Do Asymmetric Central Bank Preferences Help Explain Observed Inflation Outcomes?

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Abstract

When the central banker's loss function is asymmetric, changes in the volatility of inflation and/or unemployment affect equilibrium inflation. This suggests that changing macroeconomic volatilities may be an important driving force behind trends in observed inflation. Previous evidence, which has offered support for this idea, suffers from a spurious regression problem. Once this problem is controlled for, the evidence suggests that the volatility of unemployment does not help explain inflation outcomes. There is some evidence of a relationship between inflation and its volatility, but overall the data does not support the view that changing economic volatility, as filtered through asymmetric central bank preferences, is an important driver of inflation trends.

KEYWORDS: Inflation, Monetary Policy, Asymmetric Loss Function.
JEL CLASSIFICATION: E50, E61

1 Introduction

It is well known that U.S. inflation was low in the early 1960s, rose through the late 1960s and 1970s before falling through the 1980s, and remaining low thereafter. A similar pattern of rising then falling inflation occurred in many other OECD countries.[1] What caused this rise and fall of inflation is an open question that has attracted a great deal of recent attention. While the literature to date has largely focused on the U.S. experience, the existence of a common pattern in inflation suggests that any successful explanation ought to be robust across OECD countries.

In this paper we ask to what extent the common observed inflation trend in OECD countries is the result of the interaction of time inconsistency problems in monetary policy interacting with changes in the volatility of shocks to inflation and/or unemployment. This is a promising candidate explanation of inflation trends for a number of reasons. First, it is well known that the degree of macroeconomic volatility has fallen along with the level of inflation in many OECD countries in recent years.[2] Thus the decline in OECD inflation rates is roughly coincident with the so-called Great Moderation. Second, standard time inconsistency models of monetary policy, extended to allow for asymmetric central bank preferences, provide a clear theoretical channel through which these factors would affect trend inflation. Since these models abstract away from country specific institutional details, the theory naturally extends from the U.S. case to the cross country setting. Third, both the comments of policy insiders[3] as well as formal empirical work on monetary policy reaction functions[4] support the key feature of these models: that central bank preferences are asymmetric. Finally, existing research explicitly examining the relationship between asymmetric central bank preferences and inflation outcomes appears to support the hypothesis.[5]

We begin the paper by documenting the existence of a common trend in the inflation rates of OECD other than the U.S.. Figures 1-12 plot inflation rates, measured by annualized quarterly percentage changes in the Consumer Price Index, for 12 OECD countries. For

[1] See Rogoff (2003), Cicarelli & Mojon (2008), and Doyle & Falk (2008).

[2] Blanchard & Simon (2000), Stock & Watson (2003)

[3] Blinder (1997, 1998) and Goodhart (1998).

[4] Reduced form tests for asymmetries in policy reaction functions include Bec, Ben Salem, & Collard (2002), Kim, Osborn, & Sensier (2002), Martin & Milas (2004), Karagedikli & Lees (2004), and Bruinshoofd & Candelon (2005). Tests based on more structural models include Surico (2004, 2003), Dolado, Maria-Dolores, & Naviera (2005), Dolada, Maria-Dolores, & Ruge-Murcia (2004), and Aguiar & Martins (2008).

[5] Ruge-Murcia (2004), Surico (2006).

comparison, a 2-year centered moving average of U.S. inflation is included on each plot (the dashed red line) A common pattern is visible in the raw data (the light blue line), but more transparent in the 2-year centered moving averages also displayed in the figures (the heavy blue line), and is as follows: inflation starts out low in the early 1960s in most countries. This is followed by a period of rising inflation lasting until the late 1970s or early 1980s in all countries except Germany and Japan (where inflation peaks in the early and mid 1970s respectively). After this period of rising inflation, inflation rates then fall until the present, and are generally as low or lower by the end of the 1990s than they were in the early 1960s. The commonality of inflation outcomes over the past four decades suggests that a successful explanation of long run inflation trends ought to be applicable across OECD countries.

Theories of time inconsistent monetary policy based on asymmetric central bank preferences are a plausible candidate explanation of this common trend as these models are general enough to encompass the differing institutional arrangements across OECD countries. While early versions of these models required the policy makers target an unattainable unemployment rate, recent theoretical innovations show that monetary policy may suffer from time inconsistency even when central bankers target the NAIRU. When central banks have asymmetric preferences, policy makers care about the *sign* as well as the magnitude of deviations of unemployment and inflation from target. In this case, monetary policy suffers from a time inconsistency problem which causes equilibrium inflation rates to depend on the variance of the shocks to inflation and unemployment.[6]

Consider, by way of illustration, a policy maker who dislikes above NAIRU unemployment more than below NAIRU unemployment, and suppose that the variance of shocks to unemployment increases. With a higher variance, the probability of an episode of very high (and, due to the asymmetry of preferences, strongly disliked) unemployment increases. The central banker will respond with expansionary monetary policy, in an attempt to drive the average unemployment rate down, to reduce the likelihood of an episode of very high unemployment. This policy, however, results in an increase in equilibrium inflation. Given the asymmetry in the loss function, the policy maker would be willing to pay this price to avoid a more distasteful episode of very high unemployment.[7]

[6] Cukierman (1999), Ruge-Murcia(2003a), Gerlach (2003), and Nobay & Peel (2003).

[7] With rational expectations, of course, this policy response is anticipated by agents, and produces the rise in inflation with no offsetting change in average unemployment.

A natural test of this theory is to use a GARCH model to estimate the conditional variance of shocks to unemployment and/or inflation and then regress inflation on this conditional variance to measure any correlation. The results of this simple exercise support the proposition that asymmetric preferences and changing volatility can explain inflation trends in at least some OECD countries.[8] We argue that findings, based on this test, that changes in the conditional volatility of unemployment have statistically significant effects on changes in inflation are likely spurious, as these results are found in those countries for which the conditional variance is most persistent. We employ Monte Carlo methods to show that, when inflation is persistent, as it is in most OECD countries, simple regressions of inflation on the conditional volatility of unemployment over-reject a correct null hypothesis of no relationship if the conditional volatility of unemployment also exhibits persistence.

To fix this problem, we re-do the analysis two ways: first by estimating the model in differences of inflation and the conditional variance of unemployment, and second by estimating the model as a cointegrating relationship. The results from these exercises suggest that changes in the conditional variance of unemployment do not explain changes in inflation. When estimating the model in differences, the model fits the data only for Austria. The cointegration results are marginally more positive, as we find support for the model in three countries out of a sample of 13. To control for the possibility of changes in central banks' inflation targets, we repeat the exercise for a more recent subsample of data, where a constant inflation target is a more plausible assumption. This does not change the results. Overall the implication is that the combination of asymmetric central bank preferences and changes in the volatility of unemployment is not a promising explanation of time series inflation trends in OECD countries.

