# Is the New Keynesian Explanation of the Great Dis-Inflation Consistent with the Cross Country Data?*

Matthew Doyle†
Jean-Paul Lam‡

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**Abstract**

A leading explanation of long run U.S. inflation trends attributes both the fall of inflation in the 1980s and the subsequent years of low and stable inflation to well run monetary policy pinning down inflationary expectations. Most other OECD economies experienced a similar rise and fall of inflation, as well as subsequent low and stable inflation over the same period. This observation has been under-explored in the literature. In this paper we exploit the international dimension of the fall of inflation to investigate the hypothesis that good monetary policy is responsible for recent inflation outcomes. Our results suggest that this theory is not compatible with the cross country data.

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†Department of Economics, University of Waterloo, Waterloo ON, Canada, N2L 3G1. Email: m6doyle@uwaterloo.ca.

‡Department of Economics, University of Waterloo, and Rimini Center for Economic Analysis. Email, Lam: jplam@uwaterloo.ca.

# 1 Introduction

U.S. inflation was low in the early 1960s and rose through the late 1960s and 1970s before falling through the 1980s, and remaining low thereafter. There is an extensive literature that attempts to explain why *U.S.* inflation rose in the 1970s and then fell in the 1980s.[1] The most prominent theory may be the one suggested by New Keynesian theory and put forth by Clarida, Galí and Gertler (2000), which gives a key role to agents' expectations of inflation. A major problem in this literature, however, is that all major competing theories are broadly consistent with the U.S. data, so that debates often boil down to researchers' beliefs over the plausibility of various identification schemes and refinements in econometric technique.[2] These debates are difficult to resolve in the absence of new data.

The pattern of rising inflation in the 1970s, falling inflation in the 1980s and low and stable inflation thereafter was not isolated to the U.S., however. A similar pattern occurred across OECD countries.[3] A common pattern to OECD inflation outcomes strongly suggests a common cause, which implies that a successful explanation of the so-called Great Inflation should apply *across* OECD countries. We exploit this observation by employing a multi-country approach to evaluate the hypothesis that recent years of low and stable inflation are the result of good monetary policy pinning down inflationary expectations.

While this, New Keynesian explanation concerns both the rise and fall of inflation, we focus on the more recent period. This is for two main reasons: i) describing monetary policy via a monetary policy reaction function is a common practice in this literature, and applying this approach across OECD countries is more defensible in recent years than in earlier decades, and ii)

[1] Clarida, Galí, and Gerlter (2000), Sargent (1999), Cogley and Sargent (2005, 2001), Sims and Zha (2006), Orphanides (2004, 2003, 2002), Ireland (1999), Primiceri (2006).

[2] See for example the exchange between Benati and Surico (2009) and Canova (2007), or Cochrane's (2007) critique of the literature.

[3] Previous authors have come to a similar conclusion. For example, see Rogoff (2003), Ciccarelli and Mojon (2008), and Mumtaz and Surico (2008), Monacelli and Sala (2009), and Doyle & Falk (2008, 2010).

the theory suggests that the rise of inflation was due to indeterminacy caused by poor monetary policy, which allowed for self fulfilling expectations of high inflation. Identification of monetary policy rule parameters is more difficult when multiple equilibria are possible, so we limit our results to the more recent period where we have more confidence about what our empirical results are measuring.

Our paper starts out by documenting the key observation that low frequency movements in OECD inflation rates are strongly correlated across countries. In addition to presenting simple plots of inflation, we use the dynamic correlation measure of Croux, Forni, and Reichlin (2001) to measure the correlation between time series in the frequency domain. Using this approach we show that OECD inflation rates are highly correlated with each other, with the strongest correlation at low frequencies. In other words, the results confirm that the rise and fall of inflation was indeed an OECD-wide phenomenon.

We then turn to the main question of the paper, which is whether or not the New Keynesian theory of monetary policy improvements can explain the OECD-wide fall of inflation and subsequent low inflation. This view asserts that errors in monetary policy allowed a rise in inflation in the 1970s because central bankers, in violation of the so-called Taylor principle, failed to increase *real* interest rates in the face of rising inflationary expectations, thereby validating those expectations. The inflationary episode ended in the 1980s when policy makers adopted policies that reacted in a more contractionary way in the face of inflationary expectations. The subsequent adherence to these policies resulted in a long period of low and stable inflation that has lasted from the early or mid 1980s to the present.

We test this theory by estimating a monetary policy reaction function based on the widely used Taylor rule (Taylor (1993)) that relates central banks' nominal interest rate decisions to an output gap and a measure of expected inflation. We follow Clarida, Galí and Gertler (2000) by using GMM to estimate policy reaction functions across OECD countries for the more recent period, when inflation has been either falling or low and stable. We ask whether the estimated parameters are in the region that New Keynesian theory suggests they must be in order to not accommodate inflationary expectations.

