Available online at www.sciencedirect.com
ScienceDirect
Journal of Macroeconomics 28 (2006) 609–620
Journal of MACROECONOMICS
www.elsevier.com/locate/jmacro

The links between inflation, inflation uncertainty and output growth: New time series evidence from Japan

Bradley Kemp Wilson *

Department of Economics, University of Saint Thomas, Mail #4246, 2115 Summit Avenue, Saint Paul, MN 55105-1096, USA

Received 7 November 2001; accepted 3 November 2004
Available online 25 July 2006

Abstract

Employing a bivariate EGARCH-M model, we investigate the links between inflation, inflation uncertainty and output growth for post-war Japanese data. Our results indicate that increased inflation uncertainty is associated with higher average inflation and lower average growth in Japan. Further, we find that increased growth uncertainty is associated with higher average inflation, but unrelated to average growth. Lastly, both inflation and growth display significant asymmetry in their respective conditional variances. Specifically, negative surprises raise both inflation uncertainty and growth uncertainty more than positive surprises.
© 2006 Elsevier Inc. All rights reserved.

JEL classification: E31; E64

Keywords: Inflation; Inflation volatility; Output growth; EGARCH-M

1. Introduction

Monetary policy practitioners world wide have long believed that inflation is potentially detrimental to the growth of an economy's output. The basis for this conviction can be

* Tel.: +1 651 962 5688; fax: +1 651 962 5682.
E-mail address: bkwilson@stthomas.edu

0164-0704/$ - see front matter © 2006 Elsevier Inc. All rights reserved.
doi:10.1016/j.jmacro.2004.11.004

attributed to a hypothesis advanced by Milton Friedman in his 1977 Nobel Lecture (see Friedman, 1977). Comprised of two coherent arguments, Friedman's hypothesis first posits that rising inflation, brought about by full employment policy objectives, creates a strong pressure to counter it, and that the perception of such pressure subsequently increases private agent uncertainty about the course of future inflation. Second, the hypothesis contends that as a result of this increase in inflation uncertainty market prices become a less efficient system for coordinating economic activity, thus causing a decline in output growth.

Because Friedman's hypothesis posits a positive relation between inflation uncertainty and positive inflation surprises together with a negative relation between output growth and inflation uncertainty, policy makers have embraced it as a reinforceable argument for their continued adherence to policies of price stability. In this paper, using growth and inflation data from Japan, we present new statistical evidence on the links between inflation, inflation uncertainty and output growth. We find that while inflation uncertainty significantly negatively affects output growth in Japan, such uncertainty rises more in response to negative inflation surprises than to positive surprises.

Since its inception, Friedman's hypothesis has been the subject of a great number of empirical investigations. Empirical studies of the hypothesis have typically pursued one of two avenues of inquiry: (1) assessing the relationship between inflation uncertainty and the inflation rate; or (2) assessing the relationship between output growth and inflation uncertainty. Interestingly, while the evidence uncovered on the first of these two relationships strongly supports the hypothesis (see Okun, 1971; Logue and Willett, 1976; Wachtel, 1977; Carlson, 1977; Cukeirman and Wachtel, 1979; Logue and Sweeney, 1981; Ball and Cecchetti, 1990; Evans, 1991; Hess and Brunner, 1993; Grier and Perry, 1998; Wilson and Culver, 1999; Fountas, 2001; Fountas et al., 2002; Grier et al., 2004; Apergis, 2004; Kontonikas, 2004), the statistical evidence on the latter relation has been mixed, though recent GARCH studies of the relation have provided unanimous support.[1] Mullineaux (1980), Hafer (1986), Darrat and Lopez (1989), Davis and Kanago (1996), Al-Marhubi (1998), Wilson and Culver (1999), Grier and Perry (2000), Hayford (2000), Fountas et al. (2002), Grier et al. (2004), Apergis (2004) all uncover evidence supporting a negative relationship between output growth and inflation uncertainty, while Katsimbris (1985), Thornton (1988), Jansen (1989), Levine and Renelt (1992), Levine and Zervos (1993), Bohara and Sauer (1994) and Clark (1997) fail to provide such support.

The methodology adopted in this paper differs from most previous studies in three important respects. First, we recognize that in terms of its policy significance Friedman's hypothesis is not two separate hypotheses but rather one hypothesis about two coherent relationships. Given this, we follow Wilson and Culver (1999) and Grier et al. (2004) and employ a model that permits a holistic examination of the hypothesis. The model

[1] ARCH studies conducted by Engle (1983), Holland (1984), Cosiman and Jansen (1988) are dissenting studies of the relationship between inflation uncertainty and the rate of inflation. That is, they find little evidence of a link between inflation and inflation uncertainty. However, as discussed by Hess and Brunner (1993), the ARCH studies of the 1980s fail to uncover a relation between inflation uncertainty and the inflation rate due to the fact that the ARCH model is by construction incapable of assessing such a relation. In addition, while not a dissenting study per se, Caporale and Caporale (2002), employing a TARCH model, find that inflation uncertainty rises more in response to negative inflation surprises than to positive surprises.

we adopt is a bivariate exponential generalized autoregressive conditional heteroskedasticity in mean (EGARCH-M) model of output growth and inflation. Unlike other time-series models the bivariate EGARCH-M allows the relationships between inflation uncertainty and inflation surprises and growth and inflation uncertainty to be assessed together, as theoretically prescribed.

