a 12-month moving average of the variance of forecast error of monetary growth

$VARO_t$: variability of the aggregate supply shock measured as a 12-month moving average of the variance of the detrended log of the oil price index.

$V1, S1, \alpha$: coefficients

$\beta_1, \delta 1$: coefficients on unanticipated and anticipated monetary growth.

e_t : error term

Equation (2) relates the output *level* to anticipated and unanticipated monetary growth and the variances of the detrended oil price index and money forecast errors. Friedman's formulation allows the natural rate of output to be variable. Equation (2) is based on the assumption that anticipated and unanticipated monetary growth lead to cyclical deviations around the trend of real output while a time trend variable and the variability of demand and supply lead to variations in the natural rate of output itself. The manner in which the trend in real output is modeled has important implications for the results obtained. Parkin (1984) concludes that if trends and seasonality are modeled properly, the business cycle in Japan is a classical one. Nelson and Plosser (1982) suggest various ways to induce stationarity in the output data. They further point out that in many cases, it is incorrect to assume that the trend in real output is itself unchanging. Incorporating Nelson and Plosser's results, Gochoco (1986) shows that although expectations are formed rationally, money is not neutral in Japan, contrary to Parkin's findings.[5]

Friedman (1977) suggests that unemployment will tend to rise if

[5]Parkin, in a letter to the author, admitted that there may be some problems with his earlier findings and is re-doing his earlier study.

there are impediments to the efficient functioning of the price system such as institutionally-based rigidities in prices and high variability of inflation. Since a high degree of variability in inflation is usually accompanied by high levels of inflation, there will also tend to be a positive relationship between inflation variability and unemployment. Empirically, one would expect a positive relationship between the level of inflation and the level of unemployment. Equivalently, one would expect a negative relationship between the level of inflation and the level of output.[6]

Following Friedman's analysis, the natural rate of output depends on the variability of inflation in addition to a time trend. The variability of inflation depends on the variability of aggregate demand and supply proxied in our model by the variances of money forecast errors and the detrended log of the price index, respectively.

There are cyclical deviations from the natural rate of output. In equation (2), such effects are captured by the coefficients of anticipated and unanticipated monetary growth. Anticipated monetary growth is included in the output equation because evidence presented in Gochoco (1986) rejects neutrality. Omitting anticipated monetary growth could lead to a bias in the estimated coefficients. The lag lengths on anticipated and unanticipated monetary growth are based on the results of the F-test and are the same as in Gochoco (1986).

As in Froyen and Waud's study, the number of lags in calculating $VARM_t$ and $VARO_t$ is arbitrarily specified. Taking expectations of equation (1) yields:

$$M_t^e = Z_{t-1}\tau$$

Substituting (1a) into (2) yields

$$(3)\quad IP_t = C_o + \sum_{i=o}^{"} \beta_i (M_{t-i} - Z_{t-1-i}\tau) + \sum_{i=o}^{"} (Z^e_{t-1-i}\tau)$$
$$+ \alpha \cdot TREND + V1 \cdot VARM_t + S1 \cdot VARO_t + e_t$$

Equations (1) and (3) are estimated jointly using an iterative

[6]Ideally, one would want a formal test of the Lucas hypothesis that the responsiveness of real output to nominal shocks decreases as the variance of nominal shocks increases relative to the variance of real shocks. It is not possible to do so using equation (2) because the new classical framework posits that the coefficients on anticipated and unanticipated monetary growth are themselves functions of the variances of the shocks to the supply and demand curves.

generalized least squares procedure. As in Gochoco (1986), the τ's in (1) and the τ's in (3) are constrained to be equal. The estimation procedure corrects for serial correlation using a first-order autoregressive specification for e_t, i.e.,

$$e_t = \rho e_{t-1} + \eta\tau.$$

Seasonally unadjusted monthly data, covering the period from January 1970 to June 1985, are used. In carrying out the joint nonlinear estimation procedure, the contemporaneous variance/covariance matrix is calculated on each iteration. In order for the coefficients in (3) to be identified, the contemporaneous correlation between u_t in (1) and e_t in (3) must be known. For consistency, this correlation must be zero.

3. Empirical Results

The same money forecasting equation in Gochoco (1986) is used here. The Bank of Japan uses the monetary aggregate *M*2 + *CDs* to conduct policy. *M*2 is the sum of currency in circulation, demand and time deposits, and *CDs* are certificates of deposit. Suzuki (1985) has shown that this definition of money has a strong correlation with lagged income whereas *M*1 has a strong correlation with contemporaneous income. The relevant variables in the money growth equation are six lags each of the growth rates of money, the wholesale price index, and the industrial production index. The lag length and the relevant variables were decided on the basis of the results of the F-test for joint significance.