We then examine the relationship between inflation and its conditional variance in time series data for OECD countries to allow for the possibility that central bank preferences are asymmetric in inflation, rather than unemployment. Again we estimate the model using both first differencing and cointegration tests to reduce the likelihood of spurious results.[9] The results here are more mixed than the unemployment results. Overall there is some evidence of a relationship between inflation and its own conditional variance. It is less clear

[8] Ruge-Murcia (2004).

[9] Our econometric work follows the model by assuming that the conditional variance is predetermined relative to the level of the inflation rate, though the direction of causality could go in the opposite direction (see, for example, Friedman (1977)).

whether this relationship is due to asymmetric central bank preferences. When we estimate the model in differences, there is statistically significant evidence of an inverse relationship between inflation and its conditional variance, as would be implied if asymmetric central bank preferences were important, in about one third of the countries. When we estimate the model as a cointegrating relationship, however, the sign of the correlation between inflation and its conditional variance becomes positive in all but one country, which is less attractive from the perspective of the theory. Since differencing the data emphasizes higher frequency movements, rather than the low frequency movements we are most interested in, we interpret these results to mean that the theoretical channel by which inflation volatility affects inflation through asymmetric central bank preferences is not the correct explanation for the co-movement of inflation and its volatility at co-movements.

Our conclusion is that, while it appears at first to be a promising candidate, a careful look at the data suggests that the time inconsistency story coupled with asymmetric central bank preferences and time varying variance of economic shocks is not a likely cause of the rise and fall of inflation in OECD countries in recent years.

Our paper is most closely related to papers by Ruge-Murcia (2004) and Surico (2006) who investigate the possibility that economic volatility coupled with asymmetric central bank preferences interact to explain inflation trends. Ruge-Murcia has much the same approach as take here, though he does not control for spurious results generated by persistence in inflation and the conditional variances. Surico takes a different approach to the question, estimating the degree of asymmetry in preferences from the policy reaction function for the U.S. Federal Reserve in different periods. His research question differs from ours in that he investigates the possible effect of changes in the degree of asymmetry in preferences, rather than changes in the variances of the shocks, on inflation and finds that such changes may account for a sizable fraction of the decline in U.S. inflation. Furthermore, he only considers U.S. inflation.

More broadly, our paper fits into the literature investigating the causes of the rise and fall of U.S. inflation. Loosely speaking, this literature bifurcates around the question of whether the Great Inflation was caused by monetary policy errors[10] or by adverse shocks.[11] The monetary mistakes literature has produced a number of theories of why the U.S. Fed

[10] See Clarida, Gali, & Gertler (2000), for example.

[11] Sims and Zha (2006).

may have performed poorly including misleading real time data[12] problems with learning about key parameters of the economy[13] and changing fundamentals, such as the NAIRU, exacerbating time inconsistency problems.[14] Our paper falls into this last category, where the contribution is the use of the common international experience as a way of disciplining our empirical work, and the use of changes in economic volatility, rather than the NAIRU, as the source of the time inconsistency problem.

Finally, our paper is also somewhat related to the empirical literature investigating asymmetries central bank preferences by examining central bank policy reaction functions (see footnote 4 for references). Relative to this literature, our paper represents an indirect and somewhat inefficient test of whether central bank preferences exhibit asymmetry. However, testing whether or not central bank preferences are asymmetric is not the main objective of our paper, and our approach allows us to focus on the object of our interest: the link between economic volatility and observed inflation outcomes.

The paper proceeds as follows: Section 2 presents a model of time inconsistent monetary policy with asymmetric central bank preferences. Section 3 introduces the spurious regression problem and presents our estimates of the effect of changing volatility of unemployment shocks on inflation, correcting for this problem. Section 4 examines the relationship between inflation and its own conditional variance, as well as the general case in which inflation may depend on the conditional variance of both inflation and unemployment. Section 5 concludes.

2 Model Overview

In this section we present a model in which monetary policy is time inconsistent due to asymmetries in the central banker's loss function. The model begins with an expectations augmented short run Phillips curve, which ties deviations of unemployment from the NAIRU to unexpected inflation and exogenous shocks:

$$u_t = u^n + \lambda(\pi_t - \pi_t^e) + \eta_t \tag{2.1}$$

[12] Orphanides (2002, 2003).
[13] Sargent (1999), Primiceri (2004).
[14] Ireland (1999).

where π_t is the rate of inflation in period t, π_t^e represents households' expectations of period t inflation, u_t is the rate of unemployment, u^n is the NAIRU, and η_t is an unemployment shock, where $\eta \sim N(0, \sigma_\eta^2)$. In some applications (for example, Ireland (1999)), the NAIRU is allowed to vary over time. Here, to simplify the exposition, we use the simplest assumption, which is that the NAIRU is constant.

The central banker does not control inflation directly, but rather chooses the level of some policy instrument, i_t, which affects inflation, subject to some control error θ_t:

$$\pi_t = f(i_t) + \theta_t \tag{2.2}$$

where $f(\cdot)$ is a monotonic, continuous, differentiable function and $\theta \sim N(0, \sigma_\theta^2)$. This shock is commonly thought of as a control error, and serves to both introduce exogenous volatility into the inflation process, and break the equality between equilibrium unemployment and the NAIRU. In the absence of θ_t (i.e. if the central banker could control inflation perfectly), rational expectations implies that the private sector's forecasts of inflation would always be correct, implying that unemployment always equals the NAIRU, which does not seem reasonable.

Private agents have rational expectations, so that

$$\pi_t^e = E\{\pi_t | I_t\}, \tag{2.3}$$

where I_t is the information set of the private sector, which contains all information except the current realizations of the shocks.

The problem facing the central banker in each period is to choose i_t so as to minimize:

$$E\{L(\pi_t - \pi^*, u_t - u^*) | \Omega_t\}, \tag{2.4}$$

where $L(\pi_t, u_t)$ is the central banker's loss function, which generally depends on the deviations of inflation and unemployment from their targets, denoted by π^*, and u^*, respectively, and Ω_t is the central banker's information set, which includes all information except the current realizations of the shocks and equation (2.3).[15] The inflation target π^* is generally exogenous, and is sometimes assumed to be zero. We assume that u^* is equal to the NAIRU, thus removing the classic source of inflationary bias from the model.[16]

[15]Private sector expectations are taken as given by the central bank.

[16]Kydland and Prescott (1977), Barro and Gordon (1983).

Since the central banker takes expectations as given, he or she is unable to commit to a monetary policy rule. Instead, in each period, after the private agents have formed their expectations but before the realization of the shocks η_t and θ_t, the central banker chooses the value of the policy instrument, i_t. Thus the central banker's problem is to minimize (2.4) subject to the constraints given by (2.1) and (2.2).