Second, the paper differs from previous time series studies in the measurement of inflation uncertainty. Although the ARCH model and its more general cousin GARCH have become rather popular choices of methodology for assessing Friedman's hypothesis (see Engle, 1983; Holland, 1984; Cosiman and Jansen, 1988; Jansen, 1989; Grier and Perry, 1998, 2000; Fountas, 2001; Fountas et al., 2002; Apergis, 2004; Kontonikas, 2004), the measures of inflation uncertainty generated by these models are by construction invariant to the direction of change in inflation (see Nelson, 1991; Hess and Brunner, 1993). Because of this the ARCH/GARCH model is not capable of providing reliable assessments of either inflation-uncertainty relations or growth-uncertainty relations. The class of ARCH model we use in this paper permits its measure of inflation uncertainty to respond to the direction of change in inflation, and thus offers an important advantage over other ARCH specifications in the assessment of the hypothesis.[2]

Lastly, following Fountas et al. (2002), and Grier et al. (2004), the bivariate nature of our model permits assessments of a number of other, related predictions about the macroeconomic effects of uncertainty. For example, with respect to the inflation effects of inflation uncertainty, Cukierman and Meltzer (1986) predict that increased inflation uncertainty raises average inflation, while Holland (1995) predicts that increased inflation uncertainty lowers average inflation. With respect to the inflation effects of growth uncertainty, Devereux (1989) predicts that increased growth uncertainty raises average inflation. And finally, with respect to the growth effects of growth uncertainty, Black (1987) predicts that increased growth uncertainty raises average growth while the prediction from the literature on irreversible investment is that increased growth uncertainty lowers average growth.[3]

The plan for this paper is as follows: In Section 2, we discuss our data, and present results from a series of preliminary diagnostic tests. In Section 3, we present the time series model adopted for this study and discuss its merits. The links between inflation, inflation uncertainty and output growth are examined in Section 4, and concluding remarks are provided in Section 5.

2. Data

To empirically assess the links between inflation, inflation uncertainty and output growth we use quarterly, post-war CPI and real (in 1995 dollars) GDP data for Japan spanning from 1957:Q4 to 2002:Q3. All data is obtained from the International Financial Statistics of the International Monetary Fund.

[2] Grier and Perry (1998) estimate an asymmetric GARCH model of inflation, specifically a GJR model, and compare it to other GARCH specifications for G7 countries. They do not find evidence of asymmetry in the conditional variances of inflation for any of the G7 countries.

[3] For a more detailed summary of these predictions see Ramey and Ramey (1995), Grier and Perry (2000) and Fountas et al. (2002).

Table 1
ADF and PP tests for unit roots

Tested series	Diagnostic tests, $T(\hat{\rho}-1)$[a]			
	ADF test for a unit root		PP test for a unit root	
	With trend	Without trend	With trend	Without trend
p	−0.675	−1.974	1.447	−1.97407
Δp	−22.238**	−14.127**	−128.297***	−106.872***
q	−2.032	−2.323	−1.967	−2.318
Δq	−61.769***	−14.771**	−214.060***	−188.779***

Notes: p denotes the logarithm of the quarterly Consumer Price Index in Japan, q denotes the logarithm of the quarterly real GDP in Japan, and Δ is the difference operator. The sample period evaluated is 1959:Q1–2002:Q3. The numerical values in this table are the computed test statistics $[T(\hat{\rho}-1)]$ for the respective diagnostic test.

[a] The critical values for the tests with a trend at the 1%, 5% and 10% levels of significance are −28.4, −21.3, and −18.0, respectively. The critical values for the tests without a trend at the 1%, 5% and 10% levels of significance are −20.3, −14.0, and −11.2, respectively. The critical values are from Hamilton (1994): Table B.5, case 2 without trend and case 4 with trend. The lag lengths used in the ADF tests, determined by the recursive t-statistic procedure suggested by Campbell and Perron (1991), are as follows: 5 for p; 4 for Δp; 0 for q; and 5 for Δq. Each PP test was performed with four lags in the spectral estimation window.

* Denotes significance at the 10% level.

** Denotes significance at the 5% level.

*** Denotes significance at the 1% level.

To begin our assessment of the aforementioned links, we first perform standard Augmented Dickey Fuller (ADF) and Phillips-Perron (PP) tests on the data to detect unit roots. The tests were performed with and without time trends and the test statistics computed are of the $T(\hat{\rho}-1)$ form. The results are reported in Table 1. For both the price series and the output series, with and without a trend, the ADF and PP tests fail to reject the null hypothesis of a unit root. For both output growth and inflation, this null hypothesis is rejected by both tests, once again with and without a trend.