Equation (3) was estimated over the periods January 1970 to December 1974, January 1975 to March 1981, and April 1981 to June 1985. These break points in the data were selected arbitrarily, in part based on the picture in Appendix A which shows a shift in the Phillips curve between the early part of the '70s and the later part, as well as to be able to compare the effects of the variability of the oil shocks in 1973 and 1979 on the Japanese economy.

The results in Table 1 indicate that the variability of aggregate supply, proxied for by the variance of the detrended oil price index, had a significantly negative effect on real output in the period January 1970 to December 1984. This finding is consistent with Friedman's view that the more variable the rate of inflation is, the lower will be the natural rate of output. It is also consistent with the casual empirical evidence presented earlier of a negative correlation between output growth and inflation in this period, i.e., a positively-sloped Phillips curve. Very few

Table 1 – January 1970 to December 1974

$$IPT_t = C_o + \sum_{i=o}^{11} \beta_i (M_{t-i} - Z_{t-1-i}\tau) + \sum_{i=o}^{11} \delta_i (Z^e_{t-1-i}\tau)$$

$$+ \alpha.TREND + V1.VARM_t + S1.VARO_t + \rho e_{t-1} + \eta_t.$$

C_0 :	3.518**	(0.319)	V1 :	4.896	(13.969)
β_0 :	–0.547	(0.670)	S1 :	–1.253**	(0.615)
β_1 :	–1.201	(0.850)			
β_2 :	1.235	(1.798)			
β_3 :	2.312	(1.667)	ρ :	–1.00	(0.060)
β_4 :	–0.426	(2.031)	SSE :	0.00001	
β_5 :	–1.234	(1.990)	R^2 :	0.99	
β_6 :	–1.958	(2.072)			
β_7 :	–3.098	(2.277)			
β_8 :	3.984**	(1.945)			
β_9 :	0.017	(1.615)			
β_{10} :	0.649	(1.580)			
β_{11} :	–1.714	(1.852)			
δ_0 :	2.181	(1.878)			
δ_1 :	4.941	(1.941)			
δ_2 :	–0.227	(1.834)			
δ_3 :	2.065	(2.378)			
δ_4 :	–2.317	(2.808)			
δ_5 :	–0.201	(3.330)			
δ_6 :	4.517	(3.088)			
δ_7 :	0.689	(2.651)			
δ_8 :	2.810	(2.222)			
δ_9 :	1.195	(2.449)			
δ_{10} :	4.442**	(1.858)			
δ_{11} :	4.754**	(2.171)			
α :	0.011**	(0.003)			

Note: In this and in succeeding tables, SSE = sum of squared errors, * indicates significance at the 1% level and ** indicates significance at the 5% level. Standard errors are in parentheses.

of the coefficients of anticipated and unanticipated monetary growth are significant and that of the variability of demand is insignificant.

It must be noted, however, that the autocorrelation coefficient is unity and significant. Hence, differencing seems to be indicated. Using the alternative method of inducing stationarity in the data, namely, using the first difference of the logs of output as the dependent variable and dropping the time trend as a regressor on the right-hand side, produces unconvincing results. While the autocorrelation coefficient is lower, none of the variables including both proxies for the variability of inflation are significant.[7]

Table 2 shows the results for the period January 1975 to March 1981. Most of the coefficients on anticipated and unanticipated mone-

Table 2 – January 1975 to March 1981

$$IPT_t = C_o + \sum_{i=o}^{11} \beta_i (M_{t-i} - Z_{t-1-i}\ \tau) + \sum_{i=o}^{11} \delta_i (Z_{t-1-i}\ \tau) + \alpha.TREND + V1.VARM_t + S1.VARO_t + \rho e_{t-1} + \eta_t.$$