We assume that the loss function is asymmetric with respect to positive versus negative deviations from the targets π^* and u^*. The literature often employs a Linex specification for preferences:[17]

$$L(\pi_t, u_t) = (e^{\alpha(\pi_t - \pi^*)} - \alpha(\pi_t - \pi^*) - 1)/(\alpha^2) + \Phi(e^{\gamma(u_t - u^*)} - \gamma(u_t - u^*) - 1)/(\gamma^2). \quad (2.5)$$

In general, when the loss function is asymmetric, the deviation of inflation from target $(E(\pi|I) - \pi^*)$ will be a function of the unemployment target, the conditional variance of inflation, and the conditional variance of unemployment. The equilibrium inflation rate in this economy is:

$$E(\pi|I) = \pi^* - (\alpha\sigma_\pi^2/2) + (1/\alpha)\ln[1 + (\alpha\lambda\Phi/\gamma)(e^{\gamma^2\sigma_u^2/2} - 1)] \quad (2.6)$$

where σ_π is the conditional variance of inflation and σ_u^2 is the conditional variance of unemployment. Note that these variances can be related back to the variances of the exogenous structural disturbances: $\sigma_\pi^2 = \sigma_\theta^2$, and $\sigma_u^2 = \sigma_\theta^2 + \gamma^2\sigma_\eta^2$.

Asymmetries in the loss function mean that central bankers dislike deviations from target with one sign more than deviations from target with the other sign. If central bankers dislike high unemployment more than they dislike low unemployment (i.e. $\gamma > 0$), the central banker would be willing to accept an unemployment rate that is below the target level on average, in return for a lower chance of suffering through a period of very high unemployment. In this case, periods of highly volatile unemployment will cause the central banker to pursue inflationary policy in an attempt to drive down the average unemployment rate, so as to insure against high unemployment shocks. Of course, the form of the Phillips curve (2.1) along with the rational expectations assumption (2.3) implies that the central bank's attempts to systematically engineer lower unemployment results, in equilibrium, only in higher inflation.

[17] Ruge-Murcia (2003a, & 2003b), and Nobay & Peel(2003).

Similarly, if central bankers' preferences are asymmetric in the inflation rate (or the deviation of the inflation rate from some target), an increase in the volatility of inflation affects the conduct of monetary policy and, consequently, the average inflation rate. To the extent that central bankers dislike high inflation more than they dislike low inflation (i.e. $\alpha > 0$), an increase in the volatility of inflation is likely to lead to lower average inflation. Higher volatility of inflation causes central bankers to reduce average inflation so as to insure against high inflation shocks.

The above intuition assumes $\alpha > 0$, $\gamma > 0$. If $\alpha < 0$, the central banker dislikes low inflation more than high inflation, implying that the central banker *raises* average inflation when the volatility of inflation is high to insure against low inflation shocks. Similarly, if $\gamma < 0$, the central banker lowers inflation when unemployment is more volatile, in an attempt to insure against low unemployment shocks. While $\alpha < 0$, and/or $\gamma < 0$ are theoretically possible, $\alpha > 0$, $\gamma > 0$ is the most intuitively plausible case.

2.1 Empirical Implications

The model takes the conditional volatility of both inflation and unemployment as exogenous. If we take the model seriously, it is reasonable to use single equation methods to estimate model parameters. From (2.6), inflation depends linearly on σ_π^2 and on the natural log of the exponential of σ_u^2. A linear estimation framework is a reasonable approximation of this first order condition. Thus, the following equation can serve as the basis for empirical work:

$$\pi_{j,t} = \pi_{j,t}^* + b_1\sigma_{u,j,t}^2 + b_2\sigma_{\pi,j,t}^2 + \epsilon_{j,t} \tag{2.7}$$

where, j is the country, t is the time period, $\epsilon_{j,t}$ is an error term, and:

$$\begin{aligned} b_1 &= (\gamma\lambda\Phi/2) \\ b_2 &= -(\alpha/2). \end{aligned}$$

While individual model parameters cannot be identified in this framework, it is possible to relate the signs of various coefficient estimates back to underlying parameters. Of main interest are the parameters related to the asymmetries in the loss function.

First note that, since λ and Φ are positive by assumption, a positive estimate of b_1 implies that γ is positive. This corresponds to the intuitively reasonable case in which the central banker dislikes high unemployment more than low unemployment. Similarly, a negative value of b_2 implies that α must be positive, which corresponds to the other intuitively attractive case, in which the central banker dislikes high inflation more than low inflation.

The model nests two special cases: If central bank preferences are symmetric in unemployment, then $\gamma \to 0$, and b_1 equals zero. Similarly, if central bank preferences are symmetric in inflation, $\alpha \to 0$, corresponding to the case where b_2 equals zero.

3 Main Results

In this section we present our main empirical results. To facilitate comparison with the existing literature, we take the case where central bank preferences are asymmetric in unemployment but not inflation as our baseline model. In this case, the model's reduced form solution for the inflation rate time series is approximated by the linear regression equation:

$$\pi_t = a + b_1 \cdot \sigma_{u,t}^2 + \epsilon_t \tag{3.8}$$

where π_t is the inflation rate in period t, $\sigma_{u,t}^2$ is the conditional variance of the unemployment rate in period t, and ϵ_t is a white noise error term. Asymmetric preferences imply that the parameter b_1 will be nonzero, being positive (negative) if the central bank prefers deviations below (above) the natural rate. Note that the more reasonable case, that central bankers dislike excessively high unemployment more than they dislike excessively low unemployment, corresponds to the case where b_1 is positive.

We estimate the variance of the unemployment rate via the following GARCH(1,1) model of the innovation in the unemployment series:

$$\begin{aligned} w_t &= v_t \sqrt{h_t} \qquad (3.9) \\ h_t &= \mu + \alpha w_{t-1}^2 + \omega h_{t-1} \end{aligned}$$

where w_t is the innovation in the unemployment rate and v_t is an i.i.d. sequence of zero-mean and unit-variance random variables. It can be shown (see, e.g., Hamilton, 1994) that this GARCH(1,1) implies that w_t^2 has an ARMA(1,1) representation whose autoregressive coefficient is $\alpha + \omega$. It is worth noting that this implies that $\alpha + \omega$ measures the persistence of the conditional variance. In the limit, $\alpha + \omega = 1$ which defines the IGARCH(1,1) model.

3.1 Spurious Regression

It would appear natural to estimate the model via a two step process: first using (3.9) to estimate the variance of the unemployment shock, and then estimating (3.8) by OLS. Ruge-Murcia (2004) conducts essentially this exercise for G-7 countries and finds some support for the predictions of the model.[18]

There is a potential problem with this approach, however. For many countries, both the inflation rate and the conditional variance of the unemployment rate are highly persistent. As a result, standard estimates of the relationship between inflation and the conditional variance of unemployment may suffer from a spurious regression problem.

