Department of Economics
UNIVERSITY OF CANTERBURY
CHRISTCHURCH, NEW ZEALAND

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Alfred V. Guender and Robin Young

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Department of Economics, University of Canterbury
Christchurch, New Zealand

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ABSTRACT

This paper presents a test of the Friedman hypothesis: Friedman (1977) argues that increases in the average inflation rate are often associated with a rise in inflation variability and hence inflation uncertainty. With reference to the importance of the time horizon in analysing inflation uncertainty we utilise an unobserved components model of inflation which decomposes inflation uncertainty into two measures, one short term, the other long term. Results obtained from a panel of data for the G7 countries provide support for Friedman's basic contention that inflation uncertainty affects real output. In particular, long-term inflation uncertainty has a negative effect on real output. Our results also underscore the importance of central bank independence as a possible influence for fluctuations in real activity.

JEL Classification: E31, E58

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by

Alfred V Guender and Robin Young

Department of Economics
University of Canterbury
Private Bag 4800
Christchurch, New Zealand

Fax: 64-3-364-2635

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1. INTRODUCTION

To examine long term and short term aggregate uncertainty and their effect on real variables, it is important to consider the following: first, is there a systematic tendency for periods of high average inflation to be associated with high variability of the aggregate inflation rate and greater uncertainty about future rates of inflation? Second, does this uncertainty about inflation have effects on real variables by creating inefficiencies in the allocative workings of the price system?

These two propositions have been related via the Friedman hypothesis laid out in his 1977 Nobel lecture.[1] The assertion centres on the notion that the inflation rate and the variability in the inflation rate are positively correlated and that the positive relation between unemployment and inflation that is sometimes observed from fitted Phillips curves hides an underlying positive relationship between unemployment and inflation uncertainty. This effect can also be extended to other real variables such as industrial production or real growth, which in this framework would be negatively related to inflation uncertainty.

Our aim in this paper is to test the Friedman hypothesis by using a model of inflation uncertainty developed by Ball and Cecchetti (1990) which specifically breaks down the uncertainty relation into two effects; one short term, the other long term. The purpose of such a decomposition is to reconcile some of the ambiguities in the literature concerning the uncertainty relation with respect to the inflation rate[2] and, further, to discover if the resultant inflation uncertainty has effects on real economic variables. Such a decomposition also allows us to use a panel of data covering the countries in the Group of Seven from 1960 to 1993. This approach, it can be argued, is distinct from previous studies for the following reasons. First, studies regarding the average inflation-uncertainty thesis have tended only to look at one notion of uncertainty. Second, the majority of previous papers that have examined the Friedman hypothesis have used survey data to determine inflation uncertainty, and have focused, empirically, on the U.S.[3]

Finally, while Ball and Cecchetti look specifically at the first proposition above we are extending their analysis with respect to proposition two. That is, we examine Friedman's contention that uncertainty has real effects and these should be captured by the long run measure of uncertainty introduced in the present paper.

2. THE FRIEDMAN HYPOTHESIS

In his Nobel lecture, Friedman (1977) outlines his views on the Phillips curve relationship and takes the debate on to a "third stage". According to Friedman:

> "an increased variability of actual or anticipated inflation may raise the natural rate of unemployment in two rather different ways." (p.26.)

Firstly, higher inflation volatility produces "institutional" changes by shortening the optimal contract period for non-indexed agreements and promotes an increase in the proportion of contracts involving indexing. This process takes time and, meanwhile, prior arrangements create rigidities, and the fact that indexing arrangements are far from perfect (Gray (1976)), a reduction in productive efficiency can be postulated. However, the impact of this efficiency/real resource utilisation cost on unemployment is not all that clear.

Secondly, volatility of inflation impacts upon the efficiency of the price system in a way which complicates the recognition of relative price changes. True signals in relative price movements, as Hayek (1945) has demonstrated, serve the purpose of transmitting knowledge throughout an economy in a compact, efficient and low-cost way. So, the real function of a price system is to act as a mechanism for communicating information where the relevant information is contained in relative prices which in practice are transmitted via absolute prices. Friedman (1977) explains how the signal extraction problem is complicated by a more volatile rate of inflation.[4]

Friedman has suggested that the effect of an increase in volatility of inflation is a transitory one but the effect is lengthy in the sense that it may take decades to unwind because in the long, long run the natural rate remains independent of monetary phenomena.[5] This is because once an inflation is allowed to occur, the public's attitude towards the monetary authorities will change. This has the effect of creating a more uncertain environment which will take a reasonable length of time to correct.

