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# Do Asymmetric Central Bank Preferences Help Explain Observed Inflation Outcomes?

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**Abstract**

Recent theoretical work shows that changes in the volatility of inflation and/or unemployment affect equilibrium inflation outcomes when the central banker's loss function is asymmetric. We show that previous evidence offered in support of the proposition that the volatility of unemployment helps explain inflation outcomes 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 the data is not strongly supportive of the view that asymmetric central bank preferences are an important driver of inflation.

KEYWORDS: Inflation, Monetary Policy, Asymmetric Loss.

JEL CLASSIFICATION: E50, E61

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# 1 Introduction

The view that central bank preferences are asymmetric in deviations of inflation and/or unemployment from their target levels has received increasing attention from macroeconomists, in part due to its resonance with the comments of policy insiders (examples include Blinder (1997, 1998) and Goodhart (1998)). Empirical work investigating such asymmetries and the resulting non-linearities in monetary policy reaction functions finds evidence of asymmetries in central bank preferences.[1] Recent theoretical work (Cukierman (1999), Ruge-Murcia(2002, 2003a), Gerlach (2003), and Nobay & Peel (2003)) shows that when central banker's preferences are asymmetric, the variances of inflation and/or unemployment affect equilibrium inflation. These results suggest that the interaction of asymmetric central bank preferences and the volatilities of inflation and unemployment may be an important determinant of inflation.

In this paper we ask to what extent this mechanism can explain observed inflation trends within countries over time as well as observed cross country differences in average inflation rates. Our results consistently suggest that volatility in unemployment does not help explain either within or across country inflation outcomes. The evidence concerning whether or not inflation volatility can help explain inflation outcomes, is less conclusive. The time series evidence suggests that high inflation volatility is associated with a lower inflation rate, at least in some OECD countries, though our results imply that this mechanism does not explain much of the time series variation of inflation.

In the cross country data, we do find evidence of a strong, positive correlation between the level inflation and its variance. While this suggests a relationship between inflation and the volatility it does not constitute strong evidence in favor of the mechanism of asymmetric central bank preferences, as the sign of the correlation implies that central bankers dislike

[1]Reduced form tests for asymmetries in policy reaction functions include Bec, Ben Salem, & Collard (2002), Kim, Osborn, & Sensier (2002), Cukierman & Muscatelli (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 (2005). While the literature has yet to arrive at a consensus view, most investigations find support for some form of asymmetry.

low inflation more than they dislike high inflation.

In the classic Kydland-Prescott, Barro-Gordon (KPBG) monetary policy framework (Kydland & Prescott (1977), Barro & Gordon (1983)), if central bankers attempt to target an unemployment rate below the NAIRU, policy suffers from a time inconsistency problem, and that monetary policy exhibits an inflation bias. The standard formulation models central bankers' preferences as a symmetric, usually quadratic, loss function in the deviations of inflation and unemployment from their respective targets. Extending this framework to incorporate an asymmetric loss function allows for the possibility that central bankers may dislike deviations from target of one sign more than those of the other sign. For example, a central banker may strongly dislike shocks that drive unemployment above the NAIRU, but be more willing to tolerate shocks that drive unemployment below the NAIRU.

In an economy where the central banker's preferences are asymmetric, inflation depends on the variances of inflation and unemployment. Consider an economy in which the variance of shocks to unemployment increases. With a higher variance, the probability of an episode of very high (and 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 (were it to succeed) would make an episode of very low unemployment more likely. 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.

In equilibrium, however, private agents anticipate the central banker's expansionary policy with the result that the effect of this policy is to drive up the inflation rate. This outcome closely resembles the time inconsistency result of the original KPBG model. The main difference here being that, rather than targeting an unattainably low unemployment rate, the source of the inflationary bias is the central banker's wish to drive down the unemployment rate to reduce the probability of a very high unemployment episode.

In this paper we investigate the value of this mechanism in explaining inflation outcomes observed in both time series, and cross sectional data. We begin by asking whether changes in the volatility of inflation and unemployment over time within a country are related to times series trends in inflation that we observe within that country. Recent evidence (for example, Ruge-Murcia (2004)) offers some support for the proposition that asymmetric preferences with regards to unemployment can explain inflation trends in at least some of the G-7 countries.

We employ Monte Carlo methods to show that, when inflation is persistent (which is the case in most, if not all, 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 exhibits persistence. Existing empirical work, which is confirmed by our own investigation, finds that changes in the conditional volatility of unemployment have statistically significant effects on changes in inflation only in those countries for which the conditional variance of unemployment is most persistent. We conclude that these effects are likely spurious.

In an attempt to eliminate the spurious regression problem, we re-do the analysis by taking first differences of inflation and the conditional variance of unemployment. The results suggest that changes in the conditional variance of unemployment are not correlated with changes in inflation, implying 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 the conditional variance of inflation and inflation in time series data for OECD countries, again using first differences of inflation and its conditional volatility to reduce the likelihood of spurious results.[2] The results here are more mixed than the unemployment results. There is statistically significant evidence of an inverse relationship between inflation and its conditional volatility, as would be implied if asymmetric central bank preferences were important, in about one third of the countries.[3]

[2]Our econometric work follows the model by assuming that the conditional volatility 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)).

[3]For completeness, we estimated this relationship in levels, rather than first differences. In this case, the results suggest a strong, positive relationship between changes in the conditional variance of inflation and

On the other hand, the data rejects such a relationship in about half of the countries in the sample, with the results for the remaining countries being more inconclusive. Furthermore, even in those countries for which there is evidence of a relationship between inflation and its conditional volatility, the model generally explains very little of the variation of changes of inflation. These results suggest that while changes in the volatility of inflation, interacting with asymmetric central bank preferences, may affect inflation to a degree in a subset of countries, there is no strong support for the view that this mechanism is an important determinant of inflation trends.

Finally we ask whether cross country differences in the volatility of inflation and unemployment can explain observed cross country differences in average inflation rates in samples including both developed and developing countries. In previous work, Ruge-Murcia (2002), and Cukierman & Gerlach (2003) test versions of the asymmetric preference model on cross sectional data for samples of OECD countries. Relative to these papers, our contribution is to increase the sample of countries examined, as well as to control for additional factors that have been identified as important influences on inflation in the empirical literature on cross country inflation differences.

As in the time series case, the volatility of unemployment does not contribute much in explaining cross country differences in inflation in any of our regressions. This conclusion is robust to a variety of permutations of the basic regression model designed to control for problems that arise due to the nature of the pooled cross country data.

The volatility of inflation, in the cross sectional data, is consistently positively correlated with average inflation rates. It is not clear that this is due to the mechanism of asymmetric central bank preferences. In order to generate a *positive* correlation between inflation and the volatility of inflation, the central banker would have to dislike negative deviations from the inflation target more than he or she dislikes positive deviations. In other words, the

changing inflation outcomes. From the view of a model based on asymmetric central bank preferences, the main problem with this result is that, in order to generate a *positive* correlation between inflation and the volatility of inflation, the central banker would have to dislike negative deviations from the inflation target more than he or she dislikes positive deviations.

model would explain a positive correlation between inflation and its conditional variance as resulting from policy makers in countries with volatile shocks to inflation deliberately engineering high average inflation rates in order to avoid low inflation outcomes. While this is possible, it seems intuitively to be the less likely case, and is inconsistent with more direct estimates of asymmetries in central bank preferences.