Given that the unit-root test results presented in Table 1 indicate that the output and price series are individually integrated of order 1, we test for cointegration between these variables. Test results, reported in Table 2, are obtained from both the Engle and Granger approach and the Johansen approach. We execute these tests with and without a trend, and for the Engle and Granger approach report results for alternative lag lengths of the differenced residual. At all lag lengths, and with or without a trend, the Engle–Granger test fails to reject the null of no cointegration. With respect to Johansen's approach, we report the maximal eigenvalue test statistics for both a restricted constant (no time trend) and an unrestricted constant (allowing for a time trend).[4] Consistent with the results from the Engle–Granger test, we do not find evidence of cointegration: The Johansen test fails to reject the null of zero cointegrating vectors.

Lastly, we perform Ljung–Box and Lagrange Multiplier tests for serial correlation and autoregressive conditional heteroskedasticity (ARCH), respectively, in the inflation and growth series. Results from these tests are reported in Table 3. The results indicate that the data is conditionally heteroskedastic, but not serially dependent.

[4] A test statistic often reported along side the maximal eigenvalue test statistic is the trace test statistic. Due to the relatively small sample employed in this paper, however, we do not perform the trace test.

Table 2
Tests for cointegration

Lags	t-Statistic (with trend)	t-Statistic (without trend)
Engle–Granger tests[a]		
1	−3.057	−2.609
2	−3.143	−2.799
3	−2.248	−2.043
4	−2.503	−2.324
5	−2.546	−2.383
6	−2.690	−2.483
	Maximal eigenvalue test statistic (with a constant and trend)	Maximal eigenvalue test statistic (with a constant and without trend)
Johansen tests[b]		
6	13.629	27.355

Notes: The sample period evaluated is 1959:Q1–2002:Q3.

[a] The Engle–Granger tests conducted here are tests for cointegration between the logarithms of the CPI and real GDP in Japan. The tests were conducted with a trend and without a trend. Lags denotes the lag length of the differenced residual in each test. The numerical values in this table are the test statistics of the respective t-test. The null hypothesis of the Engle–Granger test is that there is no cointegration. For the test including a trend, the critical value at the 10% level of significance is −3.537. For the test without a trend, the critical value at the 10% level of significance is −3.070. Critical values are from MacKinnon (1991). *Denotes significance at the 10% level. **Denotes significance at the 5% level. ***Denotes significance at the 1% level.

[b] The Johansen tests conducted here are tests for cointegration between the logarithms of the CPI and real GDP in Japan. The tests were conducted with a trend and without a trend. Both tests were conducted with a constant term. Lag denotes the lag length of the undifferenced VAR used in the tests. A series of Likelihood ratio tests were performed to determine this lag length. The maximum lag length considered was 24. The numerical values in this table are the maximal eigenvalue test statistics. The null hypothesis of the Johansen tests is that $r = 0$, where r denotes the rank of cointegration. At the 5% level of significance, the critical values for the test with a constant and trend and for the test with a constant and no trend are 42.48 and 39.37, respectively. Critical values are from Osterwald-Lenum (1992). **Denotes significance at the 5% level.

Table 3
Preliminary tests for serial correlation and autoregressive conditional heteroskedasticity (ARCH)

Tested series	Q(4)	Q(8)	LM(4)
Δp	1.632	9.406	28.834
	[0.803]	[0.309]	[0.000]
Δq	1.944	6.168	21.396
	[0.746]	[0.628]	[0.000]

Notes: p denotes the logarithm of the quarterly CPI in Japan, q denotes the logarithm of quarterly real GDP in Japan, and Δ is the difference operator. Q(4) and Q(8) are the Ljung–Box statistics for fourth and eighth order serial correlation in the residuals from an AR(4) for Δp and an AR(5) for Δq. LM(4) are Lagrange Multiplier test statistics with four lags in the respective squared residual. The sample period evaluated is 1959:Q1–2002:Q3. The numbers in [ ] are p-values.

3. Empirical methodology

Since the pioneering work of Engle (1982), the ARCH class of time series models, specifically the GARCH model, has become a popular choice of methodology for assessing

Friedman's Hypothesis. Its popularity can be attributed to its ability to generate direct, time-varying measures of inflation uncertainty; measures implied by Friedman (1977) as necessary for accurate empirical examinations of his hypothesis.[5] To illustrate the usefulness of this approach, consider the following GARCH model of inflation with an *m*th-order autoregressive process for the conditional mean:

$$\Delta p_t = \mu + \sum_{i=1}^{m} \beta_i \Delta p_{t-i} + \varepsilon_t, \quad \varepsilon_t \sim N(0, h_t), \tag{1}$$

$$h_t = h + \sum_{i=1}^{r} \theta_i h_{t-i} + \sum_{j=1}^{s} \lambda_j \varepsilon_{t-j}^2, \tag{2}$$

where Δp_t denotes the rate of inflation between $t-1$ and t, and ε_t is the shock to inflation at time t. As indicated in Eqs. (1) and (2), ε_t is usually assumed to be normally distributed with a time-varying conditional variance h_t, specified as a linear function of past squared inflation surprises (the simple ARCH model) and past variances (the GARCH model). Now it is this conditional variance which is of interest, for the estimated sequence of these variances constitutes a direct, time-varying measure of inflation uncertainty.[6] Together with series for the expected inflation rate and output growth, it follows that this sequence of estimated conditional variances can be used to perform separate time series analyses of the relationships between inflation uncertainty and inflation and output growth and inflation uncertainty.[7]

There is, however, a serious drawback to using the ARCH/GARCH model to generate time-varying measures of inflation uncertainty for testing Friedman's hypothesis. As illustrated in Eq. (2), the conditional variance of inflation estimated by the GARCH model is a function only of the magnitudes of inflation shocks (ε_{t-i}^2), not also the signs of such shocks, and thus by construction is blind as to whether inflation is rising or falling. It follows then that the GARCH model does not yield a time-varying measure of inflation uncertainty capable of providing a reliable assessment of the response of inflation uncertainty to inflation surprises.