We focus on the most recent period for two reasons. First, it is more reasonable to apply a Taylor rule framework to recent data as the conduct of monetary policy has converged more closely on a common framework. Prior to the early 1980s, policy makers used a number of instruments, price and wage controls for example, as tools of monetary policy. Institutional arrangements concerning monetary policy, in particular the use of short run nominal interest rates as a main tool of policy, are much better described by a Taylor type policy reaction function in the period after the early 1980s than before. Furthermore, from an empirical perspective, standard Taylor-type policy rules describe monetary policy in other OECD countries about as well as in the U.S. for the post-1980 period.[4]

Second, and more importantly, we look at the more recent period because of the problem of accurately and properly identifying monetary policy parameters in forward looking models, particularly when using one equation or partial information approaches, as noted by Cochrane (2007). Recent research suggests that when the Taylor rule parameters are in the determinacy region, where monetary policy does not accommodate inflationary expectations, single equation estimation methods can identify the policy rule parameters, as long as the economy exhibits inflation and/or output persistence (Carillo (2008)).[5] Less is known about how to identify monetary parameters in the indeterminacy region, where policy accommodates inflationary expectations.

Carillo (2008) uses a Monte Carlo approach to evaluate the ability of single equation approaches to accurately identify monetary policy rule parameters. For simplicity, however, he restricts attention to the case of backwards looking policy reaction functions, that express interest rate decisions as a function of contemporaneous values of inflation and the output gap. Our objective in

[4]See Clarida, Galí and Gerlter (1998), for example.

[5]Actual data displays both kinds of persistence, and research extensions of New Keynesian theory focuses on building models consistent with these features of the data. For examples, see McCallum and Nelson (1999), Erceg, Henderson and Levin (2000), Mankiw and Reis (2002), and Christiano, Eichenbaum and Evans (2005).

this paper, however, requires us to estimate a *forward looking* policy rule in which monetary policy responds to expected future inflation. We therefore first employ Carillo's approach to show that a GMM approach can identify the policy parameters in a forward looking policy reaction function.

We then conduct our test of the theory by estimating monetary policy reaction functions for the recent period.[6] If the theory is correct, and the parameters are in the determinacy region, then our Monte Carlo results suggest that we are able to identify these parameters correctly. The restriction we test is that policy satisfies the Taylor principle, which says that policy should respond to increases in expected inflation by increasing the nominal interest rate by more than the increase in expected inflation, so that the real interest rate rises. In other words, that the coefficient on expected inflation in the monetary policy rule exceeds one.

We view our approach as representing a fairly weak test of the theory, for three reasons. First, we are only testing the implications of the theory concerning the fall of inflation, rather than requiring it to explain both the rise and the fall. Also, when monetary policy parameters are in the indeterminacy region the economy may produce identical dynamics to the case where these parameters are in the determinacy region. In this case, it may appear that monetary policy satisfied the Taylor principle when in fact it did not. Finally, when trend inflation is positive, satisfying the Taylor principle requires increasing nominal interest rates by substantially more than the increase in expected inflation (i.e. the coefficient on expected inflation in the monetary policy rule is substantially larger than one). This effect is particularly pronounced when the policy reaction function is forward looking. In this paper, we assume that trend inflation is zero, thus making it much easier to satisfy the Taylor principle compared to cases where trend inflation is positive.[7]

In spite of employing a relatively weak test of the theory, our main finding is

[6] We employ two approaches to determine where the recent period starts: i) we use turning points in inflation identified by previous researchers, and ii) we use a common date thought to correspond with changes in the U.S. monetary policy regime.

[7] Coibion and Gorodnichenko (2010).

that the there is little evidence that the Taylor principle was satisfied for most countries in the recent period. In fact, we find that the theory fits the data for the U.S., but only for two or three of the other 13 countries in our sample (depending on the method of estimation used). We view the combination of a relatively lax test and little evidence in favor of the theory as strongly suggestive of the implication that the theory cannot explain the fall of inflation observed in OECD countries.

The paper proceeds as follows: in Section 2 we document inflation trends in 14 OECD countries, and argue that they share common features. In Section 3 we present the empirical and theoretical framework along with our discussion of the identification problem. In Section 4 we present the main results. In Section 5 we offer concluding comments.

# 2 Inflation Trends

In this section we argue that inflation patterns in OECD countries over the past 4 decades closely mirror the well known patterns of U.S. inflation. Our measure of inflation is the annualized quarterly percentage change in the Consumer Price Index (as reported in the OECD's Main Economic Indicators).[8] We present the data for a sample of 14 countries.

Figures 1-3 plot annualized inflation rates derived from quarterly CPI data for 14 countries. While the pattern of inflation is visible in the raw data, we present the 9 quarter centered moving average of inflation to smooth out the higher frequency movements in the data. To facilitate comparison, each figure also includes the moving average of U.S. inflation (the dotted line).

The main observation for the figures is that inflation starts out relatively low in the early 1960s in all countries. This is followed by a period of rising inflation lasting until the late 1970s or early 1980s. After this period of rising, inflation rates then fall until the present, and are generally as low or lower by

[8]The observation that inflation trends are common across countries is robust to the use of the quarterly GDP deflator as the price series for those countries for which a sufficiently long sample of quarterly national accounts data is available.

the end of the 1990s than they were in the early 1960s.