The paper proceeds as follows: Section 2 outlines the model, Sections 3 and 4 provide empirical results from time series and cross-sectional tests of the model, respectively. 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 in departures of both inflation and unemployment from their respective target levels. Since the material in this section is drawn directly from the work of others, we restrict ourselves to outlining the model, and presenting the equilibrium outcome, along with the accompanying intuition. Readers desiring a more detailed exposition are referred to Ruge-Murcia (2002), from which this version of the model is taken. Cukierman (1999), Gerlach (2003), Nobay and Peel (2003), and Ruge-Murcia (2003a; 2003b; 2004) offer variants of the model.

## 2.1 The Basic Framework

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:

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

where $\pi_t$ is the rate of inflation in period $t$, $\pi_t^e$ represents households' expectations of period $t$ inflation, $u_t$ is the rate of unemployment, $u^n$ is the NAIRU, and $\eta_t$ 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 $\theta_t$:

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

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 $\theta_t$ (i.e. if the central banker could control inflation perfectly), rational expectations implies that the private sectors forecasts of inflation would always be correct, implying that unemployment always equaled 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 $\pi^*$, and $u^*$, respectively, and $\Omega_t$ is the central banker's information set, which includes all information except the current realizations of the shocks and equation (2.3) (as private sector expectations are taken as given by the central bank). The central banker's optimization is subject to the constraints given by (2.1) and (2.2).

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 $\eta_t$ and $\theta_t$, the central banker chooses the value of the policy instrument, $i_t$.

In the Kydland-Prescott Barro-Gordon version of this model, the central banker's loss function is quadratic in the deviation of unemployment from target, and in the deviation of inflation from target:

$$L(\pi_t - \pi^*, u_t - u^*) = (1/2)(\pi_t - \pi^*)^2 + (1/2)(u_t - u^*)^2 \tag{2.5}$$

The unemployment target is $k \cdot u^n$, where $0 < k \leq 1$. Time inconsistency of policy results in the case where $k < 1$, as the central banker attempts to use expansionary monetary policy to drive unemployment below the NAIRU. The private sector understands this policy, and adjusts its expectations of inflation accordingly. The result is sub-optimally high inflation, with no corresponding decline in unemployment.

### 2.2 Asymmetric Preferences

Recent work extends this basic framework by replacing the quadratic loss function with an asymmetric function. The Linex specification for preferences is popular (Ruge-Murcia (2002, 2003a, 2003b, 2004), and Nobay & Peel(2003)) as it is tractable and nests the KPBG outcome as a special case:

$$L(\pi_t, u_t) = (e^{\alpha(\pi_t - \pi^*)} - \alpha(\pi_t - \pi^*) - 1)/(\alpha^2) + \Phi(e^{\gamma(u_t - u^*)} - \gamma(u_t - u^*) - 1)/(\gamma^2). \tag{2.6}$$

While the central bank's unemployment target, $u^*$, could be anything in principle, it is usually assumed to be related to the NAIRU. Here we make the typical assumption that $u^*$ is some fraction of the NAIRU: $u^* = k \cdot u^n$, where $k \leq 1$. The inflation target $\pi^*$ is generally exogenous, and is sometimes assumed to be zero.

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:

$$E(\pi|I) = \pi^* - (\alpha\sigma_\pi^2/2) + (1/\alpha)\ln[1 + (\alpha\lambda\Phi/\gamma)(e^{\gamma(1-k)u^n(\gamma^2\sigma_u^2)/2}1)] \tag{2.7}$$

where $\sigma_\pi$ is the conditional variance of inflation and $\sigma_u^2$ 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.

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.

## 2.3 Empirical Implications

The model takes the NAIRU, and 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. Since the equilibrium condition for inflation involves taking the log of an exponential function, an estimation framework that employs a linear approximation of this first order condition is a reasonable approximation. Thus, the following equation can serve as the basis for empirical work:

$$\pi_{j,t} = \pi^*_{j,t} + b_1 u^n_{j,t} + b_2 \sigma^2_{u,j,t} + b_3 \sigma^2_{\pi,j,t} + \epsilon_{j,t} \tag{2.8}$$

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

$$\begin{aligned} b_1 &= \lambda\Phi(1-k) \\ b_2 &= (\gamma\lambda\Phi/2) \\ b_3 &= -(\alpha/2). \end{aligned}$$

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 $\lambda$ and $\Phi$ are positive by assumption, a positive estimate of $b_2$ implies that $\gamma$ 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 $b_3$ implies that $\alpha$ must be positive, which corresponds to the other intuitively attractive case, in which the central banker dislikes high inflation more than low inflation. Finally, it is possible to relate $b_1$ back to the KPBG inflation bias, as a positive value of $b_1$ corresponds to a case where there is an inflation bias, whilst a zero value of $b_1$ would correspond to the case where the central banker targets the NAIRU.

The model nests a number of special cases: As the central bank targets the NAIRU, $k \to 1$, so the KPBG inflation bias vanishes, and $b_1$ equals 0. If central bank preferences are symmetric in unemployment, then $\gamma \to 0$, and $b_2$ equals zero. Similarly, if central bank preferences are symmetric in inflation, $\alpha \to 0$, corresponding to the case where $b_3$ equals zero.

# 3 Evidence from Time Series Data

In this section we examine whether the asymmetric loss function model of the previous section helps explain time series inflation outcomes in OECD countries. We focus on the role of the conditional variances of inflation and unemployment in explaining inflation trends, and downplay the role of the KPBG inflation bias.[4]

In section 3.1 we re-examine the existing empirical literature on the role of asymmetric preferences in unemployment in understanding time series trends in inflation. We argue that findings of a statistically significant relationship between inflation and the conditional variance of unemployment are likely contaminated by a spurious regression problem. In section 3.2 we attempt to correct for this problem by re-estimating the model in first differences. We also examine the case where central bank's preferences are asymmetric in inflation.

## 3.1 Revisiting Existing Results

Ruge-Murcia (2004) formulates a model of inflation rate determination when the central banker targets the natural rate of unemployment but has asymmetric preferences in unemployment.[5] The model's reduced form solution for the inflation rate time series is the linear regression equation:

$$\pi_t = a + b \cdot \sigma^2_{u,t} + \epsilon_t \tag{3.9}$$

where $\pi_t$ is the inflation rate in period t, $\sigma^2_{u,t}$ is the conditional variance of the unemployment rate in period t, and $\epsilon_t$ is a white noise error term. Asymmetric preferences imply that the

[4]If the NAIRU is time varying, the equilibrium inflation rate will mimic any trends in the NAIRU, as the inflation bias is higher for higher NAIRUs. Ireland (1999) shows that this mechanism may help explain long term trends in U.S. inflation rates. Doyle & Falk (2004) argue that this conclusion does not extend to other OECD countries.

[5]This corresponds to a special case of the model in Section 2, where there is no Kydland-Prescott Barro-Gordon bias (i.e. $k = 1$) and the central bank's loss function is asymmetric in unemployment ($\gamma \neq 0$), but not inflation ($\alpha \to 0$).

parameter $b$ will be nonzero, being positive (negative) if the central bank prefers deviations below (above) the natural rate.