The Friedman proposition can be more neatly summarised along the lines of Levi and Makin (1980):

(i) That higher errors in forecasting inflation should be correlated with higher uncertainty about the rate of inflation, and

(ii) higher uncertainty about inflation could be associated with lower employment and industrial output. This is because the natural rate of real output, necessarily a longer term phenomena, is detrimentally affected for sustained periods. We therefore require an appropriate measure of this uncertainty.

3. THE MODEL

Fischer (1981) explains the important distinction between inflation variability and inflation uncertainty. That is, it can sometimes be possible that even when high inflation results in a high variation in the inflation rate the movements can be, to a large extent, predicted. This means that the variance of unanticipated inflation may not be all that pronounced. Furthermore, an increase in variability does not necessarily imply greater uncertainty. In this regard, Ball and Cecchetti's ((1990) henceforth B&C) paper is an attempt to resolve the empirical disputes regarding the inflation-uncertainty relation. This they do by proposing the distinction between short term and long term uncertainty.

B & C concentrate upon permanent and temporary shocks to inflation. They classify permanent shocks as shifts in the trend rate of inflation and temporary shocks as fluctuations around the trend. For example inflation uncertainty in the next period depends primarily upon the variance of temporary shocks while uncertainty about inflation over a longer time frame depends more on the variance of permanent shocks. Similarly, Evans (1991) decomposes the inflation-uncertainty notion into temporal (short term) and intertemporal (longer term) measures. The key finding of both studies is that the level of inflation has a more pronounced effect on the variance of permanent (long term) shocks than on the variance of temporary (short term) shocks, and therefore a more apparent effect on uncertainty at long horizons.

B&C utilise a univariate model in which there are two innovations to inflation. Two relations make up the basic model,

$$\pi_t = \hat{\pi}_t + \eta_t \qquad \eta_t \sim N(0,\sigma_\eta^2) \tag{1}$$

$$\hat{\pi}_t = \hat{\pi}_{t-1} + \varepsilon_t \qquad \varepsilon_t \sim N(0,\sigma_\varepsilon^2) \tag{2}$$

where $\hat{\pi}_t$ = trend inflation, which is not directly observable and follows a random walk. $\hat{\pi}_t$ is the optimal forecast of the inflation rate in all future periods.

π_t = actual inflation rate = trend inflation + white noise.

η_t = temporary shock: captures events that affect inflation temporarily but do not affect the trend. This can include non-accommodated supply shocks, fluctuations in the velocity of circulation of money, etc..

ε_t = permanent shock: captures circumstances that change trend inflation. ε_t is negative, for example, when the monetary authorities create a recession in order to disinflate and is positive when the authorities allow a supply shock to be accommodated which then leads to trend inflation rising.

η_t and ε_t are uncorrelated: $E_t[\eta_t \ \varepsilon_t] = 0$

Within this framework there are two explanations why inflation uncertainty may be high when the average (trend) inflation rate is high. Firstly, inflation may vary to a greater extent around its trend when the trend is high or, secondly, a high trend might itself promote a more erratic trend. The former may occur because of either shocks or monetary control errors which provide the impetus for fluctuations around trend. The latter is driven by changes in the trend growth of the money supply whereby shifts in trend inflation involve a shift in the policy stance of the monetary authorities. Friedman, and others, argue that because permanent changes in inflation involve movements in monetary policy high inflation, ultimately, makes policy less predictable and hence more unstable.

An important feature of the model is that since trend inflation , $\hat{\pi}_t$, follows a random walk inflation is non-stationary meaning that there are events which permanently shift trend inflation with no inclination for inflation to revert to a constant mean. This specification is motivated by the findings reported by Barsky (1987). He finds that US inflation has followed a non-stationary process since 1960. (Statistical tests fail to reject non-stationarity for most countries in B & C 's sample.). Also, although $\hat{\pi}_t$ is assumed to follow a random walk it does not necessarily mean that the transitory shock, η_t , need be white noise. However, B & C experimented with generalisations of the model in which η_t is serially correlated. They find that most of countries and time periods fit the model with white noise errors, and that movements away from trend inflation tend to endure largely for only one period.

In sum, the rate of inflation consists of the unobservable trend component which follows a random walk and a cyclical component which follows a white noise process. This unobserved components model can be shown to be observationally equivalent to an ARIMA(0,1,1) with a single shock. That is, a model in which the change in inflation is an MA(1). From equations (1) and (2) it can be shown that:

$$\Delta\pi_t = \varepsilon_t + (\eta_t - \eta_{t-1}). \tag{3}$$

This is a MA(1) process because, since ε_t and η_t are white noise, only the first autocovariance of $\Delta\pi_t$ is non-zero:

$$\Delta\pi_t = v_t + \theta v_{t-1} \tag{4}$$

It follows that, by setting these variances and covariances implied by (3) and (4) equal, we obtain (5).

$$\sigma^2_\eta = -\theta\sigma^2_v. \tag{5}$$

$$\sigma^2_\varepsilon = (1+\theta)^2\sigma^2_v.$$

The interpretation of such a formulation is that the MA coefficient, θ , lies between 0 and -1, meaning that a shock to inflation is partially reversed in the next period. This is because within this particular specification permanent shocks do not revert to the mean while temporary shocks are always entirely reversed. An implication, in terms of the ARIMA(0,1,1) model, is that a high trend inflation makes changes in inflation more persistent because as the variance of the permanent shocks rises in relation to the variance of temporary shocks θ declines in absolute value.