To remedy this shortcoming, we adopt a class of ARCH model, exponential GARCH (EGARCH), which explicitly allows its measure of uncertainty to respond to the positivity or negativity of shocks.[8] In so doing, this model permits a straightforward assessment of the response of inflation uncertainty to inflation surprises. To see this, replace Eq. (2) with the following expression to obtain an EGARCH model of inflation:

$$h_t = \exp\left(h + \sum_{i=1}^{r} \theta_j \log h_{t-i} + \sum_{j=1}^{s} \lambda_j \{|v_{t-j}| - E|v_{t-j}| + \xi v_{t-j}\}\right), \tag{3}$$

[5] See Friedman (1977), second full paragraph on page 468.

[6] All of this assumes, of course, that the inflation series possesses heteroskedastic errors, which has been well documented.

[7] Time series assessments of the inflation-uncertainty relation have involved, for example, Granger methods to test for causality between the conditional variance of inflation and expected inflation, and comparing the path of the conditional variance with the path of expected inflation. Assessments of the growth-uncertainty relation have typically involved regressing output growth on the conditional variance of inflation.

[8] Exponential GARCH (EGARCH) was developed by Nelson (1991) to remedy shortcomings of the GARCH model, for example the inability of GARCH to generate a time-varying variance for a series that is capable of responding to the series mean; a relationship observed in both economic and financial time series.

where E is the mathematical expectation, and $v_{t-j}(=\varepsilon_{t-j}/\sqrt{h_{t-j}})$ is i.i.d. with zero mean and unit variance. Like the GARCH model of inflation, the EGARCH model yields a time-varying measure of inflation uncertainty that responds to the magnitudes of inflation shocks. For example, in Eq. (3), if $\lambda_j > 0$, then a deviation of $|v_{t-j}|$ from its expected value causes inflation uncertainty (h_t) to rise. Unlike the GARCH model, however, EGARCH allows this effect to be asymmetric, thus permitting its measure of uncertainty to respond also to the signs of inflation shocks. The parameter in Eq. (3) that allows for such asymmetry is ξ. If $\xi > 0$, then h_t will rise more in response to positive inflation shocks ($\varepsilon_{t-j} > 0$) than to negative shocks ($\varepsilon_{t-j} < 0$). If $\xi < 0$, then h_t will rise more in response to negative inflation shocks than to positive shocks. And if $\xi = 0$, then a positive shock to inflation has the same effect on uncertainty as a negative shock of the same magnitude. In this case, the direction of change in inflation does not influence the path of inflation uncertainty.

It is important to note that other asymmetric GARCH models have been considered in the literature for assessing asymmetry in the conditional variance of inflation, most notably the GJR model proposed by Glosten et al. (1993) (see Grier and Perry, 1998). While the GJR model is certainly less computationally burdensome than exponential GARCH, it is subject to some of the same limitations as GARCH models, namely nonnegativity constraints on its parameters, and thus in general cannot guarantee positive values for variances (see Nelson, 1991).

Now because we are interested in assessing the relationship between inflation uncertainty and inflation together with the relationship between growth and inflation uncertainty, we specify a bivariate EGARCH-M model of inflation and output growth.[9] The bivariate framework of our model dictates that we specify whole covariance matrices which change over time. Generally, such specifications will involve the sacrifice of a large number of degrees-of-freedom, and no guarantee of numerical tractability. To avoid this, we adopt Bollerslev's (1990) specification and assume conditional correlations to be constant so that all time variations in conditional covariance terms are due only to changes in the conditional variances. Given the diagnostic test results reported in Tables 1–3, the model we propose is as follows:

$$\Delta p_t = \mu_p + \sum_{i=1}^{m} \alpha_{p,i}\Delta p_{t-i} + \sum_{j=1}^{n} \beta_{p,j}\Delta q_{t-j} + \gamma_p h_{p,t}^{.5} + \varphi_p h_{q,t}^{.5} + \varepsilon_{p,t}, \tag{4}$$

$$\Delta q_t = \mu_q + \sum_{i=1}^{m'} \alpha_{q,i}\Delta q_{t-i} + \sum_{j=1}^{n'} \beta_{q,j}\Delta p_{t-j} + \gamma_q h_{q,t}^{.5} + \varphi_q h_{p,t}^{.5} + \varepsilon_{q,t}, \tag{5}$$

$$h_{p,t} = \exp\left(h_p + \sum_{i=1}^{r} \theta_{p,j}\log h_{p,t-i} + \sum_{j=1}^{s} \lambda_{p,j}\{|v_{p,t-j}| - E|v_{p,t-j}| + \xi_p v_{p,t-j}\}\right), \tag{6}$$

$$h_{q,t} = \exp\left(h_q + \sum_{i=1}^{r'} \theta_{q,j}\log h_{q,t-i} + \sum_{j=1}^{s'} \lambda_{q,j}\{|v_{q,t-j}| - E|v_{q,t-j}| + \xi_q v_{q,t-j}\}\right), \tag{7}$$

$$h_{pq,t} = \rho\sqrt{h_{p,t}\cdot h_{q,t}}, \tag{8}$$

[9] Developed by Engle et al. (1987), GARCH-M, a simple extension of the GARCH model, allows the conditional variance of a series to feed back and influence the conditional mean.