Ruge-Murcia (2004) fits model (3.9) to quarterly time series data for Canada (1961:1-1999:2), France (1970:1-1999:2), Italy (1970:1-1999:2), the United Kingdom (1960:1-1999:2) and the United States (1960:1-1992). The inflation rate is measured as the (annualized) percentage change in the GDP deflator. The unemployment rate is the quarterly average civilian unemployment rate. The conditional variance of the unemployment rate is estimated from a GARCH(1,1) formulation of the first-difference of the unemployment rate.[6, 7]

The point estimates of $b$ are all positive, which is consistent with asymmetric preferences favoring unemployment rates below the natural rate. However, the estimated $b$'s are not statistically significant for Canada, Italy, and the United Kingdom. They are significant only for France and the United States. Ruge-Murcia (2004, p.105) concludes that "estimates of the reduced-form parameters support this hypothesis for the United States and France, but not for Canada, Italy, or the United Kingdom." Ruge-Murcia (2004, p.103) also notes that "the countries for which $b$ is statistically different from zero are also the countries for which the conditional variance of unemployment is unemployment is the most persistent."

Thus, Ruge-Murcia's (2004) time series regressions of the inflation rate on the (estimated) conditional variance of the unemployment rate provide evidence of a statistically significant relationship only for those countries with highly persistent conditional variances.[8] This is not a prediction of the theory. Perhaps it is a coincidental result. Or, it could reflect a spurious regression problem resulting from the persistence of both the dependent and independent variables in the regression. We pursue this possibility below.[9]

[6] Japan was initially included in the sample but was excluded because its unemployment rate did not display evidence of conditional heteroskedasticity.

[7] Ruge-Murcia actually estimated the model of the unemployment rate and its conditional variance jointly with model (3.9) using quasi-maximum likelihood. He notes, and our results confirm, that the results for model (3.9) obtained by a two-step procedure in which OLS is applied to (3.9) after the conditional variances are estimated in the first step provide essentially the same results.

[8] This result is not specific to the countries studied by Ruge-Murcia. We have confirmed that it holds in the larger set of countries we examine in the section 3.2.

[9] It is worth noting that Ruge-Murcia's test also relies on the assumption that the inflation rate is conditionally homoskedastic, an assumption that our subsequent tests suggest is invalid for most countries in his sample.

Consider the GARCH(1,1) model of the innovation in the unemployment series:

$$\begin{aligned} w_t &= v_t\sqrt{h_t} \\ h_t &= \mu + \alpha w_{t-1}^2 + \omega h_{t-1} \end{aligned} \tag{3.10}$$

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 $\alpha + \omega$. Therefore, $\alpha + \omega$ measures the persistence of the conditional variance. In the limit, $\alpha + \omega = 1$ which defines the IGARCH(1,1) model. The values of $\alpha + \omega$ reported by Ruge-Murcia (2004) for the United States and France are 0.87 and 0.92, respectively. The values of $\alpha + \omega$ for Canada, Italy, and the United Kingdom are 0.49, 0.70, and 0.36, respectively.

Not only are the conditional variances of the unemployment rates for the United States and France persistent, but so are their inflation rates, with first-order autocorrelation coefficients equal to 0.87 and 0.80. In fact, the same unit root tests that leads Ruge-Murcia (2004) to conclude that unemployment rates are I(1) processes, lead to the same conclusion for inflation rates in these countries. (See, for example, Ireland (1999), or Doyle & Falk (2004).)

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:

$$y_t = \rho y_{t-1} + u_t, \qquad u_t \text{ i.i.d.N}(0,1), \qquad 0 < \rho \leq 1. \tag{3.11}$$

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:

$$\begin{aligned} w_t &= v_t\sqrt{h_t} \\ h_t &= \mu + \alpha w_{t-1}^2 + \omega h_{t-1} \\ v_t &\sim \text{i.i.d. N}(0,1) \\ \mu &= 1, \quad 0 \leq \alpha \leq 1, \quad 0 \leq \omega < 1, \quad \alpha + \omega \leq 1 \end{aligned} \tag{3.12}$$

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. The actual rejection rates, compiled for 1000 simulations and sample sizes 100, 500, and 2000, are reported in Table 1. 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. Ruge-Murcia's point estimates of $\alpha$ and $\omega$ 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 (3.9) the coefficient $b$ 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!

Our conclusion is that existing findings of a statistically significant correlation between the conditional variance of unemployment and inflation may well be spurious. The next section of the paper represents our attempts at correcting for this problem.

### 3.2 Correcting for Persistence

Regressions of the inflation rate on its own conditional variance, or 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, including IGARCH-M regressions, 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. To this end, we collected an alternative data set made up of quarterly unemployment and CPI-based inflation rates for eighteen countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Sweden, Switzerland, United Kingdom, United States. We used the largest sample available in each country, which for most countries spans from the mid to late 1960s until 2003.[10]

We began by estimating the model of the previous section, this time in differences. Taking first differences of 3.9 gives

$$\Delta\pi_t = b\Delta\sigma_{u,t}^2 + \epsilon_t, \tag{3.13}$$

where $\Delta x_t = x_t - x_{t-1}$, and $\sigma_{u,t}^2$ is the conditional variance of the unemployment rate.[11] 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 $b$ were virtually identical and the intercept was not statistically significant in any case. Therefore, we only report the results for the regressions run without an intercept.

Following Ruge-Murcia, 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 $b$ in equation (1) 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 and if any of the p-values was less than or equal to 10-percent, we took this as evidence of possible conditional heteroskedasticity. 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

[10]The data appendix to the paper describes the data in more detail.

[11]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 output gap did not change the results substantively.

GARCH(1,1) conditional variance series for the differenced unemployment rate's innovations.

The results using the CPI as the price series are presented in Table 2.[12] We find evidence of time varying conditional heteroskedasticity in 9 out of the 18 countries. In two of these countries, the ARCH effects were not persistent, and in both of these cases the estimate of $b$ was positive, though statistically significant only in one (Austria) and at the 10% level. In the remaining 7 countries, we detected persistent ARCH effects. In four of these countries, the point estimate of $b$ was negative, but statistically significant in only two of these four countries. In the remaining three countries, the estimate of $b$ 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. There is no evidence of time variation in the conditional volatility in about half of the countries in the sample, and in about half of the countries for which ARCH effects are present the estimate of the key parameter, $b$, is of the intuitively unattractive sign. In none of the 18 countries do we find statistically significant evidence that the conditional variance of unemployment is positively related to the change in inflation, as the model would suggest.

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.[13] The model's reduced form solution in this case results in a linear equation relating $\pi$ to the conditional variance of $\pi$:

$$\pi_t = a + b \cdot \sigma^2_{\pi,t} + \epsilon_t \tag{3.14}$$

We estimate the model in first differences, as both the inflation rate and the conditional

[12] Unreported estimates using the GDP deflator as the price series for a smaller sample of countries mirror the results reported in Table 2.

[13] This corresponds to the special case of the model in Section 2, where there is no Kydland-Prescott Barro-Gordon bias (i.e. $k = 1$) and the central bank's loss function is asymmetric in inflation ($\alpha \neq 0$), but not unemployment ($\gamma \to 0$).

variance of the inflation rate exhibit high persistence. Taking first differences of 3.14 gives:

$$\Delta\pi_t = b\Delta\sigma^2_{\pi,t} + \epsilon_t, \tag{3.15}$$

which forms the basis of our estimation. As before, the results with respect to the parameter $b$ in regressions with and without an intercept were virtually identical and the intercept was not statistically significant in any case.