C_0 :	3.758**	(0.063)	V1 :	–15.295	(18.025)
β_0 :	2.690**	(0.638)	S1 :	0.016	(0.193)
β_1 :	1.874**	(0.729)			
β_2 :	0.088	(0.942)			
β_3 :	2.238**	(0.874)	ρ :	–0.315**	(0.151)
β_4 :	–0.966	(0.878)	SSE :	0.006	
β_5 :	2.385**	(0.914)	R^2 :	0.99	
β_6 :	–0.141	(1.018)			
β_7 :	1.990**	(0.889)			
β_8 :	–0.013	(0.917)			
β_9 :	2.961**	(0.913)			
β_{10} :	1.152	(0.870)			
β_{11} :	2.270**	(0.982)			
δ_0 :	1.621	(0.999)			
δ_1 :	0.819	(1.017)			
δ_2 :	–2.233**	(1.078)			
δ_3 :	3.138**	(1.033)			
δ_4 :	–2.697**	(1.196)			
δ_5 :	2.677**	(1.078)			
δ_6 :	–1.174	(1.276)			
δ_7 :	5.739**	(1.197)			
δ_8 :	–1.340	(1.168)			
δ_9 :	4.930**	(1.143)			
δ_{10} :	2.022**	(0.799)			
δ_{11} :	6.044**	(1.094)			
α :	0.005**	(0.0003)			

[7]These results are available from the author upon request.

tary growth are significantly positive. Neither of the measures of inflation variability is significant although $S1$ is now positive. It may also be worth noting that the correlation between output and the rate of inflation is weakest in this period (0.008).

Table 3 shows the results for the period April 1981 to June 1985. Again, most of the coefficients on anticipated and unanticipated monetary growth are significantly positive. Now, however, $S1$ is significantly positive. Again, this is consistent with Friedman's analysis and

Table 3 – April 1981 to June 1985

$$IPT_t = C_o + \sum_{i=o}^{..} \beta_i \, (M_{t-i} - Z_{t-1-i}\,\tau) + \sum_{i=o}^{..} \delta_i \, (Z_{t-1-i}\,\tau)$$

$$+ \alpha.TREND + V1.VARM_t + S1.VARO_t + \rho e_{t-1} + \eta_t.$$

C_0 :	3.529**	(0.095)	V1 :	–8.175	(28.167)
β_0 :	0.349	(1.271)	S1 :	2.726**	(0.513)
β_1 :	–0.792	(1.295)			
β_2 :	6.076**	(1.416)			
β_3 :	6.405**	(1.413)	ρ :	–0.55**	(0.199)
β_4 :	2.339**	(0.960)	SSE :	0.001	
β_5 :	2.184*	(1.171)	R^2 :	0.99	
β_6 :	5.097**	(1.450)			
β_7 :	4.411**	(1.365)			
β_8 :	0.034	(1.270)			
β_9 :	0.209	(1.323)			
β_{10} :	0.649	(1.486)			
β_{11} :	1.806	(1.215)			
δ_0 :	4.585**	(1.225)			
δ_1 :	2.557*	(1.378)			
δ_2 :	3.563**	(1.010)			
δ_3 :	5.417**	(0.876)			
δ_4 :	7.248**	(1.106)			
δ_5 :	6.160**	(1.433)			
δ_6 :	2.851**	(1.169)			
δ_7 :	1.786	(1.196)			
δ_8 :	1.146	(1.274)			
δ_9 :	4.822**	(1.692)			
δ_{10} :	3.815**	(1.629)			
δ_{11} :	3.465**	(1.229)			
α :	0.005**	(0.0005)			

with the earlier finding of a strong positive correlation between output and inflation in this period (0.028). This implies that the Japanese economy faced a more favorable tradeoff in this latter period than in the earlier period.

In general, the results show that variability of the detrended oil price index, a measure of aggregate supply variability, had significant effects on the Japanese economy in the periods January 1970 to December 1974 and April 1981 to June 1985. The direction of the effect changed from negative to positive in these two periods, implying a more favorable tradeoff between output and inflation in the latter period. The results also indicate that the factors which contribute to cyclical deviations of output about its trend were important in the middle and later periods.

4. Summary and Conclusions

In this study, an attempt was made to test whether the variability of output and the variability of nominal shocks are negatively related. Friedman's hypothesis states that higher and more variable rates of inflation reduce the efficiency of the price system and bring about an increase in unemployment and a reduction in the rate of change of real output.

The results of this study indicate that aggregate supply variability, proxied for by a 12-month moving variance of the detrended oil price index, significantly affected the Japanese economy in the periods January 1970 to December 1974 and April 1981 to June 1985. *S*1, the coefficient of aggregate supply variability, was significantly negative in the early period and significantly positive in the later period. Aggregate demand variability, proxied for by a 12-month moving variance of money forecast errors, was never statistically significant in any period.