where, as above, Δp_t denotes the rate of inflation between $t-1$ and t; Δq_t denotes the growth rate of output between $t-1$ and t; $\varepsilon_{p,t}$ is the shock to inflation at t; $\varepsilon_{q,t}$ is the shock to output growth at t; $v_{i,t-j}(=\varepsilon_{i,t-j}/\sqrt{h_{i,t-j}})$ is i.i.d. with zero mean and unit variance for $i=p,q$; $h_{i,t}=\mathrm{var}_{t-1}(\varepsilon_{i,t})$ for $i=p,q$; and $h_{pq,t}=\mathrm{cov}_{t-1}(\varepsilon_{p,t},\varepsilon_{q,t})$. Eqs. (4) and (5) describe changes in means while Eqs. (6)–(8) specify time variations in the conditional second moments: Eqs. (6) and (7) describe changes in the conditional variances, while (8), consistent with the constant correlation specification of Bollerslev (1990), describes changes in the conditional covariance. Generally, $(\varepsilon_{p,t},\varepsilon_{q,t})$ is assumed to be distributed bivariately normal or student-t. The normality assumption is adopted in this study.

In essence, Eqs. (4)–(8) constitute a VAR with exogenous variables, simply modified from the standard to allow for heteroskedastic errors.[10] The exogenous variables in the model, that is $h_{p,t}^{.5}$ and $h_{q,t}^{.5}$ in Eqs. (4) and (5), are incorporated to assess the responses of inflation and output growth to both inflation uncertainty ($h_{p,t}^{.5}$) and growth uncertainty ($h_{q,t}^{.5}$). Of primary interest in the above model are the coefficients on $h_{p,t}^{.5}$ in Eq. (5) and $v_{p,t-j}$ in Eq. (6), that is φ_q and ξ_p, respectively. The coefficient φ_q determines the response of output growth to inflation uncertainty, while ξ_p, like ξ in Eq. (3), determines the response of inflation uncertainty to inflation surprises. In addition, the coefficients on $h_{p,t}^{.5}$ and $h_{q,t}^{.5}$ in Eq. (4) permit assessments of Cukierman and Meltzer's (1986) hypothesis and Devereux's (1989) hypothesis, respectively, while the coefficient on $h_{q,t}^{.5}$ in Eq. (5) permits an assessment of Black's (1987) hypothesis. Combined, these coefficients constitute a holistic assessment of the links between inflation, inflation uncertainty and growth.

4. Estimation and results

To empirically evaluate the bivariate EGARCH-M model proposed in this paper, we employ Maximum Likelihood Estimation (MLE) using the Berndt, Hall, Hall, and Hausman (BHHH) algorithm. First, a series of Likelihood Ratio (LR) tests are performed to determine the lag structures of Eqs. (4)–(7). The maximum lag length considered in the execution of these tests is 24. Due to the large number of LR tests performed to generate the final estimated model, results from these tests will be provided upon request. After allowing for the construction of lags, the time period evaluated is 1959:Q1–2002:Q3. Results from the final model are reported in Table 4.

As these results show, the constant term and the coefficients on Δp_{t-1}, Δp_{t-2}, Δp_{t-4}, Δq_{t-1}, $h_{p,t}^{.5}$, and $h_{p,t}^{.5}$ in the conditional mean equation for inflation are all individually significant at the 5% level, whereas in the conditional mean equation for growth the constant term and the coefficients on Δq_{t-1}, Δq_{t-3}, Δq_{t-5}, and $h_{p,t}^{.5}$ are individually significant at the 5% level. The statistical significance of Δq_{t-1} in the mean equation for inflation together with the absence of lagged inflation terms in the mean equation for growth indicate unidirectional Granger causality from growth to inflation in Japan, and the battery of LR tests used to arrive at this

[10] To control for the growth effects of oil price shocks, Davis and Kanago (1996), Hayford (2000) include variables in their models to account for growth in the relative price of crude oil. We constructed such a variable for Japan, but, consistent with Hayford (2000), its estimated coefficient was found to be statistically insignificant with a p-value near 0.9. Given this finding, and given that we needed to decrease our sample size in order to assess this effect, we have excluded this variable from our formal analysis. In addition, and unlike Grier and Perry (2000), we were unable to consider relevant interest rate spread variables in our analysis due to an insufficient number of observations for such variables.