As in the previous case, the slope coefficient in 3.15 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 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, $b$, by fitting the first difference of inflation to the first difference of the estimated conditional variance series.[14]

Results using the CPI as the prices series are presented in Table 3. First off, we find evidence of a time varying conditional variance in inflation in nearly two thirds of countries in the sample. Of the 11 countries for which we find evidence of time-varying conditional heteroskedasticity, the ARCH effects are not persistent in 5 cases. In each of these cases, the estimate of $b$ is negative, and is statistically significant in 4 of the cases. In the remaining 6 countries, for which we find evidence on persistent ARCH effects, the estimates of $b$ are positive but not statistically significant in two countries, and negative in four countries, though not statistically significantly different from zero in two of the latter.

Table 4 estimates using the GDP data on a smaller sample of countries are similar to the

[14] 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 3.14. We found evidence of conditional heteroskedasticity in the CPI-based inflation rate for 12 of the 18 countries in the sample (Australia, Austria, Canada, Denmark, Finland, Germany, Italy, Japan, New Zealand, Norway, United Kingdom, and United States). Of these 12 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 12 of these countries displayed persistent conditional heteroskedasticity, with the sum of the estimated GARCH(1,1) coefficients exceeding 0.75. In 11 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.

results reported in Table 3. The main difference is that the ARCH effects are persistent for Australia, Japan, when the GDP deflator is used, and the slope coefficients become positive and statistically insignificant in these countries. Also, the U.S. exhibits (persistent) ARCH effects when the deflator is used, and the estimate for $b$ is both negative and statistically significant.

The results reported in Tables 3 and 4 offer mixed support for the mechanism of asymmetric preferences coupled with a time varying variance of inflation. However, at least judged by the metric of goodness of fit, our results do not translate into support for the view that this mechanism is an important determinant of changes in inflation. The $R^2$ values for the regressions reported in Tables 3 and 4 are almost uniformly low. Of the 18 cases, across the two tables, where we find evidence of time varying conditional heteroskedasticity, the $R^2$ is below 0.1 in 16 cases, lying below 0.04 in 14 cases. The exceptions are the U.K. using the CPI as the measure of prices, and Italy, using the GDP deflator, in which cases the $R^2$ values are 0.14 and 0.22 respectively. Even in this latter case, which provides the best fit, the model is explaining only about 20% of the variance in inflation changes.

Overall, we cannot reject the notion that support for a version of the model with asymmetric preferences in inflation is present in the time series data of at least some countries. However, the evidence in favor of this view as an important determinant of inflation is weak.

# 4 Evidence from a Cross Section of Countries

In this section we examine whether the model can explain inflation outcomes in cross sectional data. Essentially, we ignore the time series aspect of the problem by analyzing average inflation outcomes across countries over a given time period (1981-2002). Eliminating the time dimension from equation 2.8 leads to our baseline specification:[15]

$$\bar{\pi}_j = a + b_1\bar{u}_j + b_2\sigma_{u,j} + b_3\sigma_{\pi,j} + \epsilon_j, \tag{4.16}$$

where $\bar{\pi}_j$ is the measure of inflation in country $j$, $\bar{u}_j$ is the measure of the NAIRU in country $j$, and $\sigma_{u,j}$ , and $\sigma_{\pi,j}$ are our estimates of the standard deviations of unemployment and inflation, respectively, in country $j$.

According to the model, the NAIRU is exogenous. Furthermore, due to the inability of monetary policy to affect unemployment outcomes in rational expectations models of this sort, the actual unemployment rate only ever differs from the NAIRU due to exogenous shocks. Hence, assuming a constant NAIRU, it is consistent with the model to use the sample average unemployment rate as a proxy for $u^n$ in a linear regression. Thus, $\bar{u}_j$, our measure of the NAIRU for country $j$, is the sample average of our annual unemployment data.

There are a number of possible measures of the volatility of inflation and unemployment. Interpreted literally, the model implies that inflation should be related, more or less linearly, to the conditional variances of inflation and unemployment, rather than their standard deviations (see equation 2.8, for example). However, the data suggest a linear relationship between inflation and the standard deviations of inflation and unemployment. Since the use of standard deviations rather than variances does not change the main thrust of the results, and since the exact form of the equilibrium (equation 2.7) depends on the specific nature of the loss function assumed in any case, we report the results using standard deviations, rather than variances, as our measure of volatility in this section of the paper.

Our measure of the conditional standard deviations of unemployment and inflation is the sample average of the conditional standard deviations recovered from estimating a

[15] Throughout this section we present results based on the most general version of the model in which the Kydland-Prescott Barro-Gordon inflation bias is not ruled out and the central banker's objective function is allowed to be asymmetric in both inflation and unemployment. The results are robust to more restricted versions of the model in which either one of the asymmetries or the inflation bias (or both) is ruled out by assumption. So, for example, omitting $\bar{u}_j$ and $\sigma_{\pi_j}$ does not change the substantive results of the estimation. Versions of the empirical model that omit measures of the volatility of inflation display little explanatory power, at least as measured by $R^2$ values.

GARCH(1,1) model on the underlying annual time series data. Since, in this section, we are working off cross country data and ignoring the time series dimension of the problem, the sample standard deviation and the mean conditional standard deviation converge asymptotically. As a result, it is reasonable to use the sample standard deviations as an alternate measure of the volatility of unemployment and inflation to provide a check on the robustness of our results to the measure of volatility employed.

The last unobservable is the inflation target, $\pi^*$. We start with the simplest assumption, that all countries target the same rate of inflation. Note that the target rate of inflation is the rate of inflation that would obtain if average unemployment, the conditional variance of inflation and the conditional variance of unemployment were equal to zero. In other words, it is the intercept of equation 4.16.

The results of our cross sectional estimation are presented in Table 5. All of the results reported in Table 5 are based on OLS and use the mean of the conditional standard deviations of inflation and unemployment as the measures of volatility. All of these results are robust to the use of the sample standard deviations. We do not report results using the sample standard deviations mainly due to their strong similarity to the results reported. White's heteroskedasticity consistent robust standard errors are reported throughout Table 5.

Our sample also includes three outliers with regards to inflation[16] We have excluded these outliers from the sample for all of the regressions summarized in Table 5. The results presented are robust to the inclusion of the outlying countries, with one exception, which we discuss below.

The first column of Table 5 presents the results of the estimation of equation (4.16). The first thing to note is that the only statistically significant finding is that the average inflation rate depends directly on the standard deviation of inflation, which is positive and statistically significant at the 1% level. None of the other coefficients is statistically

[16]The outliers are Argentina, with an average annual inflation rate of 368%, Brazil, with average annual inflation of 529%, and Nicaragua, with an annual average inflation rate of 1248%. The next highest average inflation rate in the sample is 60% (Turkey).

significant at conventional significance levels. The fit of the model is unusually tight, as can be seen by the high $R^2$ value. This close fit is due to the strong relationship between average inflation and the standard deviation of inflation.