We now have in place a model whereby we can test the implications of the Friedman hypothesis. *A priori* we would expect that the effect on real output of our measure of the long term inflation-uncertainty would concur with Friedman's conjecture: inflation creates uncertainty regarding its future direction over long horizons which ultimately impacts upon real variables either by reducing the efficiency of the economic system or by upsetting the framework of [monetary] policy.

4. EMPIRICAL RESULTS

We used the B&C model in order to derive estimates for the two variance terms in equation (5) above. This was done for the group of seven countries between 1960 and end 1993. We used monthly observations for inflation and industrial production data.[6, 7]

Before estimating our model, augmented Dickey-Fuller tests on the respective price and production indices were carried out. In each case, we failed to reject at the 10 percent level the existence of non-stationarity.

A series of sub-samples consisting of 40 data points each were then established allowing 10 such sub-samples to be obtained, each containing an estimate of σ_η^2, σ_ε^2, and the mean and variance of industrial output growth

Once estimates were obtained we converted the variance terms to standard deviations and ran least squares regressions of the mean and standard deviation of industrial output (in log form and differenced once) against our two measures at first for each country and secondly for an enlarged G7 panel.

4.1 TESTING THE FRIEDMAN HYPOTHESIS

For the Friedman hypothesis to hold, we require that the coefficient estimate on σ_ε to be significant and negatively related to the mean growth rate of industrial production but positive for the standard deviation. This is because we would expect that an increase in the uncertainty relating to the inflation rate would have a detrimental effect on the average growth rate in the medium to long term and would tend to amplify variations around the mean growth rate.

Tables 1 and 2 summarise our initial results, using monthly data, for the G7 countries individually and as a composite group. For the individual countries there are ten observations (sub-samples) with seven degrees of freedom. For the enlarged sample we have 70 observations.[8]

(i) Mean Growth Rate

For our measure of the mean growth rate of real output the results for the individual countries are not particularly revealing. Only for Japan is the coefficient on σ_ε highly significant, and the sign is as expected. Notice though that the coefficient on σ_ε is negative for all countries except Germany. For Japan, Italy and France the coefficient on σ_η is positive and significant while it is negative and significant for the USA.[9]

For the G7 as a whole, all coefficients are statistically significant with σ_ε, our measure of longer term inflation uncertainty, taking a negative sign as predicted by the Friedman hypothesis. The short term uncertainty measure is positive and statistically significant.

(ii) Standard Deviation of the Growth Rate

Our findings in favour of the Friedman hypothesis are more clear-cut for the standard deviation[10] of industrial output in the individual countries. The coefficient on σ_ε is positive and significant, as expected, for the U.S. and the U.K. For the U.S. the coefficient on σ_η is also positive and significant as it is for Italy as well. For our G7 panel, the coefficient on σ_ε is significant and signed as expected while the coefficient on σ_η turns out to be insignificant.[11]

4.2 CENTRAL BANK INDEPENDENCE

For the panel, another issue arises which concerns the relationship between central bank independence and economic performance.

There is a widespread belief that average inflation rates are lower in countries when central banks are more independent. For example, Cukierman (1992) forwards the hypothesis that inflation should be negatively related to the level of central bank independence.

> "[This] is a consequence of either or both of the following underlying elements: a stronger degree of time preference on the part of political authorities in comparison to that of the central bank, and/or relatively higher concern of the central bank for price stability against the background of private information about its independence." (Cukierman, 1992, p.415).

Alesina (1988) and Grilli et al.(1991) have shown informally that countries with higher degrees of central bank independence tend to have lower average inflation rates. Alesina and Summers(1993) also find a significant negative relationship between indices measuring the degree of central bank independence and average inflation, but they detect no trace of a relationship between central bank independence and real economic performance in sixteen industrialised countries.[12]

Regarding real econmic performance, the basic contention is that a higher degree of central bank independence enhances political stability, reduces risk premia in interest rates and mitigates the political business cycle. This argument postulates a direct link between the extent of economic and political independence of central banks and the level of and variation in real economic performance. Using the framework laid out in section 3, we investigate the contention that central bank independence has a distinct, separately indentifiable impact on real economic performance. We ran the G7 regressions adding an indicator variable for three alternative measures of central bank independence[13] as laid out in Parkin and Bade (1985), Grilli et al. (1991) and Cukierman (1992). While the three alternatives do differ in emphasis and scoring method, they all agree in broad ranking terms that the Bundesbank in Germany, is the most independent and that, comparatively, the Federal Reserve and the Canadian Central Bank are more "independent" than the Bank of England, Banque de France and the Italian central bank, respectively.