Table 4
Estimates of the bivariate EGARCH-M model of Section 3

Conditional mean equations (4) and (5)

(4) $$\Delta p_t = \underset{(-7.888)}{-0.009} + \underset{(3.641)}{0.092}\,\Delta p_{t-1} + \underset{(10.337)}{0.285}\,\Delta p_{t-2} - \underset{(-0.452)}{0.013}\,\Delta p_{t-3} + \underset{(16.014)}{0.298}\,\Delta p_{t-4} + \underset{(4.018)}{0.062}\,\Delta q_{t-1} + \underset{(8.342)}{1.610}\,h^{.5}_{p,t} + \underset{(3.224)}{0.020}\,h^{.5}_{q,t}$$

(5) $$\Delta q_t = \underset{(2.327)}{0.008} + \underset{(1.901)}{0.137}\,\Delta q_{t-1} + \underset{(0.529)}{0.037}\,\Delta q_{t-2} - \underset{(2.700)}{0.161}\,\Delta q_{t-3} + \underset{(1.682)}{0.108}\,\Delta q_{t-4} + \underset{(3.840)}{0.195}\,\Delta q_{t-5} + \underset{(0.852)}{0.052}\,\Delta p_{t-1} - \underset{(-0.042)}{0.018}\,h^{.5}_{q,t} - \underset{(-1.930)}{0.424}\,h^{.5}_{p,t}$$

Conditional variance/covariance equations (6)–(8)

(6) $$h_{p,t} = \exp\Big(-\underset{(-80.086)}{11.095} - \underset{(-7.648)}{0.115}\log h_{p,t-1} + \underset{(7.435)}{1.261}\Big\{|v_{p,t-1}| - E|v_{p,t-1}| - \underset{(-2.168)}{0.149}\,v_{p,t-1}\Big\} - \underset{(-2.621)}{0.266}\Big\{|v_{p,t-2}| - E|v_{p,t-2}| - \underset{(-2.168)}{0.149}\,v_{p,t-2}\Big\}$$
$$- \underset{(-2.657)}{0.240}\Big\{|v_{p,t-3}| - E|v_{p,t-3}| - \underset{(-2.168)}{0.149}\,v_{p,t-3}\Big\} + \underset{(4.566)}{0.346}\Big\{|v_{p,t-4}| - E|v_{p,t-4}| - \underset{(-2.168)}{0.149}\,v_{p,t-4}\Big\}\Big)$$

(7) $$h_{q,t} = \exp\Big(-\underset{(-6.049)}{3.206} + \underset{(10.769)}{0.644}\log h_{q,t-1} + \underset{(0.161)}{0.019}\Big\{|v_{q,t-1}| - E|v_{q,t-1}| - \underset{(-2.023)}{0.806}\,v_{q,t-1}\Big\} - \underset{(-2.121)}{0.347}\Big\{|v_{q,t-2}| - E|v_{q,t-2}| - \underset{(-2.023)}{0.806}\,v_{q,t-2}\Big\}$$
$$+ \underset{(1.280)}{0.184}\Big\{|v_{q,t-3}| - E|v_{q,t-3}| - \underset{(-2.023)}{0.806}\,v_{q,t-3}\Big\} - \underset{(-1.559)}{0.265}\Big\{|v_{q,t-4}| - E|v_{q,t-4}| - \underset{(-2.023)}{0.806}\,v_{q,t-4}\Big\}\Big)$$

(8) $$h_{p,q,t} = -\underset{(-0.631)}{0.055}\sqrt{h_{p,t}\cdot h_{q,t}}$$

Standardized residual diagnostics

	Mean	Variance	Q(4)	Q(8)
$v_{p,t}$	−0.032	1.049	5.995	12.024
	[0.676]	[0.310]	[0.199]	[0.150]
$v_{q,t}$	−0.096	1.093	4.736	9.624
	[0.214]	[0.186]	[0.315]	[0.292]

This table displays parameter estimates for the bivariate EGARCH-M model presented in Section 3. p denotes the logarithm of the quarterly CPI in Japan; q denotes the logarithm of quarterly real GDP in Japan; Δ is the difference operator; $h_{p,t}$ is the conditional variance of inflation; $h_{q,t}$ is the conditional variance of output growth; and $h_{pq,t}$ is the conditional covariance between inflation and growth. The sample period evaluated is 1959:Q1–2002:Q3. The numbers in ( ) are asymptotic t-statistics; the numbers in [ ] are p-values; and Q(4) and Q(8) are the Ljung–Box statistics for fourth and eighth order serial correlation in the standardized residuals. With respect to the mean and variance reported in the lower table, the null hypothesis is that such are equal to 0 and 1, respectively.

model support this finding. This result, however, runs contrary to that found by Fountas et al. (2002) for industrial production growth and producer price index inflation in Japan.

Consistent with the Lagrange Multiplier test results reported in Table 3, the results for the variance equations demonstrate that the variances of both inflation and output growth are time varying, display asymmetry, and exhibit statistically significant GARCH terms. Of greatest note are the estimated asymmetry effects. The coefficients on $v_{p,t-i}$ and $v_{q,t-i}$ in the conditional variance equations for inflation and growth, respectively, are both negative and statistically significant at the 5% level. In addition, the constant correlation coefficient in the conditional covariance equation is statistically insignificantly different from zero. Overall, the model appears to be well specified. The standardized residuals, $v_{p,t}$ and $v_{q,t}$, both possess means and variances that are statistically insignificantly different from 0 and 1, respectively, as required, and both satisfy the nulls of no fourth and eighth order serial dependence, as indicated by Ljung–Box tests.