The second point of interest is that the estimate of the coefficient for the standard deviation of inflation has the "wrong," or at least less intuitive, sign. Recall (see equation 2.8 and the surrounding discussion) that a positive value of $b_3$ corresponds to a negative value of $\alpha$ (the parameter governing the asymmetry in inflation in the central banker's loss function). While the model does not place restrictions on the sign of $\alpha$, a positive relationship between the volatility of inflation and average inflation corresponds to the case where the central banker finds inflation outcomes below target *more* costly than high inflation outcomes. This is the intuitively unappealing case, and it seems unlikely that such a mechanism could be responsible for such a strong relationship as the one detected in the data here. The coefficients on average unemployment and the mean of the conditional standard deviation of unemployment do have the "correct" signs, but the estimates are not statistically significantly different from zero.

One concern with the results presented in the first column of Table 5 is the use of the raw level of inflation as the dependent variable where the model predicts that the conditional variances of unemployment and inflation affect the deviation of equilibrium inflation from target. The main concern, in a cross country regression, is that the central bank's target inflation rate might differ across countries. In this case, running the simple regression of inflation against the conditional variance of unemployment will be misleading. This is an important issue given that our data set contains some fairly heterogeneous countries and that countries' inflation targets are unobservable. We take three approaches to deal with this problem: first, we re-estimate the model using a more homogenous sample of countries, second, we try to control for differences in inflation targets by including in the regression proxy variables that are thought to affect average inflation outcomes, and finally, we re-estimate the model in differences, which eliminates any cross country differences in inflation targets to the extent that these cross country differences are constant over time.

The first approach consists simply of re-estimating equation (4.16) using a more homogenous sample of 21 OECD countries.[17] The second column of Table 5 presents the results. The results closely resemble the results reported in the first column of Table 5: the coefficient on the standard deviation of inflation is both positive and statistically significant at the 1% level, while the other coefficients are not statistically significant at conventional significance levels. Again, the close relationship between average inflation and the standard deviation of inflation results in a tight fit, as evidenced by the high $R^2$ values, though, as before, the coefficient on the mean conditional standard deviation of inflation enters with the wrong sign. As in the previous case, these results are robust to the inclusion of outliers and the use of the sample standard deviations as the measures of volatility.

The second approach requires finding proxy variables that might allow us to control for cross country differences in inflation targets. To this end, we looked at the empirical literature on cross country inflation differences to see what variables are thought to influence average inflation outcomes. Commonly used control variables are lagged average inflation, lagged real income, openness, central bank independence, and G/Y.[18]

We model these variables as forming the typical inflation rate of a country around which discretionary monetary policy then operates. We allow our control variables to influence average inflation linearly, resulting in the following estimating equation:

$$\begin{aligned}\bar{\pi}_j &= a + b_1 \cdot \bar{u}_j + b_2 \cdot \sigma_{u,j} + b_3 \cdot \sigma_{\pi,j} + c_1 \cdot CBI_j + c_2 \cdot OPEN_j \\ &\quad + c_3 \cdot G/Y_j + c_4 \cdot L(\bar{\pi}_j) + c_5 \cdot L(Y_j) + \epsilon_j, \qquad (4.17)\end{aligned}$$

where $CBI_j$ is the measure of central bank independence in country $j$, $OPEN_j$ is the measure of openness in country $j$, $(G/Y)j$ is government expenditures relative to GDP in country $j$, $L(\bar{\pi}_j)$ is lagged average inflation in country $j$, and $L(Y_j)$ is lagged real per capita

[17] The countries are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Luxembourg, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, the United Kingdom, and the United States.

[18] Campillo and Miron (1997) provide a synthesis of the literature on cross country inflation differences, incorporating a wide range of control variables common in the literature. Our control variables are taken from their paper. Detailed information concerning the construction of the data set, along with a discussion of each of these variables and why they are used in inflation regressions is contained in the data appendix to the paper.

income in country $j$. According to this specification then our estimate of the inflation target is $\hat{\pi}_j^* = c_1 \cdot CBI_j + c_2 \cdot OPEN_j + c_3 \cdot G/Y_j + c_4 \cdot L(\bar{\pi}_j) + c_5 \cdot L(Y_j)$.

Column 3 of Table 5 presents the results. As in the previous cases, the coefficient on the volatility of inflation is positive and statistically significant at the 1% level. The signs of the other coefficients have the intuitively attractive signs, but the estimates are not statistically significant from zero. Again, the $R^2$ is high.

These results are by and large robust to the use of the sample standard deviations, and the inclusion of outlier observations. The one exception is the case where the outliers are included and the conditional standard deviation is used as the measure of volatility. In this case the coefficients on average unemployment and the conditional standard deviation of unemployment are statistically significant at the 10% level. Experiments with various permutations on the set of included control variables did not change the results in any substantive way.

Our final approach to the problem of unobserved $\pi_j^*$ is to estimate the model in differences so as to eliminate any cross country differences in inflation targets. The estimating equation is:

$$\Delta\bar{\pi}_j = a + b_1 \Delta\sigma_{u,j} + b_2 \Delta\sigma_{\pi,j} + \Delta b_3 \bar{u}_j + \epsilon_j, \tag{4.18}$$

where $\Delta X_j = X_{j,t} - X_{j,t-1}$. We estimate the model using approximately decade long samples, so that $t$ is the sample period 1990-2002, and $t-1$ is the period 1980-1989. Then, $\Delta\bar{\pi}_j$, for example, is the difference between average inflation from 1980-1989 to 1990-2002. We include an intercept term to capture any common time trend in inflation targets across countries.

Column 4 of Table 5 presents the results. Again,the results closely resemble the previously reported results: the coefficient on the volatility of inflation is positive and statistically significant at the 1% level. In this case, the sign of the coefficient on the volatility of unemployment is positive, though the coefficient is not statistically significant and the sign of the coefficient on average unemployment is negative, though, again it is not statistically significant. The $R^2$ values are again high. The positive sign on the coefficient of the volatility of unemployment is not a robust result, and only occurs in the case when outliers are excluded and the means of the conditional standard deviations of unemployment and inflation are used as the measures of volatility. In all other permutations, the sign of this coefficient is negative, though statistically insignificant.

The results reported in Column 4 of Table 5 are not robust to the inclusion of the three high inflation countries to the estimation.[19] When these countries are included in the sample, the coefficient on the standard deviation of inflation remains positive,but the estimate is no longer statistically significant.

The results presented in columns 2 through 4 of Table 5 represent attempts to control for country specific differences in the intercept term in our baseline regression model. A related issue that arises because of the pooled nature of the data is the possibility that some slope parameters may differ across countries.