To investigate this, we constructed the following Monte Carlo experiment. First, the time series $y_t, t = 1, \ldots, T$, was constructed according to the AR(1) model:

$$y_t = \rho y_{t-1} + u_t, \qquad u_t \text{ i.i.d.N}(0,1), \qquad 0 < \rho \leq 1. \tag{3.10}$$

Second, the time series $h_t, t = 1, \ldots, T$ was independently constructed as the conditional variance of the GARCH(1,1) process w_t according to:

$$\begin{aligned} w_t &= v_t\sqrt{h_t} \\ h_t &= \mu + \alpha w_{t-1}^2 + \omega h_{t-1} \\ v_t &\sim \text{i.i.d. N}(0,1) \\ \mu &= 1, \quad 0 \leq \alpha \leq 1, \quad 0 \leq \omega < 1, \quad \alpha + \omega \leq 1 \end{aligned} \tag{3.11}$$

The regression of y_t on a constant and h_t was run and a t-test was applied to the (true) null hypothesis that the regression coefficient on h_t equals zero, using a nominal test size of five-percent. Table 1 reports the rejection rates for two-sided t-tests of the null hypothesis that the slope coefficient in the regression of an AR(1) process on an independently generated GARCH(1,1) process is equal to zero, using a nominal test size of 5-percent (i.e. using a critical value equal to 1.96). The results are derived from Monte Carlo simulations using 1000 simulations for each parameter combination. T refers to the sample size. Initial values

[18] Ruge-Murcia actually estimates the model of the unemployment rate and its conditional variance jointly with (3.8) using quasi-maximum likelihood. He notes, and we have confirmed, that the results obtained by a two-step procedure in which OLS is applied to (3.8) after the conditional variances are estimated in the first step provide essentially the same results.

were set equal to zero and 500-period burn-in periods were used. The actual rejection rates, compiled for 1000 simulations and sample sizes 100, 500, and 2000, are reported.

It is clear from Table 1 that regressions of persistent time series on persistent conditional variance series suffer from the spurious regression problem, with actual rejection rates much greater than nominal rejection rates for both small and large sample sizes. This is a serious problem. For example, in Ruge-Murcia (2004), the point estimates of α and ω for France are 0.12 and 0.80, respectively, based on a sample size of approximately 120 observations. Assuming that the largest autoregressive root in the inflation rate series is at least 0.80, our results suggest that if in equation (3.8) the coefficient b_1 is equal to zero, the null hypothesis that it is zero would be incorrectly rejected at the five-percent level over 40-percent of the time!

This result is not specific to France. We have confirmed in the larger set of OECD countries that we examine in the following sections that the null hypothesis that b_1 equals zero is rejected most frequently in countries where the conditional variance of unemployment is highly persistent. Our conclusion is that this approach to estimating the model and testing the null hypothesis is likely to over-reject the null in many countries. Findings of a statistically significant correlation between inflation and the conditional variance of unemployment estimated in this framework may well be spurious. In the remainder of the paper, we attempt to correct for this problem.

3.2 Differencing

Regressions of the inflation rate on the conditional variance of the unemployment rate, are problematic in those cases where the conditional variance is an exact or approximate IGARCH process. Neither the asymptotic nor finite sample properties of regressions with IGARCH regressors have been developed, other than the simulation results we presented earlier which showed that such regressions appear to be contaminated by the spurious regression problem. The problem may be further complicated by the fact that the dependent variable, i.e. the inflation rate, appears to be an exact or approximate unit root process.

Given the highly persistent behavior of the inflation rate and, in some cases, the conditional variance of the unemployment rate, it seems reasonable to consider estimating the model in first differences. Unreported Monte Carlo results suggest that tests of the first differenced model do suffer from size distortions. However, these distortions were always in the direction of over-rejecting a true null. Given that our results imply that the null of no relationship can generally not be rejected, any size distortions in the test would appear to reinforce, rather than overturn, our conclusions.

To estimate the first differenced model, we collected a data set made up of quarterly unemployment and CPI-based inflation rates for thirteen countries:[19] Australia, Austria, Canada, Denmark, Finland, France, Germany, Italy, Japan, Norway, Sweden, the United Kingdom, and the United States. The data are taken from the OECD's Main Economic Indicators database. We used the largest sample available in each country, which for most countries spans from the mid to late 1960s until 2004.

We began by estimating the model of the previous section, this time in differences. Taking first differences of 3.8 gives

$$\Delta\pi_t = b_1 \Delta\sigma^2_{u,t} + e_t, \tag{3.12}$$

where $\Delta x_t = x_t - x_{t-1}$, and $\sigma^2_{u,t}$ is the conditional variance of the unemployment rate.[20] Assuming that the inflation target is constant, the model suggests that the differenced form of the regression should be fit without an intercept. We ran the differenced regression with and without an intercept. The results with respect to the parameter b_1 were virtually identical and the intercept was not statistically significant in any case.[21]

We first tested for a unit root in the unemployment rate and found that the unit root null could not be rejected for any of these countries. Since the parameter b_1 in equation (3.8) is not identified in the absence of time-varying conditional heteroskedasticity, we next tested for ARCH effects in the first difference of the unemployment rate. Each differenced unemployment series was fit to an AR(p) model, where p was selected by the AIC, and the squared residual series was fit to AR(1), AR(2), AR(3), and AR(4) models. We applied Engle's (1982) LM test for neglected ARCH effects, and if any of the p-values was less than

[19] Our conclusions are robust to the use of the GDP deflator as the price measure for countries for which this series is available.

[20] We also estimated the model using the conditional variance of the unemployment gap, recovered from a cubic de-trending of the unemployment rate. This allows for the possibility that monetary policy makers interpret low frequency changes in the unemployment rate as changes in the NAIRU, and target monetary policy towards the deviations of the actual unemployment rate from the NAIRU. The use of the conditional variance of the unemployment gap did not change the results substantively.

[21] The results we report are for the regressions run with an intercept.

or equal to 10-percent, we took this as evidence of possible conditional heteroskedasticity.[22] If the differenced unemployment series displayed such evidence of conditional heteroskedasticity, we fit the first difference of the rate of inflation to the first difference of the estimated GARCH(1,1) conditional variance series for the differenced unemployment rate's innovations.

The results are presented in Table 2. Column 1 of the table reports only the lowest of the p-values from the LM tests, as we included in the sample countries displaying any evidence of possible conditional heteroskedasticity. We find evidence of time varying conditional heteroskedasticity in seven out of the 13 countries. In two of these countries (Austria and the U.K.), the ARCH effects were not persistent, and in both of these cases the estimate of b_1 was positive, though statistically significant only in one (Austria) and then only at the 10% level. In the remaining five countries, we detected persistent ARCH effects. In three of these countries, the point estimate of b_1 was negative, but statistically significant in only two of these three countries. In the remaining three countries, the estimate of b_1 was positive, but not statistically significant.

These results suggest that the mechanism posited by the model does not fit very well as an explanation of time series patterns in inflation in our sample of countries. First, there is no evidence of time variation in the conditional volatility in about half of the countries in the sample, which suggests that the model cannot explain changing inflation rates in these countries. Furthermore, in about half of the countries for which ARCH effects are present the estimate of the key parameter, b_1, is of the theoretically unattractive sign. In only one of the 13 countries, Austria, do we find both time varying ARCH effects and statistically significant evidence that the conditional variance of unemployment is positively related to the change in inflation, as would be required for the theory to serve as a good explanation of observed inflation trends.