Anticipated and unanticipated monetary growth, assumed to lead to deviations of actual output from its natural rate, were largely significantly positive in the periods January 1975 to March 1981 and April 1981 to June 1985.

Given that the correlation coefficient between inflation and output growth is negative in the period January 1970 to December 1974, and positive in the periods January 1975 to March 1981 and April 1981 to June 1985, the results obtained in this study seem to support Friedman's hypothesis. When the overall rate of inflation is highly variable as in the first period, the lower will be the natural rate of output and the less favorable the policy tradeoff will be between output growth and inflation. The converse holds for the April 1981 to June 1985 period.

The study, however, does not address the important question of the role of policy accommodation of relative shocks, such as energy and food price shocks, in determining the behavior of both inflation and relative price variability (Fisher 1981, 1982). Such policy actions would have an effect on real output via their effect on inflation and the variability of inflation, according to Friedman's hypothesis. Perhaps such a study would shed light on why output growth and inflation are only weakly correlated in the period January 1985 to March 1981, for example, and hence why the measures of aggregate demand and supply variability do not have significant effects on the natural rate of output. It would also seem that the other factors such as the degree of wage indexation would have an effect on the degree of inflationary impact of the oil price changes in addition to the degree of accommodation by policymakers. Hamada and Hayasahi (1985, p. 105) contend that,

> "In spite of declining productivity due to the first oil crisis, the Japanese labor negotiation process kept real wages constant or even increasing. After the second oil crisis, however, real wages moved rather flexibly. Labor union leaders seemed to have learned the lesson."

If this is the case, the effective indexation of real wages after the firs oil crisis would have led to a larger inflationary impact on the Japanes economy, and, by extension of Friedman's analysis, a less favorabl policy tradeoff. These are subjects of further research.

List of Variables

M2+CD (Average Outstanding), 100 Million Yen

IP Industrial Production Index (Based on Value Added), 1977 Average = 100

WPI Wholesale Price Index, 1980 Average = 100

Oil Price Index (Nikkei Index), in terms of Yen, 1970 Average = 100

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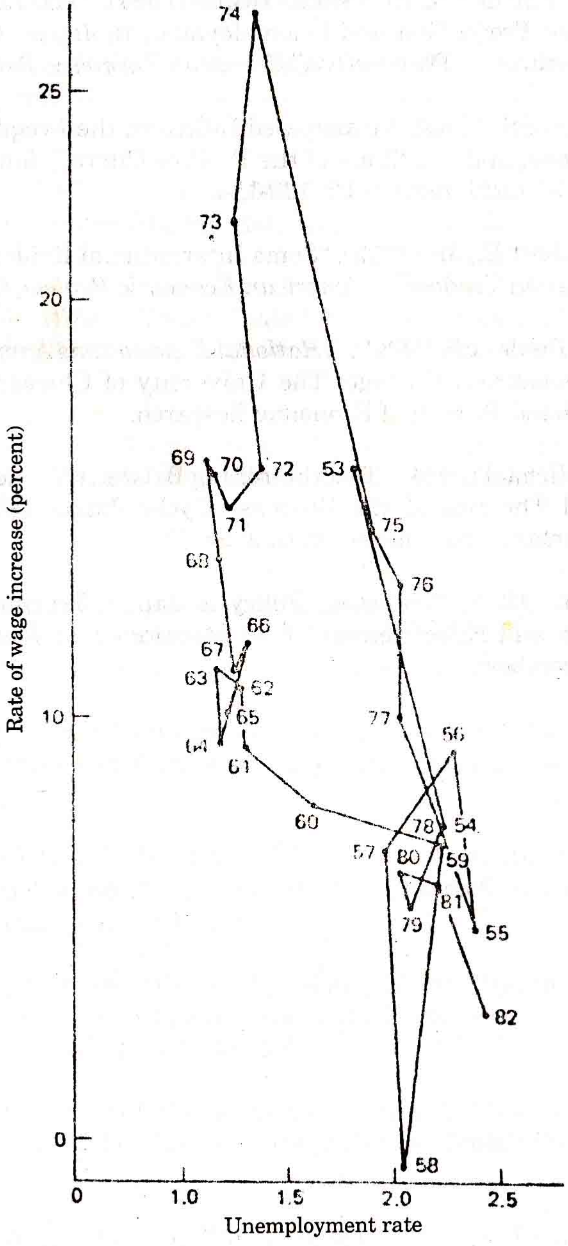

Appendix A – The Phillips Relationship in Japan

Source: (Hamada and Hayashi, 1985, p. 104).