The findings are reported in Tables 3 and 4 for our two measures of real economic performance. It appears that the degree of central bank independence impacts on the standard deviation of the growth rate of real output. The coefficients of the first two central bank indices are negative and significant at the 5 percent level while the coefficient of the third index is negative and significant at the 10 percent level. The negative coefficient of the central bank index ties in with our intuition in the sense that the higher the degree of independence, the lower we would expect the amplitude of fluctuations in the real variable. Notice that the sign on the coefficient on σ_ϵ remains positive and significant. When run against the mean growth rate, the central bank index appears to add very little extra information. A reason for this may be that independent central banks are an insurance against more pronounced variations in growth rather than in the absolute level of growth.[14]

4.3 AN EXTENSION TO QUARTERLY DATA

To test the robustness of the above results, we extended our analysis to include quarterly data. The model remained the same but in order to achieve meaningful sub-periods we adjusted the length of each sub-period outward. This alternative yielded 7 instead of 10 sub-samples, allowing for 49 (against 70) data points when our G7 panel was estimated.

With such a change in the horizon we might well expect our estimates for σ_η to be altered although we would not expect the coefficients on σ_ϵ to be changed in a meaningful way. With our monthly data series the coefficient on σ_η, for the mean growth rate, was positive compared with a negative coefficient on σ_ϵ (for the standard deviation the coefficient on σ_η was insignificant). The implications of such a positively signed value is that inflation uncertainty over short horizons (one month in this case) tends to have a positive impact on the mean growth rate. Such an occurrence can be explained via the existence of a short run Phillips curve trade-off. (In terms of a Lucas (1973) type model, this occurs when agents are confused between relative and aggregate price movements.)

Results from the regressions conducted are contained in Tables 5 and 6. The important point to note here is that the signs and significance levels regarding σ_ϵ are unchanged, while the coefficient on σ_η (for our measures of the mean growth rate) is no longer significant.[15] Although this may be owing to the smaller sample size, it could also be explained by the Phillips curve "trade-off" being of a very short duration. Surprisingly, the effect of central bank independence on real output found in the monthly data is not present in the quarterly data.

4.4 SUMMARY OF RESULTS OBTAINED

It would appear from the results presented in this paper that the coefficient estimate on σ_ε is significant and negative for the average monthly growth rate of industrial output but positive for the standard deviation. The initial results for σ_ε are unaffected by the extension from monthly to quarterly data although we do see an interesting implication when applied to our σ_η measure. The basic result (with respect to σ_ε) for the variation in the growth rate is in effect reinforced.

The above results are of interest in the sense that they are in line with the spirit of the Friedman hypothesis. Furthermore, extensions of our inquiry show that central bank independence is likely to have some importance regarding the variation in the monthly growth rate. Coefficients on dummy variables applied were negative, suggesting that the higher the degree of independence the lower the variations around the mean growth rate.

5. CONCLUSION

We have tested the Friedman hypothesis within a framework that models inflation uncertainty with reference to two notions of the time horizon. The central finding is that Friedman's contention is supported and that his emphasis on the longer term is appropriate. The importance of the decomposition of inflation uncertainty into shorter and longer term aspects would appear relevant to such an outcome. This distinction also allowed us a glimpse of the important contrary positive effects made apparent by a short run Phillips curve trade-off. The model could have implications for the duration of such a trade-off and, importantly, the ultimate negative repercussions further out.

There is also evidence suggesting that central bank independence could be a source of insurance against fluctuations around the average growth rate. By, potentially, mitigating the uncertainty effects associated with inflation, a more independent central bank may reduce the longer term costs of inflation.

Our research also shows that a fruitful avenue for further investigation would be concerned with measuring the effects through time of country specific factors. Such a framework would not only highlight the countries where improvements in the inflation-uncertainty relation have been made but would allow an analysis of changing regimes. For example, it would be interesting to see the consequences of the Reserve Bank Act in New Zealand or (for our sample at least) the end result of making the Bank of Italy fully independent.