3.3 Cointegration

One problem with the results in the previous section is that differencing the data emphasizes the high frequency features of the data at the expense of lower frequency movements. Given that our motivation is to ask whether time inconsistency problems can explain long

[22]The LM test is based on the distribution of the $T \cdot R^2$ statistic from the regression of the squared residuals from the AR(p) model.

run trends in OECD-wide inflation, this loss of information at the low frequency is an unattractive feature of differencing as a corrective to the spurious regression problem.

Consequently, we would like a solution for the spurious regression problem that preserves information about low frequency movements in the data. Our solution is to re-cast the model as a cointegrating relationship. Observe, from (2.7) and (3.8), that the model predicts that inflation should inherit the time series properties of the conditional variance of unemployment. If this conditional variance is I(1), then inflation should also be I(1), and the model predicts that these two variables will be cointegrated. We test this implication of the theory by applying Johansen's (1988) maximum likelihood approach to test for cointegration between a country's inflation rate and the the estimated conditional variance of unemployment.

We report Johansen's λ-max test statistic, which tests the null of no-cointegration against the alternative that the two series are cointegrated. All tests are performed with unrestricted intercepts and no deterministic time trend component in the VECM. We selected a relatively long lag length (8 lags) in the VECM in the hope of alleviating any serial correlation problems. Critical values for the statistic, obtained using the procedure of MacKinnon, Haug, and Michelis (1999), are: 12.30 (10%), 14.26 (5%), 18.52 (1%).

Column 4 of Table 2 reports the results of Johansen's λ-max test.[23] The results suggest there is some support of the hypothesis that long run trends in inflation rates are related to trends in conditional volatility in a cointegrating framework. The null hypothesis of no cointegration is rejected at standard significance levels for four of the seven countries in our sample for which there is evidence of time varying ARCH effects.

Under plausible beliefs about the nature of the asymmetry in central bank preferences, the model predicts that the sign of this relationship is positive. Column 5 of Table 2 reports the sign of the estimated cointegrating vector. Of the four countries for which there is evidence of a cointegrating relationship, the estimate of b_1 has the wrong sign in one of these (Germany). Thus, viewing the model as a cointegrating relationship does improve its fit marginally, but does not alter the overall conclusion that there is little evidence that this mechanism explains OECD inflation patterns.

[23] In the bivariate setting with both series assumed to be unit root processes, Johansen's λ-max and λ-trace tests are equivalent, since in both cases the null hypothesis is that the cointegration rank is zero and the alternative hypothesis is that the cointegration rank is one.

3.4 Subsample Analysis

A concern with the previous results is that they are derived holding central banks' inflation targets constant. Given the large movements in inflation and well known changes in the conduct of monetary policy that took place during our sample period, we would like to check the robustness of our results to this assumption. Hence, we re-estimate the model on a more recent subsample of data. In particular, we break the sample in 1979 Q3 and then estimate the model on the recent data only. In essence we disregard the rise of inflation and ask whether or not the model can help explain the decline in inflation to the more recent low and stable levels, with the advantage being that the conduct of monetary policy, including central banks' inflation targets, are more likely to be constant over the more recent years.[24]

As before, we first test for time varying ARCH effects in unemployment over the new sample. Column 1 of Table 3 reports the results. We find evidence of such effects in five of the thirteen countries. Relative to the full sample, Denmark, Germany and the U.S. drop out of the set of countries for which we find time varying ARCH effects in unemployment, and France joins that set.

We proceed to estimate the theoretical model for these five countries, first differencing the data, as in section 3.2, to correct for the spurious regression problem. Column 2 of the Table 3 reports the estimated b_1 coefficients. Of the five countries for which there is evidence of time varying ARCH effects in the later sample, only one country (Austria, again) shows evidence of a positive and statistically significant relationship between the conditional volatility of unemployment and the inflation rate. Of the remaining four countries, the point estimate of b_1 is negative in two, but statistically significant in neither of these.

Again, we are concerned that differencing the data emphasizes high frequency movements in the data at the expense of the low frequency effects we are most interested in, so we also estimate the model as a cointegrating relationship. Columns 4 and 5 report the results from the Johansen cointegration test described in section 3.3. The null hypothesis of no cointegration is rejected, at conventional significance levels, in four of the five countries for which we found evidence of time varying conditional heteroskedasticity. In each of these

[24]Our break date is also widely thought to represent a change in the conduct of monetary policy away from the approach of the 1970s to the modern regime.

four cases, the point estimates of b_1 are positive, which is the theoretically attractive sign.

Overall, focusing only on the, more recent, decline in inflation rather than the whole rise and fall does not change the results very much. When we estimate the model we find essentially the same results as before: the model fits nicely for Austria, but not elsewhere. When viewed as a cointegrating relationship, there is stronger evidence in favour of the model in the more recent data than in the full sample. Even here, however, the evidence suggests that that the relationships suggested by the model are present in the data for only about one third of the countries in our sample.

4 Inflation Volatility

Given the lack of support for the view that changes in the conditional variance of unemployment cause changes in inflation, we turn our attention to the case where preferences are asymmetric in inflation, but not unemployment. The model's reduced form solution in this case results in a linear equation relating π to its conditional variance:

$$\pi_t = a + b_2 \cdot \sigma^2_{\pi,t} + \epsilon_t \tag{4.13}$$

We estimate the model in first differences, as both the inflation rate and the conditional variance of the inflation rate exhibit high persistence. Taking first differences of 4.13 gives:

$$\Delta\pi_t = b_2 \Delta\sigma^2_{\pi,t} + e_t, \tag{4.14}$$

which forms the basis of our estimation. [25] Asymmetric preferences in inflation imply that the parameter b_2 will be nonzero, being negative (positive) if the central bank prefers deviations below (above) the target rate of inflation. Note that for positive inflation rates, as observed in our sample, the more reasonable case is that central bankers dislike excessively high inflation more than they dislike excessively low inflation, which corresponds to the case where b_2 is negative.

As in the previous case, the slope coefficient in 4.14 is only identified if the conditional variance is time-varying, so we began by testing for ARCH effects in the first difference of the inflation rate using the procedure outlined above for the unemployment rate. If the

[25] As before, the results with respect to the parameter b_2 in regressions with and without an intercept were virtually identical and the intercept was not statistically significant in any case. The reported results include an intercept.

inflation series displayed evidence of conditional heteroskedasticity, we fit the series to a GARCH-M regression with a GARCH(1,1) error specification, using the quasi-maximum likelihood estimator to get an estimate of the conditional variance of the inflation rate. We then estimate the slope coefficient, b_2, by fitting the first difference of inflation to the first difference of the estimated conditional variance series.[26]

Results are presented in columns 1-3 of Table 4. First off, we find evidence of a time varying conditional variance in inflation in 10 of the countries in the sample. Of these 10 countries, the ARCH effects are not persistent in three cases. In each of these cases, the estimate of b_2 is both negative and statistically significant. In the remaining six countries, for which we find evidence on persistent ARCH effects, the estimates of b_2 are positive but not statistically significant in two countries, and negative in four countries, though not statistically significantly in three of the latter.