In particular, there is a literature that argues that the slope of the short run Phillips curve, $\lambda$ and the Barro-Gordon inflationary bias, $k$, are endogenous. The degree of central bank independence is thought to affect the inflationary bias parameter, and openness is thought to affect the slope of the short run Phillips curve (Romer (1993)). We incorporate these effects by interacting our measures of central bank independence and openness with the relevant regressor: average unemployment. Allowing for cross country slope parameters to vary in this way results in:

$$\begin{aligned} \bar{\pi}_j &= a + b_1 \cdot \bar{u}_j + b_2 \cdot \sigma_{u,j} + b_3 \cdot \sigma_{\pi,j} + c_1 \cdot CBI_j + c_2 \cdot OPEN_j + c_3 \cdot G/Y_j \\ &\quad + c_4 \cdot L(\bar{\pi}_j) + c_5 \cdot L(Y_j) + d_1 \cdot CBI_j \cdot \bar{u}_j + d_2 \cdot OPEN_j \cdot \bar{u}_j + \epsilon_j \end{aligned} \tag{4.19}$$

Column 5 of Table 5 reports the results of the estimation of equation 4.19 The results closely resemble the previously reported results: the coefficient on the volatility of inflation is positive and statistically significant at the 1% level. The sign of the coefficient on the

[19]These three countries, Argentina, Brazil, and Nicaragua, are also outliers in terms of the change in average inflation, which is equal to 664%, 808%, and 107% respectively. The majority of the countries in the sample experienced a decline in average inflation between these two periods, and the next highest value in the sample is 32% (Venezuela).

volatility of unemployment is negative in each cases, but not statistically significant, and the sign of the coefficient on average inflation is positive, but not statistically significant. The $R^2$ values are again high. The results are largely robust to the use of the sample standard deviations, and the inclusion of outlier observations, though in this last case the coefficients on average unemployment and the mean of the conditional standard deviation of unemployment are statistically significant at the 10% level.

A final issue is that if inflation is skewed then measurement error in inflation is correlated with the conditional variances. The OLS estimates of the slope then reflect correlation between inflation and the conditional variances, but also correlation between measurement error and the conditional variances (which is proportional to the skewness of the independent variable). Ruge-Murcia (2002) deals with this by using one sample period to estimate the conditional variances and a different sample period to derive average inflation.

Table 6 presents the result of a re-estimation of our base model (equation 4.16) on decade by decade data, using the lagged, rather than contemporaneous, value of $\sigma_\pi$ as the measure of inflation for the sample of 44 countries. In other words, we estimate the base model (equation 4.16) using data based on one sample period (for example 1990-2002), but base our measure of $\sigma_\pi$ on the previous period (that is, 1980-1989). Column 1 reports the results using the 1990-2002 as the base period, and column 2 reports the results where 1980-1989 is the base period (in which case $\sigma_\pi$ is based on 1970-1979 data). In both cases, the mean of the conditional standard deviation is the measure of volatility. Under this specification, none of the estimated coefficients (aside from the constant term in the 1980s regressions) are statistically significantly different from zero, and the model fits the data quite poorly.

Table 6 also reports results from the same analysis applied to a sample of 21 OECD countries. Column 3 reports the results for the 1990s, while column 4 reports the results from the 1980s. For this more homogenous group of countries, the use of lagged, instead of contemporaneous, $\sigma_\pi$ does not change the results so drastically. In this case, the coefficient on the (lagged) standard deviation of inflation is positive in both cases, statistically significant at the 1% level in both the 1990s and 1980s regression. The coefficient on the conditional standard deviation is negative and statistically insignificant for both decades, and the coefficient on average unemployment is positive in both cases but not statistically significant in either case. The fit of these equations, as measured by the $R^2$'s is not as tight as in the cases where the contemporaneous standard deviation of inflation was used, but still quite high for cross sectional estimation.

Ruge-Murcia (2002), and Cukierman & Gerlach (2003) have previously tested versions of the asymmetric preference model on cross sectional data for samples of OECD countries. Ruge-Murcia (2002) finds evidence that the volatility of unemployment, but not the volatility of inflation, contributes to average inflation using quarterly data for a sample of OECD countries in the 1990s. Cukierman and Gerlach (2003) study the case where preferences are asymmetric only in unemployment, and there is no KPBG bias, using sample standard deviations as their measure of the volatility of unemployment. They find some evidence that the volatility of unemployment contributes to average inflation prior to 1985, but not afterwards, a result which they argue is likely due to increased central bank independence in the latter period. Their results are consistent with our findings, that the volatility of unemployment does not contribute to inflation post 1980.

## 5 Conclusion

Overall, the results presented in this paper provide, at best, weak support for the view that the interaction of asymmetric central bank preferences and the volatilities of inflation and unemployment are important determinants of inflation. In particular, neither the time series nor cross sectional data supports the view that the volatility of unemployment helps explains inflation. Estimated coefficients on variables measuring this volatility are consistently not statistically significantly different from zero, and frequently posses the wrong sign.

The results concerning the relevance of asymmetric preferences in inflation are more difficult to interpret. The time series results provide support for the view that the interaction between a changing volatility of inflation and a central bank endowed with an asymmetric loss function affects inflation in some OECD countries. However, this mechanism appears to explain only a small portion of the variance of changes in inflation in these countries.

Furthermore, while incorporating measures of the volatility of inflation seems to help explain cross country differences in average inflation, it is not at all clear, however, whether this correlation is due to asymmetric central bank preferences. While the model does not place restrictions on the sign this correlation, our results corresponds to the case where the central banker finds inflation outcomes below target *more* costly than high inflation outcomes. This seems to be an unlikely description of central banker's preferences.

## Data Appendix

### Quarterly Data for Section 3

We collected data on quarterly unemployment and CPI-based inflation rates for eighteen countries, listed below. The data are taken from the OECD's Main Economic Indicators database. Series numbers for both the unemployment and CPI data are listed in the table below:

| Country | Sample | Series # u | Series # CPI |
|---|---|---|---|
| Australia | (1966:3-2003:3) | 544113DSA | 545241K |
| Austria | (1964:1-2003:3) | 704115DSA | 705241K |
| Belgium | (1979:3-2003:3) | 224111DSA | 225241K |
| Canada | (1964:1-2003:3) | 444113DSA | 445241K |
| Denmark | (1970:1-2003:3) | 304111DSA | 305241K |
| Finland | (1964:1-2003:3) | 644113DSA | 645241K |
| France | (1967:4-2003:3) | 114113DSA | 145241K |
| Germany | (1964:1-2002:4) | 134111DSA | 125241K |
| Ireland | (1983:1-2003:3) | 284113DSA | 285241K |
| Italy | (1964:1-2003:3) | 164113DSA | 165241K |
| Japan | (1964:1-2003:3) | 464113DSA | 465241K |
| Netherlands | (1988:2-2003:3) | 184111D | 185241K |
| New Zealand | (1985:4-2003:2) | 594113DSA | 595241K |
| Norway | (1972:1-2003:3) | 584113DSA | 585241K |
| Sweden | (1970:1-2003:3) | 604113DSA | 605241K |
| Switzerland | (1983:1-2003:3) | 684111DSA | 685241K |
| United Kingdom | (1964:1-2003:3) | 264111DSA | 265241K |
| United States | (1964:1-2003:3) | 424113DSA | 425241K |

### Annual Data for Section 4

We collected annual data from 1981-2002 on unemployment and CPI-based inflation for 47 countries: Argentina, Australia, Austria, Barbados, Belgium, Brazil, Canada, Chile, Colombia, Costa Rica, Cyprus, Denmark, El Salvador, Finland, France, Germany, Greece, Iceland, Ireland, Israel, Italy, Jamaica, Japan, Korea, Luxembourg, Malaysia, Malta, Mexico, Netherlands, New Zealand, Nicaragua, Norway, Pakistan, Panama, Philippines, Portugal, Singapore, Spain, Sweden, Switzerland, Thailand, Trinidad and Tobago, Turkey, United

Kingdom, United States, Uruguay, and Venezuela.