The policy implications of our analysis would suggest the simplistic view of minimising the degree of uncertainty about future inflation. With respect to monetary policy, this may give support to the prescription of following a fixed growth rate rule for the money supply. The rationale for a fixed money growth rate rule would stem from Friedman and Okun's observations that stop-go monetary policies produce uncertain fluctuations in longer term inflation and have a knock-on effect to output. This is reinforced by a declining credibility of the monetary authorities.

It would appear that the costs of inflation are, in more than the pejorative sense of the word, real costs.

APPENDIX

Ball and Cecchetti's U.C.M. of Inflation

The model considers more than one kind of innovation to inflation. We have seen that:

(1) $$\pi_t = \hat{\pi}_t + \eta_t \qquad \eta_t \sim N(0,\sigma_\eta^2)$$

(2) $$\hat{\pi}_t = \hat{\pi}_{t-1} + \varepsilon_t \qquad \varepsilon_t \sim N(0,\sigma_\varepsilon^2)$$

where, π_t: actual inflation
$\hat{\pi}_t$: trend inflation
η_t: temporary shock
ε_t: permanent shock

We know that from equations (1) and (2) the change in inflation is:

$$\Delta\pi_t = \hat{\pi}_t - \hat{\pi}_{t-1} + \eta_t - \eta_{t-1}$$

Therefore,

(3) $$\Delta\pi_t = \varepsilon_t + (\eta_t - \eta_{t-1}).$$

This is an MA(1) process because, since ε_t and η_t are white noise, only the first autocovariance of $\Delta\pi_t$ is non-zero. From equation (3), it can be seen that:

$$E[\Delta\pi_t\Delta\pi_{t-1}] = E[(\varepsilon_t+\eta_t-\eta_{t-1})(\varepsilon_{t-1}+\eta_{t-1}-\eta_{t-2})]$$

$$= E[\varepsilon_t\varepsilon_{t-1} + \varepsilon_t\eta_{t-1} - \varepsilon_t\eta_{t-2} + \varepsilon_{t-1}\eta_t + \eta_t\eta_{t-1} - \eta_t\eta_{t-2} - \eta_{t-1}\varepsilon_{t-1} - \eta_{t-1}^2 + \eta_{t-1}\eta_{t-2}]$$

$$= E[\eta_{t-1}^2] = -V[\eta_{t-1}] = -\sigma_\eta^2$$

Next,

$$E[\Delta\pi_t^2] = E[(\varepsilon_t+\eta_t-\eta_{t-1})(\varepsilon_t+\eta_t-\eta_{t-1})]$$

$$= E[\varepsilon_t^2+\eta_t^2+\eta_{t-1}^2]$$

$$= \sigma_\varepsilon^2 + 2\sigma_\eta^2.$$

Hence we now have

(i) $$E[\Delta\pi_t^2] = \sigma_\varepsilon^2 + 2\sigma_\eta^2$$

and

(ii) $$E[\Delta\pi_t\Delta\pi_{t-1}] = -\sigma_\eta^2.$$

For the MA(1) process we have

(4) $$\Delta\pi_t = v_t + \theta v_{t-1}.$$

Therefore,

$$E[\Delta\pi_t^2] = E[(v_t+\theta v_{t-1})(v_t+\theta v_{t-1})]$$

$$= E[v_t^2+\theta v_t v_{t-1}+\theta v_t v_{t-1}+\theta^2\theta v_{t-1}^2]$$

$$= \theta_v^2 + \theta^2\sigma_v^2$$

(iii) $$= (1+\theta^2)\sigma_v^2$$

and

(iv)
$$E[\Delta\pi_t \Delta\pi_{t-1}] = E[(v_t+\theta v_{t-1})(v_{t-1}+\theta v_{t-2})]$$
$$= E[\theta v_{t-1}^2] = \theta\sigma_v^2.$$

By setting the variances and covariances implied by (ii) and (iv), and (i) and (iii) respectively, we obtain

(5)
$$\sigma_\eta^2 = -\theta\sigma_v^2 \quad \text{(Temporary shock)}$$

and

$$\sigma_\varepsilon^2 + 2\sigma_\eta^2 = (1+\theta^2)\sigma_v^2$$
$$\sigma_\varepsilon^2 = (1+\theta^2)\sigma_v^2 - 2\sigma_\eta^2$$
$$\sigma_\varepsilon^2 = (1+\theta^2)\sigma_v^2 - 2(-\theta)\sigma_v^2$$
$$\sigma_\varepsilon^2 = (1+2\theta+\theta^2)\sigma_v^2$$
$$\sigma_\varepsilon^2 = (1+\theta)^2\sigma_v^2 \quad \text{(Permanent shock).}$$

TABLE 1: Mean Growth Rate Of Industrial Production Against Decomposed Inflation (Monthly Data: 1960-1993)