As before, we are concerned that by differencing the data we lose information on low frequency movements. Again, we address this problem by adopting a cointegration perspective, and apply Johansen's (1988) maximum likelihood approach to test for cointegration between a country's inflation rate and the the estimated conditional volatility of inflation.

As column 4 of Table 4 shows, there is strong support for the view that inflation and the conditional volatility of inflation are cointegrated. Of the 10 countries for which there is evidence of time varying heteroskedasticity, the null hypothesis of no cointegration is rejected at standard significance levels for seven. Column 5 reports the estimates of b_2 in the cointegrating vector. Here we find that, in all cases but one, the estimates of b_2 are positive. However, since the most plausible version of central bank preferences would imply a negative b_2, it is unclear that these results represent evidence in favor of the theory.

The results in this section are mixed. In the differenced version of the model, there is a statistically significant relationship of the correct sign in 5 of the 10 countries for which we find evidence of time varying ARCH effects in inflation. However, when we turn to

[26]For completeness, we also analyzed the effect of time-varying conditional variances using the level of inflation, by fitting a GARCH-M in inflation as suggested by equation 4.13. We found evidence of conditional heteroskedasticity in the CPI-based inflation rate for 10 of the 18 countries in the sample (Australia, Austria, Canada, Finland, Germany, Italy, Japan, Norway, United Kingdom, and United States). Of these 10 countries, all except for Austria had a positive and statistically significant estimated coefficient on the estimated conditional variance of the inflation rate in the GARCH-M regression. All 10 of these countries displayed persistent conditional heteroskedasticity, with the sum of the estimated GARCH(1,1) coefficients exceeding 0.75. In 9 cases, the GARCH(1,1) coefficients exceeded 0.960 and in five cases the constraint that the sum of the GARCH(1,1) coefficients cannot exceed unity was binding.

the cointegration estimation, the sign of the relationship changes, suggesting that the time inconsistency story may not be the correct explanation for the co-movement of inflation and its volatility. Overall, while there does appear to be evidence of a relationship between inflation and its conditional volatility, the results do not represent compelling evidence in favor of the theoretical model as a driver of inflation trends. Even in the best case, which is the differenced estimates, the model fits in only five of the 13 countries in the sample.

4.1 Unemployment and Inflation Volatility

In the previous sections we examined separately the hypothesis that changes in the conditional variance of unemployment cause changes in inflation and the hypothesis that changes in the conditional variance of inflation drive inflation trends. However, central bank preferences may be asymmetric in both inflation and unemployment. In this case, the conditional variance of both inflation and unemployment should affect inflation, meaning that our previous results suffer from an omitted variable problem.

Consequently, we estimate the model with asymmetric preferences in both inflation and unemployment. The model's reduced form solution in this case results in a linear equation relating π to the conditional variance of π and u:

$$\pi_t = a + b_1 \cdot \sigma^2_{u,t} + b_2 \cdot \sigma^2_{\pi,t} + \epsilon_t \tag{4.15}$$

We estimate the model in first differences, as both the inflation rate and the conditional variance of the inflation rate exhibit high persistence. Taking first differences of 4.15 gives:

$$\Delta\pi_t = b_1 \Delta\sigma^2_{u,t} + b_2 \Delta\sigma^2_{\pi,t} + e_t, \tag{4.16}$$

which forms the basis of our estimation. As in the previous cases, the results with respect to the parameters b_1, and b_2 in regressions fit with and without an intercept were virtually identical and the intercept was not statistically significant in any case.

Model 4.16 is only relevant for countries in which the conditional variance of both unemployment and inflation is time-varying. For these countries, we estimate the slope coefficients, b_1 and b_2, by fitting the first difference of inflation to the first difference of the conditional variance series, where these series are estimated as described in the previous subsections. The main limitation of this analysis being that there are only five countries for which both inflation and unemployment display evidence of conditional heteroskedasticity: Austria, Canada, Denmark, Germany, and the U.K..

The results are presented in Table 5. In the differenced specification (columns 1-3) only two of the 10 estimated slope coefficients, are statistically significant: the coefficient on the volatility of unemployment for Denmark, and the volatility of inflation for Germany. In both cases, these coefficients take on the theoretically unattractive sign.

Again, we also apply Johansen's (1988) maximum likelihood approach to test for cointegration between a country's inflation rate and the the estimated conditional volatilities of unemployment and inflation. In the tri-varate case, critical values for the statistic are: 18.8928 (10%), 21.1314 (5%), 25.8650 (1%). The results, reported in columns 4-6 of the table, are that the null hypothesis of no cointegration is rejected at standard significance levels for three of the five countries. In this case, the estimates of b_1 have the correct sign but, as before, the estimates of b_2 are positive rather than negative.[27]

Overall, these results, though limited in scope, suggest that the conclusions of the previous sections are not sensitive to the possibility that central bank preferences are asymmetric in both inflation and unemployment.

5 Conclusion

The results presented in this paper provide little support for the view that the interaction of asymmetric central bank preferences and the volatilities of inflation and unemployment are important determinants of inflation. The data does not support the view that the volatility of unemployment helps explains inflation. Estimated coefficients on measures of this volatility are generally not statistically significantly different from zero, and frequently possess the wrong sign. In the best case (the cointegration estimates in the more recent sample) the theory seems capable of explaining inflation outcomes in only one third of the countries in the sample.

The results concerning the relevance of asymmetric preferences in inflation are more difficult to interpret. There is some evidence of a relationship between inflation and its own volatility, but this relationship only takes on the sign suggested by the theory when

[27] In this case, however, there may be multiple cointegrating vectors, so the interpretation of these coefficients is unclear.

we difference the data, which emphasizes the shorter run features of the data. Even in this case, the evidence statistically significant in only five of the 13 countries we examine.

Overall, our conclusion is that while it appears promising on the surface, a careful investigation of the data does not provide much support for the view that the combination of asymmetric central bank preferences and changing economic volatility is an important driver of recent inflation trends in OECD countries.