The main exceptions to the pattern are Germany, Japan and Switzerland. Each of these countries did experience a rise of inflation in the early to mid 1970s. However, inflation quickly fell back to low levels in each of these countries. These countries represent some cross sectional variation in inflation trends. Even in these countries, however, the more recent period is one of low and stable inflation.

## 2.1 Dynamic co-movements

While figures can be suggestive, their interpretation leaves much to the discretion of the viewer. To confirm that our interpretation is valid, we conduct a more formal analysis. We follow the literature by treating inflation as an I(0) process. This makes documenting a common pattern in inflation rates more complicated as it is difficult to talk about trends with respect to stationary series. In order to address this issue, we employ a measure of dynamic co-movements due to Croux, Forni and Reichlin (2001).

Correlation is often used in the literature on business cycles to measure co-movement between time series. Correlation measures, however, are static and do not capture the dynamics in the co-movement between different time series. Moreover measures of static correlation do not discriminate between co-movements at different frequencies, thus fail to establish whether co-movements are driven by high frequency components of the data or by a common trend.

Croux, Forni and Reichlin (2001) propose a measure of dynamic co-movement between time series which they have labeled dynamic correlation in the bivariate case, and cohesion in the case of more than two time series. Their measure of dynamic correlation is a study of the co-movements of different time series by frequency. We use this technique to describe the dynamic correlation between inflation in the U.S and the rest of the OECD countries. We are particularly interested in the correlation at low and business cycles frequencies.

The dynamic correlation measures the correlation between the $x_t$ and $y_t$ at different frequencies. The dynamic correlation measure is given by:

$$\rho_{xy}(\omega) = \frac{c_{xy}(\omega)}{\sqrt{s_x(\omega)s_y(\omega)}} \tag{2.1}$$

where $s_x(\omega)$, $s_y(\omega)$ denote the power spectrum of $x_t$ and $y_t$ respectively and $c_{xy}$ is the real part of the co-spectrum between $x$ and $y$. Similar to a classical correlation coefficient, this dynamic correlation measure takes values between -1 and 1.

We compute the dynamic correlation of inflation between the U.S. and the other OECD countries in our sample and these are shown in Figures 4-6. The figures plot the correlation of national inflation to U.S. inflation at different frequencies. The dotted vertical lines in the figures correspond to business cycles frequencies commonly defined between 6 ($\pi/3$) and 32 ($\pi/16$) quarters. Low frequency, or long-run, movements are to the left of the frequency line $\pi/16$.[9] The section that is to the right of frequency $\pi/3$ describes short-run (mostly seasonality) co-movements. We do not pay any attention to this part of the spectrum.

The main result of these tests is that dynamic correlation peaks at low frequency, that is within the frequency band $[0, \pi/16]$, for all countries except Switzerland and Japan. The dynamic correlation takes its highest value of around 0.7 at very low frequencies. In most cases, the peak is reached near the frequency zero. This indicates that low frequency, or long run, movements in inflation across OECD countries have closely mirrored the long-run movements in U.S. inflation. At business cycles frequencies, that is at frequencies $[\pi/16, \pi/3]$, the dynamic correlation is lower for all countries. The dynamic correlation declines in most cases and is low at frequencies corresponding to 6 quarters or more.

The positive and high dynamic correlation values at low frequencies between U.S inflation and other OECD countries reveal that there is indeed a common long-run trend in OECD inflation rates. This suggests, though ob-

[9]Note that if the spectral density at frequency zero has rank one, the two processes are cointegrated.

viously does not necessitate, that there is a common cause of the rise in the 1970s and the subsequent fall in the 1980s of inflation across OECD countries.

## 3 Model Specification and Identification

Can the common inflation trends discussed in the previous section be explained by common changes in the conduct of monetary policy? We address this question by first assuming that central bank behavior has a systematic component that can be described by a relatively parsimonious monetary policy reaction function that relates monetary policy variables to macroeconomic factors. In particular, we adopt the widely used Taylor type formulation for the policy reaction function.[10] We attempt to describe the behavior of monetary policy makers by estimating the parameters of this policy reaction function.

The use of this approach has been recently criticized, perhaps most tellingly by Cochrane (2007). Cochrane argues that it is simply impossible to identify structural monetary policy parameters through using single equations approaches to estimate parameters of a Taylor rule when the underlying DGP takes the form of a New-Keynesian type model. The essential problem is that New Keynesian models determine only expected inflation rather than actual, ex-post inflation. The consequence of this is that these models generically exhibit multiple equilibria, and the monetary policy rule determines current inflation through its influence on off the equilibrium path behavior. Taylor rule parameters in these models must be chosen to rule out bubble equilibria, which then forces the actual rate of inflation to jump to the unique level which is consistent with non-bubble outcomes. Since monetary policy in these models works only through influencing off the equilibrium path behavior, it is essentially impossible to identify monetary policy parameters based on observing equilibrium outcomes.

Cochrane's analysis focuses on purely forward looking versions of the New

[10] This kind of reaction function has been shown to closely track interest rate behavior in the U.S. (Taylor 1993) as well as in other OECD countries, at least for recent decades (Clarida, Galí, and Gerlter (1998)).