The CPI data were taken from the IMF International Financial Statistics Database. The unemployment data were taken from the same IMF database, and are supplemented with additional information taken from the ILO Labor Statistics Database.

We augmented this data with a number of other variables, including lagged average inflation, real income, openness, central bank independence, and G/Y used in the estimation of equations (4.17) and (4.19) to control for potential differences in inflation targets across countries. What follows is a description of the construction of these variables, along with a brief explanation of why these variables are thought to matter for inflation outcomes.

Lagged average inflation is measured using average CPI-based inflation for the years 1970-1980. The underlying CPI data is again taken from the IMF International Financial Statistics database. This is included to capture a number of possibilities: it may capture unmeasured country specific factors that affect inflation outcomes, so that high past inflation may suggest high future inflation; countries that have experienced high inflation in the past may find, after experiencing the costs of inflation, that reform is needed, making high inflation in the future less likely; or countries that have experienced high inflation historically may have invested in technologies or institutions to reduce the costs of inflation, making high inflation in the future less objectionable.

Real income is measured using the log of per capital real GDP in 1980, using the RGDPCH series from the Penn World Tables. This is included in case there are any rich country specific effects on inflation. For example, rich countries may have better developed tax and financial systems, which would reduce reliance on an inflation tax. Alternately, rich countries may be more readily able to afford technologies which reduce the costs of inflation.

Openness is measured as the average of the ratio of Exports plus Imports to GDP from 1981-2002. It has been argued (by Romer (1993), for example) that openness affects inflation by making it easier for the central bank to commit to low inflation. The idea is that inflationary policy leads to real exchange rate depreciations, which are more costly in countries that trade more.

There are a couple of viable measures of Central Bank Independence available in the literature. We use two. The first is due to Cukierman and Webb (1995), which bases a measure of Central Bank Independence on the probability that a central banker will be replaced shortly after a change in government, and covers the period for 1980-1989. This data covers most, but not all of, the countries in our sample. We supplement it with data from Sturm and de Haan (2001) which extend this measure to a large set of developing countries, for the period 1980-1998. A more independent central bank is thought to have a greater ability to commit to inflation than one that is more subject to political control.

Government expenditures to GDP (G/Y) is measured as the average of government expenditures to GDP, with the data taken from the IMF's International Financial Statistics database. This is included to capture the idea that governments experiencing difficulty financing their expenditures may inflate their currencies to enjoy seignorage revenues. Alternative, and probably better, measures include the ratio of the government budget deficit or debt to GDP, but this data was not readily available for many of the countries in our sample.

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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

| $(\alpha, \omega)$ | $\rho = 0.00$ | $\rho = 0.40$ | $\rho = 0.80$ | $\rho = 0.90$ | $\rho = 0.95$ | $\rho = 0.99$ | $\rho = 1.00$ |
|---|---|---|---|---|---|---|---|
| (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

| $(\alpha, \omega)$ | $\rho = 0.00$ | $\rho = 0.40$ | $\rho = 0.80$ | $\rho = 0.90$ | $\rho = 0.95$ | $\rho = 0.99$ | $\rho = 1.00$ |
|---|---|---|---|---|---|---|---|
| (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 |

Notes: This table contains 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. The AR(1) process has the form $y_t = \rho y_{t-1} + u_t$, $u_t \sim$ i.i.d.$N(0,1)$. The GARCH(1,1) process has the form $w_t = v_t\sqrt{h_t}$, $h_t = 1 + \alpha w_{t-1}^2 + \omega h_{t-1}$ where $v_t \sim$ i.i.d.$N(0,1)$. Initial values were set equal to zero and 500-period burn-in periods were used.

Table 2. Change in Inflation on the Change in the Conditional Variance of Unemployment

| Country | Sample Period | LM-test p-values | $\hat{b}$ | $t(b)$ | $\alpha + \omega$ |
|---|---|---|---|---|---|
| Australia | 1966:3-2003:3 | 0.20 | - | - | - |
| Austria | 1964:1-2003:3 | **0.10** | 0.27 | 1.83* | 0.124 |
| Belgium | 1979:1-2003:3 | **0.05** | 0.00 | 0.77 | 0.800 |
| Canada | 1964:1-2003:3 | **0.00** | 0.00 | 0.32 | 0.671 |
| Denmark | 1970:1-2003:3 | **0.00** | 0.00 | 0.07 | 0.996 |
| Finland | 1964:1-2003:3 | 0.26 | - | - | - |
| France | 1967:1-2003:3 | 0.42 | - | - | - |
| Germany | 1964:1-2002:4 | **0.06** | -0.04 | -0.94 | 0.715 |
| Ireland | 1983:1-2003:3 | **0.05** | -0.07 | -0.55 | 0.910 |
| Italy | 1964:1-2003:3 | 0.11 | - | - | - |
| Japan | 1964:1-2003:3 | 0.30 | - | - | - |
| Netherlands | 1988:2-2003:3 | 0.42 | - | - | - |
| New Zealand | 1985:4-2003:2 | 0.13 | - | - | - |
| Norway | 1972:1-2003:3 | 0.56 | - | - | - |
| Sweden | 1970:1-2003:3 | **0.00** | -0.10 | -1.98** | 0.903 |
| Switzerland | 1983:1-2003:3 | 0.43 | - | - | - |
| UK | 1964:1-2003:3 | **0.00** | 0.03 | 0.44 | 0.301 |
| US | 1964:1-2003:3 | **0.00** | -0.02 | -2.01** | 0.897 |

Notes:

- The LM test column provides p-values for the Engle (1982) LM test for neglected ARCH effects. The LM test is based on the distribution of the $T \cdot R^2$ statistic from the regression of the squared residuals from an AR(p) model of the first-difference of the unemployment rate on a constant and one, two, three, and four lags of the squared residuals. The value of p, for the AR(p), was selected by applying the AIC. 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.

- The $t(b)$ column provides the t-ratio for the slope coefficient in the regression of the first-difference of the inflation rate on the first-difference of the conditional variance of the unemployment rate.

- The $\alpha + \omega$ column provides the sum of the estimates slope coefficients in the GARCH(1,1) model of the errors from the AR(p) model of the first-difference of the unemployment rate.