EQUATION: $m.d.y. = a_{10} + a_{11}\,\sigma_\eta + a_{12}\sigma_\varepsilon + error$

	U.K.	U.S.A.	GERMANY	JAPAN	CANADA	ITALY	FRANCE	G7#
a_{10}	0.650* (0.331)	0.906** (0.284)	-0.012 (0.397)	0.928* (0.429)	0.985 (0.608)	0.052 (0.193)	-0.128 (0.225)	0.181** (0.070)
$a_{11}(\sigma_\eta)$	-0.882 (0.687)	-2.354* (1.094)	0.823 (1.158)	2.610** (0.671)	-2.009 (1.713)	0.731* (0.313)	1.715* (0.758)	0.483** (0.144)
$a_{12}(\sigma_\varepsilon)$	-0.152 (0.628)	-1.803+ (1.061)	0.078 (1.442)	-6.683** (1.731)	-0.848 (2.604)	-0.822 (0.826)	-2.507 (1.922)	-1.187** (0.422)
D.W. R^2	1.720 0.277	2.174 0.266	2.147 -0.164	1.762 0.641	1.838 -0.074	2.442 0.351	2.041 0.337	- 0.180
Regression F(2,67)	-	-	-	-	-		-	8.500

NOTE:
1. For each country there are 10 data points (sub-samples) derived from 40 observations per sample. For the G.7 we have 70 data points.
2. m.d.y. = mean growth rate of industrial production over each sub-sample
3. σ_η = standard deviation of cyclical inflation over each sub-sample.
4. σ_ε = standard deviation of trend inflation over each sub-sample.
5. Standard errors are reported in parentheses.
6. Growth rate is the log difference of industrial output.

*significant at the 10% level.

**significant at the 5% level.

#if panel is estimated excluding Japan coefficients are not affected in any meaningful way. The coeff. on σ_ε has a ρ value of .07. See also Endnote 8

+ p value is .13.

TABLE 2: Standard Deviation of Growth Rate Of Industrial Production Against Decomposed Inflation (Monthly Data: 1960-1993)

EQUATION: $s.d.y. = b_{10} + b_{11}\sigma_{\eta} + b_{12}\sigma_{\varepsilon} + b_{13}D_F + error$

	U.K.	U.S.A.	GERMANY	JAPAN	CANADA	ITALY	FRANCE	G7
b_{10}	2.754* (1.210)	-0.088** (0.204)	1.895** (0.732)	1.717** (0.270)	1.581* (0.808)	0.052 (0.193)	1.510** (0.445)	0.339** (0.202)
$b_{11}(\sigma_{\eta})$	-3.172 (2.510)	3.589** (0.786)	0.604 (2.134)	-0.702 (0.422)	-1.946 (2.277)	0.731* (0.313)	-0.698 1.622	0.173 (0.411)
$b_{12}(\sigma_{\varepsilon})$	4.981* (2.293)	1.987** (0.762)	-1.746 (2.659)	1.985# (1.088)	6.211# (3.461)	-0.822 (0.827)	4.546 (4.248)	2.318* (1.208)
$b_{13}(d_F)$	-	-	-	-	-	-	6.588** (0.458)	6.463** (0.766)
D.W.	2.689	2.230	1.986	1.595	1.339	2.144	2.100	-
R^2	0.259	0.681	-0.100	0.158	0.249	-0.208	0.966	0.501
Regression F(3,66)	-	-	-	-	-		-	24.049

NOTE:
1. For each country there are 10 data points (sub-samples) derived from 40 observations per sample. For the G.7 we have 70 data points.
2. s.d.y. = standard deviation of growth rate of industrial production over each sub-sample.
3. σ_{η} = standard deviation of cyclical inflation over each sub-sample.
4. σ_{ε} = standard deviation of trend inflation over each sub-sample.
5. Standard errors are reported in parentheses.
6. Growth rate is the log difference of industrial output.
7. d_F: Dummy variable in order to adjust for the massive increase in the variance of industrial output in France, during the sub-sample containing the social unrest of 1968.

*significant at the 10% level.