Keynesian framework. These models, however, do not fit the data well. Models in this literature that fit more closely the data, tend to incorporate additional elements, such as inertia in price setting, rule of thumb behavior, and habit formation in consumption. This results in a hybrid model containing both backward and forward looking elements.[11] These models have become increasingly common in the New Keynesian literature of monetary policy because they allow the modeler to capture important aspects of the data, such as output and inflation persistence. Purely forward looking models, like those used by earlier modelers working in the New Keynesian framework, are unable to replicate these features.[12]

It is not clear that the identification problem highlighted by Cochrane for purely forward looking models carries through to the hybrid models. Essentially, in backwards or partially backwards looking models, both historical and expected inflation are determined by the model. The inertia in inflation in hybrid models limits the possibility for bubble equilibria and also implies that inflation cannot jump in response to shocks but is pinned down by its past history. Essentially, the presence of higher order dynamics in the model and data, namely persistence in inflation and/or output, opens up the possibility that structural monetary policy parameters can be recovered from observations of equilibrium outcomes.[13]

Recent work by Carillo (2008) has examined this issue. He shows that, in the case where the Taylor rule is stated in terms of contemporaneous inflation and output, the inability to identify monetary policy parameters is particular to purely forward looking models. In backward looking and hybrid models, however, single equation estimation approaches can accurately recover the structural parameters of the monetary policy rule. In what follows, we show that this conclusion can be extended to a widely used Taylor rule specification

[11] McCallum and Nelson (1999), Erceg, Henderson and Levin (2000), Mankiw and Reis (2002), and Christiano, Eichenbaum and Evans (2005).

[12] For inflation persistence, see Fuhrer and Moore (1995), Estrella and Fuhrer (2002), and Rudd and Whelan (2006). For output persistence see Fuhrer and Rudebusch (2004).

[13] This resembles what is already known for purely backward looking models: higher order dynamics on the regressors or instruments are necessary in order to identify structural parameters (Carare and Tchaidze (2005)).

that includes expected future inflation as well as lagged values of the nominal interest rate.

## 3.1 Monte Carlo Experiment

The model we use for our Monte Carlo experiment is a hybrid New Keynesian model that features backward and forward-looking expectations. The framework is a small scale structural model and is very similar to the models currently used by policy-makers for forecasting and policy evaluation. Our model is similar to Amato and Laubach (2003) and Galí, López-Salido and Vallés (2004). We assume that the economy is populated by households, two types of firms; firms that produce differentiated intermediate goods and a perfectly competitive firm that produces a final good and a central bank that conducts monetary policy using a Taylor rule. The model is comprised of optimizing as well as rule of thumb consumers and producers.

The log-linearized conditions for the hybrid IS and Phillips curve are given by equations 3.2 and 3.3 respectively. Equation 3.4 describes the monetary policy rule of the central bank. We assume that the central bank uses a partial adjustment rule, that is one with interest rate smoothing. Potential output in the model is assumed for simplicity to follow an AR(1) process and equation 3.6 describes the various shocks of the model. The model is described by the following equations:

$$\begin{aligned}
y_t &= (1-\delta)y_{t-1} + \delta E_t y_{t+1} - \frac{1}{\tilde{\sigma}}\left(i_t - E_t\pi_{t+1}\right) + \epsilon_{y,t} && (3.2)\\
\pi_t &= \phi^f E_t\pi_{t+1} + \phi^b\pi_{t-1} + \tilde{\kappa}\left(y_t - \bar{y}_t\right) + \epsilon_{\pi,t} && (3.3)\\
i_t &= \rho_1 i_{t-1} + \rho_2 i_{t-2} + [1-\rho_1-\rho_2][con + \gamma(y_t - \bar{y}_t) + \beta\pi_{t+1}] + \xi_{i,t} && (3.4)\\
\bar{y}_t &= \eta\bar{y}_{t-1} + \mu_t && (3.5)\\
\epsilon_{j,t} &= \theta_j\epsilon_{j,t-1} + \eta_{j,t} \qquad for\ j \in \{y,\pi,i\} && (3.6)
\end{aligned}$$

where $E_t$ is the expectations operator, conditional on the information set at date $t$, and $y_t, i_t$ and $\pi_t$ are respectively output, the nominal interest rate and inflation, each expressed as deviations form their steady-state values. The *con* term is equal to $\bar{r} - (\beta - 1)\bar{\pi}$ where $\bar{r}$ and $\bar{pi}$ are respectively the equilibrium real interest rate and an inflation target. Potential output is denoted by $\bar{y}_t$ and we assume for simplicity that it follows and AR(1) process. The errors in the model are all serially correlated with the variance of $\eta_{j,t}$ given by $\sigma^2_{j,t}$.