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Table 1. Testing the Significance of the Slope Coefficient in a Spurious Regression: Actual Rejection Rates for a Test with Nominal Size Equal to 0.05

T=100

(α, ω)	$\rho = 0.00$	$\rho = 0.40$	$\rho = 0.80$	$\rho = 0.90$	$\rho = 0.95$	$\rho = 0.99$	$\rho = 1.00$
(0.1,0.1)	0.054	0.083	0.103	0.091	0.091	0.110	0.113
(0.1,0.4)	0.044	0.118	0.203	0.233	0.243	0.239	0.241
(0.1,0.8)	0.043	0.168	0.416	0.495	0.568	0.611	0.595
(0.1,0.9)	0.049	0.207	0.474	0.596	0.698	0.733	0.755
(0.01,0.99)	0.052	0.202	0.495	0.631	0.665	0.765	0.749
(0.4,0.1)	0.065	0.092	0.162	0.184	0.196	0.181	0.194
(0.8,0.1)	0.042	0.108	0.198	0.253	0.276	0.285	0.276
(0.9,0.1)	0.043	0.106	0.232	0.266	0.303	0.300	0.312
(0.95,0.05)	0.043	0.102	0.230	0.252	0.240	0.270	0.305
(1.0,0.0)	0.070	0.105	0.181	0.225	0.239	0.260	0.275

T=2000

(α, ω)	$\rho = 0.00$	$\rho = 0.40$	$\rho = 0.80$	$\rho = 0.90$	$\rho = 0.95$	$\rho = 0.99$	$\rho = 1.00$
(0.1,0.1)	0.044	0.079	0.090	0.108	0.087	0.106	0.114
(0.1,0.4)	0.067	0.140	0.205	0.233	0.236	0.279	0.265
(0.1,0.8)	0.057	0.186	0.405	0.530	0.598	0.627	0.667
(0.1,0.9)	0.040	0.180	0.518	0.624	0.727	0.837	0.877
(0.01,0.99)	0.056	0.201	0.497	0.641	0.766	0.883	0.941
(0.4,0.1)	0.046	0.114	0.173	0.190	0.213	0.238	0.244
(0.8,0.1)	0.052	0.136	0.247	0.276	0.340	0.352	0.357
(0.9,0.1)	0.035	0.129	0.242	0.291	0.341	0.331	0.388
(0.95,0.05)	0.040	0.114	0.227	0.298	0.334	0.336	0.357
(1.0,0.0)	0.051	0.117	0.233	0.282	0.279	0.304	0.332

Table 2. Inflation and the Conditional Variance of Unemployment

	LM-test	Differenced		Cointegration	
Country	p-values	$\hat{b}_1$	$\alpha + \omega$	λ-max	$\hat{b}_1$
Australia	0.20	-	-	-	-
Austria	**0.10**	106.4*	0.124	7.43	53.34
Canada	**0.00**	1.5	0.662	20.73***	5.88
Denmark	**0.00**	0.9	0.995	9.37	31.90
Finland	0.26	-	-	-	-
France	0.42	-	-	-	-
Germany	**0.06**	-18.4	0.716	26.03***	-15.12
Italy	0.11	-	-	-	-
Japan	0.30	-	-	-	-
Norway	0.56	-	-	-	-
Sweden	**0.00**	-39.2**	0.903	9.62	-38.96
UK	**0.00**	10.8	0.301	20.19***	35.41
US	**0.00**	-7.1**	0.897	17.18**	7.93

Notes: * = reject at the 10% level, ** = reject at the 5% level, *** = reject at the 1% level

Table 3. Inflation and the Conditional Variance of Unemployment: Late Sample

	LM-test	Differencing		Cointegration	
Country	p-values	$\hat{b}_1$	$\alpha + \omega$	λ-max	$\hat{b}_1$
Australia	0.95	-	-	-	-
Austria	**0.02**	85.41**	0.224	12.48*	82.60
Canada	**0.01**	-0.24	0.775	20.21***	5.19
Denmark	0.23	-	-	-	-
Finland	0.25	-	-	-	-
France	**0.05**	11.07	0.305	32.03***	42.12
Germany	0.29	-	-	-	-
Italy	0.23	-	-	-	-
Japan	0.52	-	-	-	-
Norway	0.50	-	-	-	-
Sweden	**0.00**	-9.6	0.922	8.75	-22.92
UK	**0.00**	66.3	0.974	13.02*	161.92
US	0.17	-	-	-	-

Notes: * = reject at the 10% level, ** = reject at the 5% level, *** = reject at the 1% level

Table 4. Inflation and the Conditional Variance of Inflation

	LM-test	Differencing		Cointegration	
Country	p-values	$\hat{b}_2$	$\alpha + \omega$	λ-max	$\hat{b}_2$
Australia	**0.01**	-0.037**	0.350	24.27***	0.097
Austria	**0.00**	-0.078***	1.000	13.95*	0.006
Canada	**0.01**	-0.087	0.732	11.86	-0.085
Denmark	**0.04**	0.026	0.928	15.42**	0.118
Finland	**0.01**	-0.080	0.918	10.33	0.788
France	0.18	-	-	-	-
Germany	**0.05**	-17.363**	0.390	18.42**	51.548
Italy	**0.01**	0.040	1.000	20.11***	0.260
Japan	**0.00**	-0.066***	0.369	24.94***	0.213
Norway	**0.00**	-0.004	0.644	3.90	0.025
Sweden	0.36	-	-	-	-
UK	**0.00**	-0.113***	0.332	19.93***	0.286
US	0.14	-	-	-	-

Notes: * = reject at the 10% level, ** = reject at the 5% level, *** = reject at the 1% level

Table 5. Inflation and the Conditional Variance of Inflation and Unemployment

	Differencing			Cointegration		
Country	$\hat{b}_1$	$\hat{b}_2$	$\alpha + \omega$	λ-max	$\hat{b}_1$	$\hat{b}_2$
Australia	-	-	-	-	-	-
Austria	60.66	0.013	0.993	12.37	71.77	0.000
Canada	0.05	-0.03	0.722	19.73*	13.70	0.06
Denmark	-27.86**	0.01	0.929	18.67	30.67	0.06
Finland	-	-	-	-	-	-
France	-	-	-	-	-	-
Germany	-1.595	73.59***	0.342	27.86***	2.64	5.84
Italy	-	-	-	-	-	-
Japan	-	-	-	-	-	-
Norway	-	-	-	-	-	-
Sweden	-	-	-	-	-	-
UK	-33.42	0.03	0.337	28.64***	38.68	0.09
US	-	-	-	-	-	-

Notes: * = reject at the 10% level, ** = reject at the 5% level, *** = reject at the 1% level