The economic significance of our results are as follows. In the conditional mean equation for inflation, the results provide support for both Cukierman and Meltzer's (1986) prediction and Devereux's (1989) prediction: The estimated coefficients on $h_{p,t}^{.5}$ and $h_{q,t}^{.5}$ indicate that both increased inflation uncertainty and increased growth uncertainty raise the mean rate of inflation in Japan, respectively. Both of these findings run contrary to Fountas et al. (2002). In their study of the Japanese economy, using industrial production and the PPI, they find that increased inflation uncertainty lowers average inflation and increased growth uncertainty has no effect on average inflation.

In the conditional mean equation for output growth, the parameter estimate on $h_{q,t}^{.5}$ is negative and statistically insignificant, indicating that increased growth uncertainty has no effect on average output growth in Japan. In other words, consistent with Fountas et al. (2002), we find no evidence in support of Black's (1987) prediction in Japan. We do, however, find evidence in support of Friedman's (1977) hypothesis, the key findings of this paper. The estimated coefficient on $h_{p,t}^{.5}$ in the conditional mean equation for growth indicates that increased inflation uncertainty lowers average output growth in Japan. Interestingly, however, the estimated coefficient on the asymmetry term $\mathrm{v}_{p,t-i}$ in the conditional variance equation for inflation indicates that inflation uncertainty in Japan rises more in response to negative inflation surprises than to positive surprises. Caporale and Caporale (2002), employing a TARCH model, find a similar effect in US inflation data. Finally, the estimated coefficient on the asymmetry term $\mathrm{v}_{q,t-j}$ in the conditional variance equation for growth indicates that growth uncertainty in Japan rises more in response to negative growth shocks than to positive shocks.

5. Conclusion

In this study, we have constructed a bivariate EGARCH-M model of Japanese inflation and growth to examine the links between inflation, inflation uncertainty and growth. We find strong evidence for the predictions that increased inflation uncertainty raises average inflation and lowers average growth. In addition, while we fail to uncover a statistically significant relationship between growth uncertainty and average growth, we do find evidence for the prediction that increased growth uncertainty raises average inflation. Lastly, unlike that typically found in US data, we find that inflation uncertainty in Japan rises more in response to negative inflation surprises than to positive surprises. Together, these findings support price stability as a fundamental objective of central banks.

Acknowledgements

I would like to thank Sarah E. Culver, Seung Dong Lee, Stephen M. Miller, Baldev Raj, and two anonymous referees for helpful comments.

References

Al-Marhubi, F., 1998. Cross-country evidence on the link between inflation volatility and growth. Applied Economics 30, 1317–1326.

Apergis, N., 2004. Inflation, output growth, volatility and causality: Evidence from panel data and the G7 countries. Economics Letters 83, 185–191.

Ball, L., Cecchetti, S.G., 1990. Inflation and uncertainty at short and long horizons. Brookings Papers on Economic Activity 1, 215–254.

Black, F., 1987. Business Cycles and Equilibrium. Basil Blackwell, New York.

Bohara, A.K., Sauer, C., 1994. The role of inflation uncertainty in Germany: Friedman's hypothesis revisited. Empirical Economics 19, 611–627.

Bollerslev, T., 1990. Modeling the coherence in short-run nominal exchange rates: A multivariate generalized ARCH model. Review of Economics and Statistics 72, 498–505.

Campbell, J., Perron, P., 1991. Pitfalls and opportunities: What macroeconomists should know about unit roots, NBER Macroeconomics Annual. MIT Press, Cambridge, MA.

Caporale, B., Caporale, T., 2002. Asymmetric effects of inflation shocks on inflation uncertainty. Atlantic Economic Journal 30, 384–387.

Carlson, J.A., 1977. A study of price forecasts. Annals of Economic and Social Measurement 6, 27–56.

Clark, T.E., 1997. Cross-country evidence on long-run growth and inflation. Economic Inquiry 35, 70–81.

Cosiman, T.T., Jansen, D.W., 1988. Estimates of the variance of US inflation based upon the ARCH model. Journal of Money, Credit and Banking 20, 409–421.

Cukeirman, A., Wachtel, P., 1979. Differential inflationary expectations and the variability of the rate of inflation: Theory and evidence. American Economic Review 69, 595–609.

Cukierman, A., Meltzer, A., 1986. A theory of ambiguity, credibility, and inflation under discretion and asymmetric information. Econometrica 54, 1099–1128.

Darrat, A.F., Lopez, F.A., 1989. Has inflation uncertainty hampered economic growth in Latin America? International Economic Journal 3, 1–14.

Davis, G., Kanago, B., 1996. On measuring the effect of inflation uncertainty on real GNP growth. Oxford Economic Paper 48, 163–175.

Devereux, M., 1989. A positive theory of inflation and inflation variance. Economic Inquiry 27, 105–116.