The coefficients $\delta, \tilde{\sigma}, \phi^f, \phi^b$ and $\tilde{\kappa}$ can be written as functions of five structural parameters of the underlying optimization problems that generate the reduced form model given above. These structural parameters are the proportion of households that are optimizers, the intertemporal elasticity of substitution in consumption, the discount factor, the proportion of firms that cannot change their prices and the proportion of firms that reset their prices to last period's prices.[14]

We calibrate the model to a quarterly frequency using the parameter values of Amato and Laubach (2003). For the baseline case, the discount factor is set to 0.96, implying a steady-state real rate of return of 4%. The intertemporal elasticity of substitution is set to 1, implying logarithmic utility for consumption. We set the proportion of rule-of-thumb consumers to 0.5. Regarding price-setters, we assume that the fraction of firms that do not reset prices, is 0.5, implying that on average, prices are assumed to be sticky for two quarters. Finally, we set the proportion of firms that are rule of thumb to 0.5. These values imply that $\phi^f = 0.49$ and $\phi^b = 0.51$. We set $\beta = 1.5$ and $\gamma = 0.5$ as the baseline values for the interest rate rule as in the Taylor rule. The coefficients for the interest smoothing $\rho_1$ and $\rho_2$ are set respectively to 0.6 and 0.15 following values typically found in the literature. In the baseline case, we assume that the shock processes are *i.i.d.*. We relax this assumption when we perform some sensitivity tests. The variance of the shocks are calibrated using the same values as Carillo (2008) where $\{\sigma_{y,t}, \sigma_{\pi,t}, \sigma_{i,t}\} = \{0.23, 0.14, 0.24\}$.

## 3.2 Simulation and estimation

The Monte Carlo experiment is carried out by simulating the above model by generating 10,000 samples of 500 observations of output, inflation and the nominal interest rate. We then use the generated data to estimate a single-

[14]See Amato and Laubach (2003) for details of the derivation.

equation policy rule of the form:

$$i_t = \rho_1 i_{t-1} + \rho_2 i_{t-2} + [1 - \rho_1 - \rho_2] [con + \gamma(y_t - \bar{y}_t) + \beta E_t \pi_{t+1}] + \xi_t \quad (3.7)$$

We follow the same methodology as Clarida, Galí and Gertler (2000) and use the Generalized Methods of Moments (GMM) to estimate equation 3.7 instead of OLS. OLS would produce biased estimates of $\beta$ and $\gamma$ even if we replace the expectations of inflation with its contemporaneous value for at least two reasons. Actual inflation is an imperfect measure of expected inflation and thus substituting expected inflation by its actual value would likely make the error term correlated with the future inflation rate (errors in variables). Moreover, OLS is not appropriate here because of an endogeneity bias since future inflation is influenced by the interest rate.

We perform the estimation using 2-step GMM on each of the 10,000 samples of 500 observations and use four lags of output, inflation and interest rate and a constant as the set of instruments.[15] We then use a normal kernel density function to estimate the probability density function of $\beta$ and $\gamma$. In each case, we check whether the mean of the reported estimated $\beta$ and $\gamma$ is close to the "true" parameters of 1.5 and 0.5 respectively.

In addition to using the baseline parameter values for our hybrid model, we also perform several sensitivity tests. We run the same Monte Carlo simulations in a completely forward-looking model ($\delta = 1$, $\phi^f = \beta$, $\phi^b = 0$) and a completely backward-looking model ($\delta = 0$, $\phi^f = 0$, $\phi^b = 1$). We also allow for serial correlation in the disturbances of the model and redo the simulations for the three different versions of the model. We thus report six sets of simulation results: our hybrid, completely backward and forward-looking models with no serially correlated shocks and the same models with serially correlated shocks. When serially correlated shocks are assumed, we set a mild degree of serial correlation and set the coefficient to 0.2 for all the disturbances.

[15] We estimate the policy rule with different set of instruments. The results with these different set of instruments are very similar to our baseline case. We also estimated the policy rule using iterative GMM. In this case also, our results are not very different from our baseline case.

Our results are reported in both Table 1 and Figure 7. The table reports the mean of the 10,000 estimates that we perform each time. The size of the bias shows how far we are from the true value for $\beta$ and $\gamma$. The standard deviations are shown in parentheses.

The results suggest that a great deal of bias occurs when the data is generated by the purely forward looking version of the model. In this case, the value of $\gamma$ is negative while we obtain a mean estimate for $\beta$ close to zero. This is not surprising given the findings of Cochrane (2007). On the other hand, when the data is generated by the purely backwards looking model, the size of the bias is basically zero on both parameters. This result holds even when we allow for serial correlation. In the hybrid model, absent serial correlation, the size of the bias is close to zero for both parameters. However, with serial correlation, the size of the bias on the coefficient $\beta$ quickly increases and is no longer small. The size of the bias on $\alpha$, on the other hand, is less affected by the introduction of serial correlation.

The main lesson that we learn from our simulations is that as long as the shocks are not serially correlated, the coefficient $\beta$ can be estimated without bias when the data generating process takes the form of our hybrid model. This is an important result for our estimation strategy since this implies that identification is indeed possible as long as these conditions are met.

# 4 Estimation

In this section we estimate policy reaction functions of the form given by equation 3.7 on data from 14 OECD countries. We use annualized changes in the quarterly CPI as our measure of inflation. Since long time series for quarterly GDP are not available for very many countries, we use quarterly industrial production as our measure of economic activity. In order to generate an estimate of the output gap, we de-trend quarterly IP using an H-P filter.[16] The measure of interest rates varies across countries. We generally use short

[16] Alternate de-trending procedures, such as using a cubic time trend or a bandpass filter, do not change the main results.

term T-bill or money market rates as the measure of nominal interest rates. For countries with multiple available short term rate series extending back far enough, we have replicated the results with alternate measures of the nominal interest rate.