Fig. 1. Australia

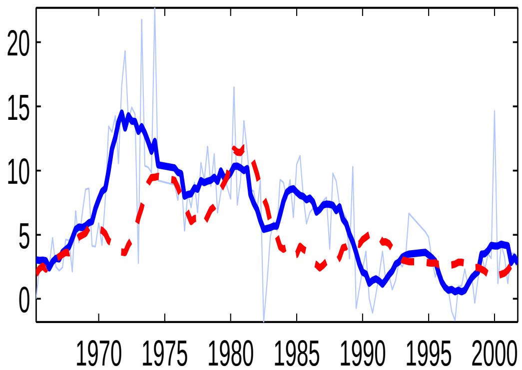

Fig. 2. Austria

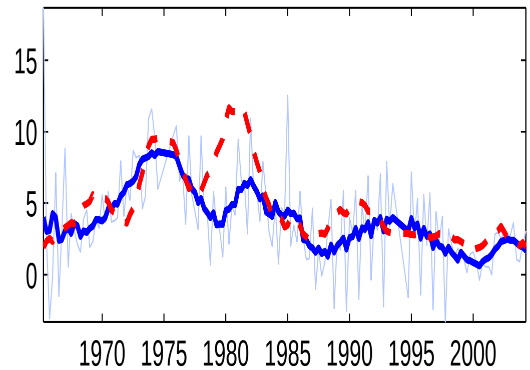

Fig. 3. Canada

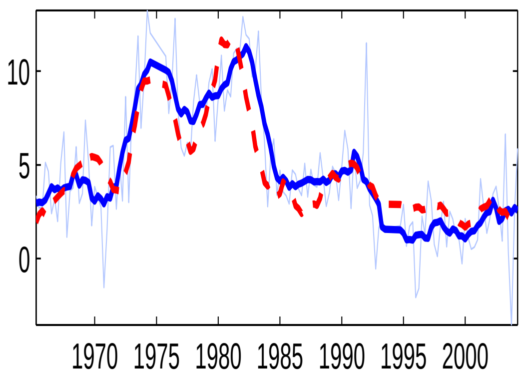

Fig. 4. Denmark

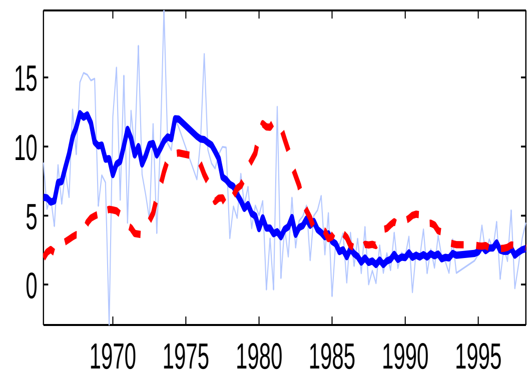

Fig 5. Finland

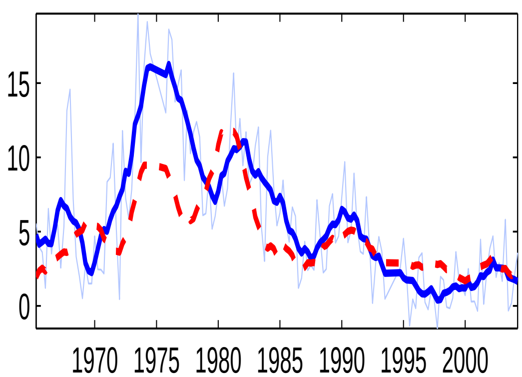

Fig. 6. France

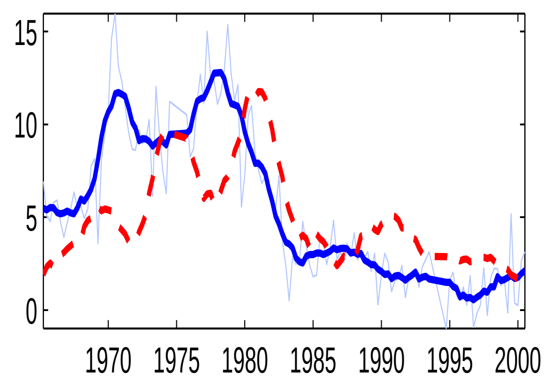

Fig. 7. Germany

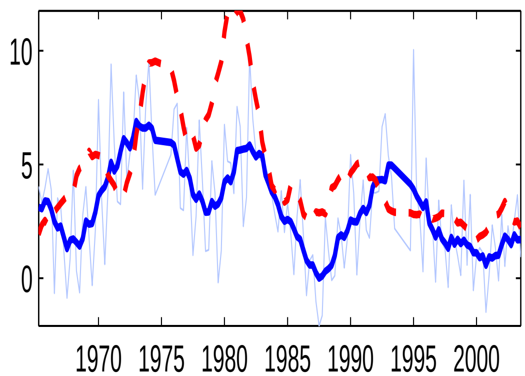

Fig. 8. Italy

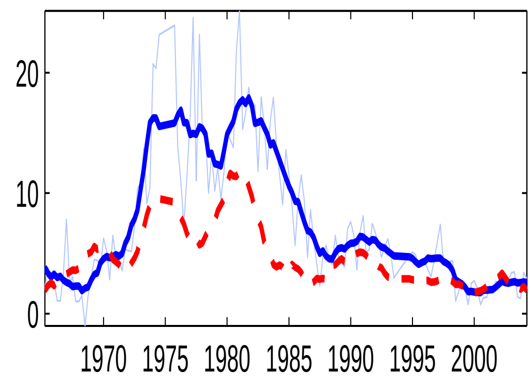

Fig. 9. Japan

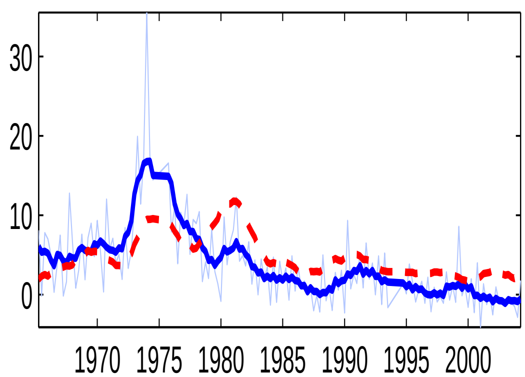

Fig. 10. Norway

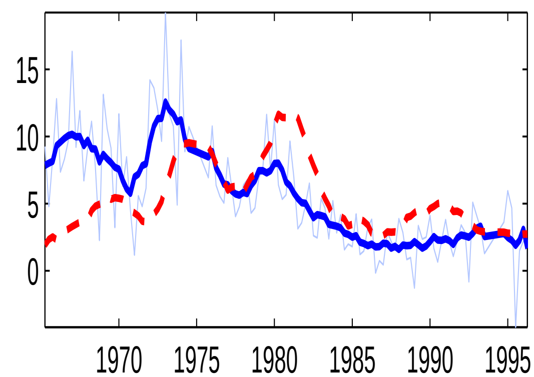

Fig. 11. Sweden

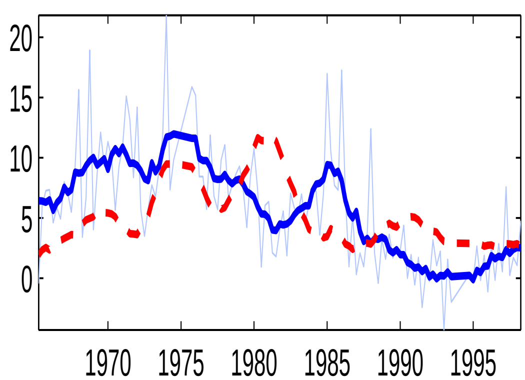

Fig. 12. U.K.