**significant at the 5% level. # ρ value for Japan and Canada is .115 and .111, respectively.

TABLE 3: Mean Growth Rate and Central Bank Independence - Monthly Data

EQUATION: m.d.y. = $a_{10} + a_{11}\sigma_\eta + a_{12}\sigma_\varepsilon + a_{14}DCB$ + error

G7 - 70 observations

	Parkin & Bade	**Grilli**	**Cukierman**
a_{10}	0.132 (0.109)	0.143 (0.135)	0.118 (0.148)
$a_{11}(\sigma_\eta)$	0.495** (0.146)	0.499** (0.154)	0.535** (0.182)
$a_{12}(\sigma_\varepsilon)$	-1.180** (0.424)	-1.173** (0.426)	-1.208** (0.426)
a_{14}(DCB)	0.018 (0.031)	0.006 (0.017)	0.116 (0.242)
R^2 Regression F(3,66)	0.170 5.727	0.168 5.628	0.169 5.679

NOTE:

1. For each country there are 10 data points (sub-samples) derived from 40 observations per sample. For the G7 we have 70 data points.
2. σ_η = standard deviation of cyclical inflation over each sub-sample.
3. σ_ε = standard deviation of trend inflation over each sub-sample.
4. Standard errors are reported in parentheses.
5. Growth rate is the log difference of industrial output.
6. d_F: Dummy variable in order to adjust for the massive increase in the variance of industrial output in France, during the sub-sample containing the social unrest of 1968..
7. See Endnote 13 for explanation of central bank dummies (DCB).

*significant at the 10% level.
**significant at the 5% level.

TABLE 4: The Standard Deviation of Growth Rate
and Central Bank Independence - Monthly Data

EQUATION: $s.d.y. = b_{10} + b_{11}\sigma_\eta + b_{12}\sigma_\varepsilon + b_{13}DF + b_{14}DCB + error$

G7 - 70 observations

	Parkin & Bade	Grilli	Cukierman
b_{10}	2.121** (0.287)	2.975** (0.304)	1.974** (0.415)
$b_{11}(\sigma_\eta)$	-0.020 (1.382)	-0.552 (0.345)	-0.356 (0.506)
$b_{12}(\sigma_\varepsilon)$	2.173* (1.113)	1.685* (0.963)	2.505** (1.194)
$b_{13}(DF)$	6.336** (0.707)	6.306** (0.608)	6.342** (0.758)
$b_{14}(DCB)$	-0.293** (0.082)	-0.245** (0.039)	-1.178* (0.676)
R^2	0.576	0.686	0.515
Regression F(4,65)	24.431	38.702	19.350

NOTE:
1. For each country there are 10 data points (sub-samples) derived from 40 observations per sample. For the G7 we have 70 data points.
2. σ_η = standard deviation of cyclical inflation over each sub-sample.
3. σ_ε = standard deviation of trend inflation over each sub-sample.
4. Standard errors are reported in parentheses.
5. Growth rate is the log difference of industrial output.
6 d_F: Dummy variable in order to adjust for the massive increase in the variance of industrial output in France, during the sub-sample containing the social unrest of 1968.
7. See Endnote 13 for explanation of central bank dummies (DCB).

*significant at the 10% level.
**significant at the 5% level.

TABLE 5: The Mean Growth Rate
for the G7 - Quarterly Data 1960-1993 - 49 Observations

Equation: m.d.y. = $a_{10} + a_{11}\sigma_{\eta} + a_{12}\sigma_{\epsilon} + a_{14}DCB$ + error

		Central Bank Index		
		Parkin & Bade	**Grilli**	**Cukierman**
a_{10}	0.748** (0.216)	0.772** (0.378)	0.836* (0.440)	1.179** (0.435)
$a_{11}(\sigma_{\eta})$	0.480* (0.290)	0.479 (0.294)	0.466 (0.300)	0.377 (0.303)
$a_{12}(\sigma_{\epsilon})$	-0.822** (0.406)	-0.829* (0.421)	-0.838* (0.416)	-0.911** (0.412)
$a_{14}(DCB)$	-	-0.009 (0.114)	-0.014 (0.061)	0.893 (0.783)
$\bar{R}^2$	0.064	0.043	0.044	0.070
F(2,46)	2.64	-	-	-
F(3,45)	-	1.72	1.74	2.20

NOTE: 1. See Tables 1-3 for explanation of terms above.

TABLE 6: The Standard Deviation of the Growth Rate Of Industrial Production for the G7 - **Quarterly Data** 1960-1993 - 49 Observations

Equation: s.d.y. = $b_{10} + b_{11}\sigma_{\eta} + b_{12}\sigma_{\epsilon} + b_{13}DF + b_{14}DCB$ + error

		Central Bank Index		
		Parkin & Bade	**Grilli**	**Cukierman**
b_{10}	1.362** (0.216)	1.691** (0.372)	1.913** (0.427)	1.289** (0.443)
$b_{11}(\sigma_{\eta})$	0.265 (0.286)	0.249 (0.286)	0.174 (0.289)	0.283 (0.304)
$b_{12}(\sigma_{\epsilon})$	0.867** (0.400)	0.767* (0.410)	0.762* (0.401)	0.883** (0.413)
$b_{13}(DF)$	3.976** (0.790)	3.913** (0.791)	3.916** (0.781)	3.997** (0.806)
$b_{14}(DCB)$	-	-0.122 (0.112)	-0.087 (0.059)	0.149 (0.790)
$\bar{R}^2$	0.365	0.368	0.382	0.352
F(3,45)	10.21	-	-	-
F(4,44)	-	7.98	8.42	7.56

NOTE: 1. See Tables 1-3 for explanation of terms above.