We employ two approaches to determine where the recent period, during which monetary policy parameters are supposed to be in the determinacy region, starts. Our first approach to determining where the recent period begins is to employ turning points in national inflation rates identified by previous researchers.[17] Our second approach is to adopt a break date commonly used in studies of U.S. monetary policy, thought to correspond to change in the conduct of U.S. monetary policy (1979q4). We then simply apply this break to all countries in the sample. Table 2 reports the dates.

The policy rule with forward-looking inflation expectations is usually estimated by replacing the unobserved expectations term, $E_t\pi_{t+1}$ by $\pi_{t+1} + \epsilon_{t+1}$ where $\epsilon_{t+1}$ is an expectational error that is assumed to be orthogonal to the information set at time $t$. We can therefore rewrite our policy rule as:

$$i_t = \rho_1 i_{t-1} + \rho_2 i_{t-2} + [1 - \rho_1 - \rho_2]\,[con + \gamma(y_t - \bar{y}_t) + \beta\pi_{t+1}] + e_t \tag{4.8}$$

where $e_t = \xi_t - (1 - \rho_1 - \rho_2)\beta\epsilon_{t+1}$. The moment condition is thus given by:

$$E\left[\left(i_t - \rho_1 i_{t-1} - \rho_2 i_{t-2} - [1 - \rho_1 - \rho_2]\,[con + \gamma(y_t - \bar{y}_t) + \beta\pi_{t+1}]\right) Z_t\right] = 0 \tag{4.9}$$

where $Z_t$ is a vector of predetermined variables or instruments. This orthogonality condition forms the basis for estimating the policy rule by GMM.

We use two lags instead of one lag on the interest rate to remove any serial correlation from our estimation. We formally test for the presence of serial correlation and find that our residuals are indeed not serially correlated. Given that our Monte Carlo simulations show that the parameters from a single equation are well identified in a hybrid model with no serial correlation, this increases our confidence that we have indeed identified the parameters correctly.

[17]The dates we use are taken from Corvoisier and Mojon (2005).

We follow Clarida, Galí and Gertler (2000) and use their set of instruments. Our instrument list includes a constant, four lags of the interest rates, four lags of the output gap and four lags of the inflation rate. The instruments are chosen based on the assumption that they are correlated with future inflation and orthogonal to the error term. To obtain the variance-covariance matrix of the moment conditions, unlike Clarida, Galí and Gertler (1998) who selects a 12-lag Newey-West estimate, we use a 4-lag Newey-West.[18] We estimate the policy rule using two alternative versions of GMM: the two-step GMM of Hansen and Singleton(1992) and the iterative GMM approach of Hansen, Heaton, and Yaron (1996).

As the number of instruments and hence orthogonality conditions exceed the number of parameter estimates, the model will be overidentified. We test the overidentifying restrictions using the $J$ test of Hansen (1982). Under this test, the null hypothesis that the overidentifying restrictions on the moment condition are valid for the vector of estimated parameters, has an asymptotic chi-square distribution with $n$ degrees of freedom where $n$ is the difference between the number of moment conditions and the number of parameters. A rejection of these restrictions would indicate that some of these moment conditions are not valid for the estimated parameters, thus implying that the model is mis-specified.

In addition to testing the whole set of moment conditions, we also test subsets of the orthogonality conditions using the Eichenbaum, Hansen and Singleton (1988) test, usually known as the $C$-test. Under this test, the null hypothesis is that the overidentifying restrictions on the moment condition in the unrestricted model are valid whereas the alternative hypothesis postulates that the null is not valid and the model estimated with the excluded instruments is valid (the restricted model). The difference between the $J$-statistic of the two models has an asymptotic chi-squared distribution with degrees of

[18]We estimated the model also by allowing for an automatic lag selection of the Newey-West. In most cases, the automatic selection returned a lag-length of around 4. For this reason, we choose to fix these lags at 4 in all of our regressions. The HAC regression allows us to obtain a convergent estimator for the variance but it has no effect on the values of the estimated coefficients.

freedom equal to the number of excluded instruments. The EHS test has been shown to have greater local power than the Hansen *J*-test.

We perform the EHS test by excluding one instrument at a time and then excluding two and three instruments at a time. Our block of moment conditions passes both the *J*-test and the EHS test. The two tests of overidentifying restrictions do not reject the null hypothesis that all moment conditions are correct for critical values at the 5% level. The third and fourth lagged value of the interest rate can, however, be excluded at the 10% level but not at the 5% level. We thus proceed by estimating our model with the block moment conditions comprising of four lags of the interest rate, output gap, inflation and a constant.[19]

The *J*-test and the EHS test provide some indication on the exogeneity of the instruments. However, the recent literature has also emphasized the importance of testing for the relevance of the instruments in order to detect whether they are strong or weak. Weak instruments lead to GMM (or IV) statistics that are non-normal, thus making the point estimates, hypothesis tests, and confidence intervals unreliable. According to Stock, Wright and Yogo (2002), in GMM, weak instruments in general is related to the weak identification of some or all of the unknown parameters. Although, we do not test directly how strong or weak our set of instruments are, we provide some sensitivity tests by estimating our model using Limited-Information Maximum Likelihood (LIML) and the continuous-updating estimator (CUE) of Hansen, Heaton and Yaron (1996). These two methods have been shown to be partially robust to the presence of weak instruments.