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

Table 3. Change in Inflation on the Change in the Conditional Variance of Inflation (CPI)

| Country | Sample Period | LM-test p-values | $\hat{b}$ | $t(b)$ | $\alpha + \omega$ |
|---|---|---|---|---|---|
| Australia | 1966:3-2003:3 | **0.01** | -0.15 | -2.09** | 0.349 |
| Austria | 1964:1-2003:3 | **0.00** | -0.31 | -2.87*** | 1.000 |
| Belgium | 1979:1-2003:3 | **0.10** | -0.28 | -0.90 | 0.304 |
| Canada | 1964:1-2003:3 | **0.01** | -0.35 | -1.35 | 0.730 |
| Denmark | 1970:1-2003:3 | **0.04** | 0.10 | 0.50 | 0.929 |
| Finland | 1964:1-2003:3 | **0.01** | -0.32 | -0.50 | 0.817 |
| France | 1967:1-2003:3 | 0.18 | - | - | - |
| Germany | 1964:1-2002:4 | **0.05** | -87.97 | -2.03** | 0.396 |
| Ireland | 1983:1-2003:3 | 0.21 | - | - | - |
| Italy | 1964:1-2003:3 | **0.00** | 0.16 | 0.91 | 1.000 |
| Japan | 1964:1-2003:3 | **0.01** | -0.25 | -3.37*** | 0.372 |
| Netherlands | 1988:2-2003:3 | 0.22 | - | - | - |
| New Zealand | 1985:4-2003:2 | 0.55 | - | - | - |
| Norway | 1972:1-2003:3 | **0.00** | -0.05 | -0.34 | 1.000 |
| Sweden | 1970:1-2003:3 | 0.71 | - | - | - |
| Switzerland | 1983:1-2003:3 | 0.82 | - | - | - |
| UK | 1964:1-2003:3 | **0.00** | -0.52 | -4.94*** | 0.363 |
| US | 1964:1-2003:3 | 0.14 | - | - | - |

Notes:

- The LM test column provides p-values for the Engle (1982) LM test for neglected ARCH effects. The LM test is based on the distribution of the $T * R2$ statistic from the regression of the squared residuals from an $AR(p)$ model of the first-difference of the inflation rate on a constant and one, two, three, and four lags of the squared residuals. The value of p, for the AR(p), was selected by applying the AIC. 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.
- The $t(b)$ column provides the t-ratio for the slope coefficient in the regression of the first-difference of the inflation rate on the first-difference of the conditional variance of the inflation rate.
- The $\alpha + \omega$ column provides the sum of the estimated ARCH and GARCH coefficients in the GARCH-M model of the inflation rate on the conditional variance of the inflation rate. The model was fit by QMLE using a GARCH(1,1) specification.

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

Table 4. Change in Inflation on the Change in the Conditional Variance of Inflation (GDP Deflator)

| Country | Sample Period | LM-test p-values | $\hat{b}$ | $t(b)$ | $\alpha + \omega$ |
|---|---|---|---|---|---|
| Australia | 1966:3-2003:3 | **0.00** | 0.03 | 1.96* | 1.000 |
| Canada | 1964:1-2003:3 | **0.03** | 0.10 | 0.70 | 0.990 |
| France | 1970:2-1998:4 | **0.03** | -0.03 | -0.35 | 1.000 |
| Germany | 1964:1-2002:4 | **0.00** | -0.25 | -3.92*** | 0.961 |
| Italy | 1964:1-2003:3 | **0.00** | 6.36 | 6.61*** | 0.002 |
| Japan | 1964:1-1999:4 | **0.01** | 0.04 | 0.94 | 0.962 |
| Switzerland | 1983:1-2003:3 | 0.24 | - | - | - |
| UK | 1964:1-1998:4 | **0.01** | -0.02 | -1.23 | 0.837 |
| US | 1964:1-2003:3 | **0.02** | -0.52 | -1.95* | 0.953 |

Notes:

- The LM test column provides p-values for the Engle (1982) LM test for neglected ARCH effects. The LM test is based on the distribution of the $T * R2$ statistic from the regression of the squared residuals from an $AR(p)$ model of the first-difference of the inflation rate on a constant and one, two, three, and four lags of the squared residuals. The value of p, for the AR(p), was selected by applying the AIC. 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.
- The $t(b)$ column provides the t-ratio for the slope coefficient in the regression of the first-difference of the inflation rate on the first-difference of the conditional variance of the inflation rate.
- The $\alpha + \omega$ column provides the sum of the estimated ARCH and GARCH coefficients in the GARCH-M model of the inflation rate on the conditional variance of the inflation rate. The model was fit by QMLE using a GARCH(1,1) specification.

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

Table 5. Cross Sectional Estimation

| | Base | OECD | $\pi^*$ Proxies | Differenced | Slope Interactions |
|---|---|---|---|---|---|
| Constant | 1.06 | -0.07 | 7.14 | -1.59 | 7.66 |
| (std. error) | (0.95) | (0.26) | (5.56) | (1.47) | (5.60) |
| *p-value* | *0.27* | *0.78* | *0.21* | *0.61* | *0.18* |
| $\sigma_u$ | -1.17 | -0.11 | -0.25 | 0.03 | -0.28 |
| (std. error) | (1.07) | (0.20) | (1.35) | (0.02) | (1.23) |
| p-value | *0.28* | *0.58* | *0.85* | *0.13* | *0.82* |
| $\sigma_\pi$ | 0.62*** | 0.75*** | 0.62*** | 0.81*** | 0.61*** |
| (std. error) | (0.11) | (0.03) | (0.07) | (0.02) | (0.07) |
| p-value | *0.00* | *0.00* | *0.00* | *0.00* | *0.00* |
| u | 1.32 | 0.14 | 0.43 | -0.67 | 0.58 |
| (std. error) | (1.11) | (0.19) | (1.38) | (0.40) | (1.26) |
| p-value | *0.24* | *0.47* | *0.76* | *0.10* | *0.65* |
| CBI | - | - | 0.07 | - | -5.21 |
| (std. error) | | | (3.53) | | (5.16) |
| Open | - | - | -2.13** | - | -0.26 |
| (std. error) | | | (0.84) | | (1.27) |
| G/Y | - | - | -60.34** | - | -56.04** |
| (std. error) | | | (25.38) | | (24.94) |
| $L(\pi)$ | - | - | 0.00 | - | -0.02 |
| (std. error) | | | (0.02) | | (0.02) |
| $L(Y)$ | - | - | 0.57 | - | 0.38 |
| (std. error) | | | (0.71) | | (0.75) |
| CBI$\cdot u$ | - | - | - | - | 0.77 |
| (std. error) | | | | | (0.69) |
| Open$\cdot u$ | - | - | - | - | -0.33 |
| (std. error) | | | | | (0.21) |
| $R^2$ | 0.90 | 0.97 | 0.94 | 0.97 | 0.94 |

* = significant at the 10% level, ** = significant at the 5% level,
*** = significant at the 1% level

Table 6. Cross Sectional Estimation: Lagged $\sigma_\pi$

| | Full Sample | | OECD Countries | |
|---|---|---|---|---|
| | 1990s | 1980s | 1990s | 1980s |
| Constant | -0.27 | 10.92* | 1.44** | -2.87** |
| (std. error) | (4.35) | (6.16) | (0.52) | (1.34) |
| *p-value* | *0.95* | *0.08* | *0.01* | *0.05* |
| $\sigma_u$ | -0.02 | -0.06 | 0.01 | -0.03** |
| (std. error) | (0.05) | (0.06) | (0.00) | (0.01) |
| *p-value* | *0.74* | *0.34* | *0.11* | *0.03* |
| L($\sigma_\pi$) | 0.16 | 0.16 | 0.18*** | 0.96*** |
| (std. error) | (0.12) | (0.14) | (0.04) | (0.16) |
| *p-value* | *0.19* | *0.25* | *0.00* | *0.00* |
| u | 0.88 | 0.92 | -0.12 | 0.39 |
| (std. error) | (1.23) | (1.24) | (0.08) | (0.25) |
| *p-value* | *0.43* | *0.46* | *0.17* | *0.14* |
| $R^2$ | 0.17 | 0.08 | 0.66 | 0.71 |

* = significant at the 10% level, ** = significant at the 5% level,
*** = significant at the 1% level