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ENDNOTES

1. A third proposition concerns whether there is also a strong correlation between relative price dispersion and the variability of the aggregate inflation rate.

2. Although it has long been suspected that inflation rates and inflation uncertainty are closely linked, the econometric evidence is not particularly clear cut. The area of contention is whether high variability translates into greater uncertainty about these rates. There have been three broad categories of research:(i) Cross-country (i.e. Logue & Willet(1976) amongst many others); (ii) Cross-sectional surveys (associated with Cukierman & Wachtel);and, (iii) Time series analysis (i.e.Engle(1983)). In essence the first two support the contention, while the third rejects it. See B & C (1990) and Evans(1991) for a full description of the debate.

3. The exception being Froyen and Waud (1984, 1987) who measure the "Friedman Effect" within a natural rate model using a time varying measure of inflation variability. However, as explained later, we believe our analysis is distinct with respect to the time horizon element of inflation uncertainty.

4. This effect has been modeled by Lucas (1973) and explained, in terms of relative price variability, inflation and allocative efficiency, by Cukierman (1982).

5. Azariadis (1977) believes that this volatility effect could well be permanent. He disagrees with Friedman's idea of a natural rate theory and provides a plausible argument, via an incomplete information model, of how inflation variability can permanently increase the level of unemployment. This would translate Friedman's "decades" into the long, long run and such a "time" period would be difficult to test within the confines of our sample period (1960-1993).

6. Inflation data was contained in the standard CPI measures for the G7 (RPI for the United Kingdom).For industrial output, manufacturing or industrial production data was used. All data was obtained from the I.M.F.'s *International Financial Statistics* database.

7. To test the assumption that both permanent and temporary inflation shocks are white noise, Ball and Cecchetti compared the implied MA(1) model for $\Delta \pi_t$ with more general models. Using the Schwartz criterion in order to select between alternative ARMA(p,q) model specifications, they find that an MA(1) specification is suitable for the majority of their 40 countries (including the G7). The benefit of such a specification is the ease at which the derived variance measures drop out. Because of this, experimenting with more fully specified processes would require a far greater level of accuracy in order to overcome the computational and formal complexities implied by such an approach. For our data set, such an MA(1) formulation appeared to fit well for all of the countries bar Germany, where the

resultant lags exhibited negative serial correlation.

8. Similar results were found to those detailed below with an augmented data set using overlapping time periods with respect to the panels. This gave us 133 observations for the G7 and 19 for each country.Tables of results are available from the authors upon request.
Excluding Germany or France from the the panel does not compromise the results reported. In fact, when Germany is omitted, the results are even more supportive of the Friedman hypothesis and of the importance of central bank independence.

9. In all of the countries, the Durbin-Watson statistics indicated no existence of autocorrelation.

10. For the standard deviation in output we have utilised a dummy variable for France in order to adjust for the high degree of volatility in output around the time of the May 1968 uprisings.

11. Application of White's test at the 10% level indicated that the residuals were free of heteroscedasticity.

12. For a comprehensive analysis of such results see Johnson and Siklos (1993).

13. The dummies for the respective G7 countries are as follows: (with the highest numbers implying a greater degree of central bank independence).

	Parkin & Bade	Grilli et al.	Cukierman
Scale:	4 to 1	7 to 0	1 to 0
Germany	4	7	0.66
USA	3	7	0.51
Canada	2	7	0.46
Japan	3	5	0.16
UK	2	5	0.32
France	2	5	0.28
Italy	1/2	1	0.25

Parkin & Bade's specification refers to policy type; Grilli et al. to the degree of economic independence while Cukierman's measure is associated with overall independence (which includes a combined "score" legal independence, economic freedom, survey attitudes and soon). We did not allow for the index to change over time even though in the case of Canada and Italy the central bank achieved a greater measure of independence. As these changes occurred towards the end of the sample period, any effects attributable to the change in status should be limited.

14. We also took into account country specific factors because the cross-sectional time-series estimates obtained concern fixed effects estimators. For example we, further, investigated whether each country within the panel required a separate intercept by employing country specific dummies. The results suggested that country specific factors are of relevance. This would appear to be intuitively plausible owing to the distinct inflation experiences of the countries within our sample. Results can be obtained from the authors upon request.

15. Application of White's test indicated that the residuals were free of heteroscedasticity at the 10 percent level.

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(Continued on next page)

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