## 4.1 Results

Table 3 presents the results, using the Corvoisier and Mojon dates to determine the samples. The first thing to observe about the results is that the

[19] We estimated the model with three lags of the same set of instruments and found little difference in our results.

Taylor rule specification is a reasonable description of monetary policy for the countries in the sample. To measure the fit of the Taylor rule, we compare the actual and fitted values. In general, the policy rule we estimate fit the data very well.[20] This is not a surprising since policy rules that incorporate a large degree of interest rate inertia tend to fit the data very well.

Furthermore, the parameters of the various Taylor rules suggest common features to monetary policy in the countries in our sample. The coefficients on the first lag of inflation are close to one, a result that is often interpreted as evidence of interest rate smoothing, for all of the countries in the sample. Except for Finland, all countries have positive $\gamma$ coefficients, and these are statistically significant in only 5 out of the 14 countries when we employ the 2-step and iterative GMM. Finally, all countries have positive $\beta$ coefficient, and these coefficients are statistically significantly difference from zero in 7 of the 14 countries when we employ the 2-step GMM and in 9 out of 14 countries when we use iterative GMM.

Hansen's $J$-statistic suggests that we cannot reject our model specification for any country. However, since the sample size we use in most case is small, this test is known to have low power in small samples. We have estimated our model with different block of moment conditions, especially with fewer instruments. We do not find any qualitatives difference in our results.

Given Equation 3.7, the Taylor principle is generally interpreted as implying that, to ensure that monetary policy does not induce indeterminacy, the coefficient on expected inflation in the monetary policy rule should be greater than 1. In some models, the fact that output and inflation are jointly determined means that the central bank's response to output also influences determinacy. However, when the monetary policy reaction function exhibits interest rate smoothing, as in our Equation 3.7, the simple Taylor principle is both a necessary and sufficient restriction to ensure determinacy.[21]

[20] We do not present information concerning the ability of the policy rule to track interest rate movements, but the results are available from the authors.

[21] Evans and McGough (2005) demonstrate this result in a hybrid model New Keynesian model similar to ours. See also Clarida, Galí and Gertler (2000), Table 4.

The theory that the return to low inflation in the 1980s and 1990s was driven by a renewed adherence to the Taylor principle by monetary policy makers therefore implies that the coefficient on expected inflation to exceed 1 in our sample. The point estimate for this coefficient, which is reported in column 5, exceeds 1 in 7 of the 14 countries in our sample when we employ a 2-step GMM estimation (9 out of 14 when we employ iterative GMM). However, in only 3 of these cases is the coefficient statistically significantly larger than 1 when we estimate our model using 2-step GMM (France, Italy and the US) and in only 4 cases if we use iterative GMM (France, Italy, Ireland and the U.S).

One of the 3-4 countries for which $\beta$ exceeds 1 with statistical significance is the U.S. This is the well known result of Clarida, Galí and Gertler (2000). Of the other 13 countries[22] that experienced the low inflation of the 1980s and 1990s, there is statistical evidence for adherence to the Taylor principle in the post 1980 period only for 2-3 countries: France, Italy and (depending on the estimation method used) Ireland.

Table 5 presents our results when we employ 1979q4 as a common break date. This date is thought to represent a change in the conduct of monetary policy in the U.S and coincides with the arrival of Paul Volcker as chairman of the Fed and the start of strong dis-inflationary policies in the U.S. Our results using this common break date provide even less support for the New Keynesian story that the fall in inflation was due to the adherence to the Taylor principle. Only 5 countries have a value of $\beta$ that exceeds one and only the U.S has a coefficient of $\beta$ that is statistically larger than one. Corvoisier and Mojon (2005) find that the turning point in inflation in most countries occurred well after 1979. This may explain why we find even less support for the New Keynesian story when we use this common break date since the break in monetary policy happened much later.