Engle, R.F., 1982. Autoregressive conditional heteroskedasticity with estimates of the variance of United Kingdom inflation. Econometrica 50, 987–1007.

Engle, R.F., 1983. Estimates of the variance of US inflation based upon the ARCH model. Journal of Money, Credit and Banking 15, 286–301.

Engle, R.F., Lilien, D.M., Robins, R.P., 1987. Estimating time varying risk premia in the term structure: The ARCH-M model. Econometrica 55, 391–407.

Evans, M., 1991. Discovering the link between inflation rates and inflation uncertainty. Journal of Money Credit and Banking 23, 169–184.

Fountas, S., 2001. The relationship between inflation and inflation uncertainty in the UK: 1885–1998. Economics Letters 74, 77–83.

Fountas, S., Karanasos, M., Kim, J., 2002. Inflation and output growth uncertainty and their relationship with inflation and output growth. Economics Letters 75, 293–301.

Friedman, M., 1977. Nobel lecture: Inflation and unemployment. Journal of Political Economy 85, 451–472.

Glosten, L.R., Jagannathan, R., Runkle, D.E., 1993. On the relation between the expected value and the volatility of the nominal excess return on stocks. Journal of Finance 48, 1779–1801.

Grier, K., Perry, M.J., 1998. On inflation and inflation uncertainty in the G-7 countries. Journal of International Money and Finance 17, 671–689.

Grier, K., Perry, M.J., 2000. The effects of real and nominal uncertainty of inflation and output growth: Some GARCH-M evidence. Journal of Applied Econometrics 15, 45–58.

Grier, K., Henry, Ó.T., Olekalns, N., Shields, K., 2004. The asymmetric effects of uncertainty on inflation and output growth. Journal of Applied Econometrics 19, 551–565.
Hafer, R.W., 1986. Inflation uncertainty and a test of the Friedman hypothesis. Journal of Macroeconomics 8, 365–372.
Hamilton, J.D., 1994. Time Series Analysis. Princeton University Press, Princeton, NJ.
Hayford, M.D., 2000. Inflation uncertainty, unemployment uncertainty and economic activity. Journal of Macroeconomics 22, 315–329.
Hess, G.D., Brunner, A.D., 1993. Are higher levels of inflation less predictable? A state dependent conditional heteroskedasticity approach. Journal of Business and Economic Statistics 11, 187–197.
Holland, S.A., 1984. Does higher inflation lead to more inflation uncertainty? Federal Reserve Bank of St. Louis Economic Review February, 15–26.
Holland, S.A., 1995. Inflation and uncertainty: Tests for temporal ordering. Journal of Money, Credit, and Banking 27, 827–837.
Jansen, D.W., 1989. Does inflation uncertainty affect output growth. Federal Reserve Bank of St. Louis Economic Review July/August, 43–54.
Katsimbris, G.M., 1985. The relationship between the inflation rate, its variability, and output growth variability: Disaggregated international evidence. Journal of Money, Credit and Banking 17, 179–188.
Kontonikas, A., 2004. Inflation and inflation uncertainty in the United Kingdom: Evidence from GARCH modeling. Economic Modeling 21, 525–543.
Levine, R., Renelt, D., 1992. A sensitivity analysis of cross-country growth regressions. American Economic Review 82, 942–963.
Levine, R., Zervos, S., 1993. What have we learned about policy and growth from cross-country analysis. American Economic Review Papers and Proceedings 83, 426–430.
Logue, D.E., Sweeney, R.J., 1981. Inflation and real growth: Some empirical results. Journal of Money, Credit and Banking 13, 497–501.
Logue, D.E., Willett, T.D., 1976. A note on the relation between the rate and variability of inflation. Econometrica 43, 151–158.
MacKinnon, J.G., 1991. Critical values for cointegration tests. In: Engle, R.F., Granger, C.W.J. (Eds.), Long-Run Economic Relationships: Readings in Cointegration, Advanced Texts in Econometrics. Oxford University Press, Oxford, New York, Toronto, Melbourne, pp. 267–276.
Mullineaux, D.J., 1980. Unemployment, industrial production, and inflation uncertainty in the United States. Review of Economics and Statistics 62, 163–169.
Nelson, D.B., 1991. Conditional heteroskedasticity in asset returns: a new approach. Econometrica 59, 347–370.
Okun, A., 1971. The mirage of steady inflation. Brookings Papers on Economic Activity 2, 485–498.
Osterwald-Lenum, M., 1992. A note with quantiles of the asymptotic distribution of the maximum likelihood cointegration rank test statistics. Oxford Bulletin of Economics and Statistics 54, 461–472.
Ramey, G., Ramey, V.A., 1995. Cross-country evidence on the link between volatility and growth. American Economic Review 85, 1138–1151.
Thornton, J., 1988. Inflation and output growth: Some time series evidence, a note. American Economist 32, 55–58.
Wachtel, P., 1977. Survey measures of expected inflation and their potential usefulness. In: Popkin, J. (Ed.), Analysis of Inflation 1965–1974. Ballinger, Cambridge, MA.
Wilson, B.K., Culver, S.E., 1999. On measuring the response of real GDP growth to changes in inflation volatility. Quarterly Journal of Business and Economics 38, 3–15.