Our results provide little support for the hypothesis that the fall in inflation in the 1980s and 1990s was due to a return to determinacy and an adherence

[22] Or perhaps 12, if Sweden's episode of high inflation in the early 1990s disqualifies it.

to the Taylor principle that central banks failed to follow prior to this period. Excluding the U.S., we find that the point estimate for $\beta$ exceeds one in about half of the countries in the sample. However, the coefficient is statistically larger than one only in 2-3 countries. Overall, the New Keynesian explanation of the fall of inflation fits the U.S data well but fails to perform well when confronted to international evidence.[23]

## 4.2 Robustness

It is well-known that efficient GMM estimation has the advantage of consistency in the presence of unknown heteroskedasticity, but the finite sample performance is poor. Recent research suggest that although the continuous-updating estimator (CUE) of Hansen, Heaton and Yaron (1996) and Limited Information Maximum Likelihood (LIML) offer no difference in efficiency asymptotically, it can perform better than the 2-step and iterative GMM in the presence of small sample and weak instruments. We therefore estimate our model using these two methods and verify that our results are robust to this change in estimation method. Since we are specifying a linear model, CUE-GMM is asymptotically equivalent to LIML under conditions of homoskedasticity. We find that in most cases, our estimation results from the CUE-GMM is identical to LIML.[24]

Our results are reported in Table 6. The main conclusion here is that our results are not substantively changed. The coefficient on inflation is statistically bigger than one, a condition needed for determinacy, in only a handful of countries. The three GMM procedures and the LIML however produce different standard errors. In general smaller standard errors are obtained with the CUE-GMM and the LIML compared to the 2-step and iterative GMM.

[23]We have also estimated our model with a common break date, 1979Q4. This date corresponds to the start of Paul Volcker at the Fed and is regarded as the beginning of a new monetary policy regime. Our results are similar to the baseline case. They are available on request.

[24]Although, we find that for a vast majority of countries the results from the CUE-GMM and LIML are similar, the standard errors under LIML are larger in most cases.

Based on these results, we again conclude that there is little support for the hypothesis that a renewed adherence to the Taylor principle explains the fall of inflation in recent decades.

Since our measures of dynamic correlation reveal that U.S inflation is highly correlated with the inflation of the other countries at low frequency, we use the current and lagged values of U.S inflation as instruments and restimate our model for all countries except the U.S. Using U.S inflation as instrument imply that the latter is assumed to be correlated with future inflation in these countries as well as well as satisfying the condition of exogeneity. In addition to current and lagged values of U.S inflation, we also use four lags of the country's output-gap and the country's interest rate as instruments. Our results are very similar to the baseline case. We have evidence to reject the New Keynesian story that the fall in inflation was caused by good monetary policy.

# 5 Conclusion

Our paper investigated the hypothesis that the fall of inflation was driven by renewed adherence to the Taylor principle by examining the experience of a number of OECD countries which have experienced similar inflation outcomes in recent decades. The test we proposed was a relatively weak test because: i) we could only test the implications of the theory for the fall of inflation, and not the rise, ii) we cannot rule out the possibility that indeterminate monetary policy might yield parameter estimates that look like the Taylor principle is satisfied, and iii) with positive trend inflation, satisfying the Taylor principle requires a stronger response to expected inflation than our test imposes. Despite the relative weakness of the test we apply, we find that the theory is not capable of explaining the cross country data. The results, therefore, constitute fairly serious support for the view that the behavior of inflation in OECD countries in recent decades is not largely driven by the impact of monetary policy on expected inflation.

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Table 1: Mean estimates and std deviation

| | $\gamma$ | *bias* | $\beta$ | *bias* |
|---|---|---|---|---|
| *i.i.d. shocks* | | | | |
| Baseline | 0.5 | 0.00 | 1.52 | 0.02 |
| | (0.20) | | (0.40) | |
| Backward-Looking | 0.50 | 0.00 | 1.51 | 0.01 |
| | (0.08) | | (0.07) | |
| Forward-Looking | -0.50 | -1.00 | 0.05 | -1.45 |
| | (0.05) | | (0.05) | |
| *Serial correlation* | | | | |
| Baseline | 0.48 | -0.02 | 1.25 | -0.25 |
| | (0.21) | | (0.37) | |
| Backward-Looking | 0.50 | 0.00 | 1.51 | 0.01 |
| | (0.08) | | (0.08) | |
| Forward-Looking | -0.41 | -0.91 | 0.21 | -1.29 |
| | (0.21) | | (0.37) | |

Table 2. Break Dates

| Country | Common | Corvoisier and Mojon |
|---|---|---|
| Australia | 1979:4 | -* |
| Canada | 1979:4 | 1982:3 |
| Denmark | 1979:4 | 1985:1 |
| Finland | 1979:4 | 1985:1 |
| France | 1979:4 | 1985:2 |
| Germany | 1979:4 | 1982:3 |
| Ireland | 1979:4 | 1984:2 |
| Italy | 1979:4 | 1985:4 |
| Japan | 1979:4 | 1981:2 |
| Netherlands | 1979:4 | 1982:2 |
| Spain | 1979:4 | 1982:2 |
| Sweden | 1979:4 | -* |
| Switzerland | 1979:4 | -* |
| U.K. | 1979:4 | 1981:4 |
| U.S. | 1979:4 | 1982:2 |

* No break date available between 1971 and 1989. We use 1982:1 as the start of the sample in these cases.

Table 3. Two Stage GMM, Corvoisier and Mojon sample

| Country | *con* | $\rho_1$ | $\rho_2$ | $\gamma$ | $\beta$ | *J*-stat |
|---|---|---|---|---|---|---|
| Australia | $3.82^a$ | $1.05^a$ | $-0.16^c$ | $1.05^c$ | 0.83 | 6.20 |
| | (0.89) | (0.10) | (0.09) | (0.47) | (0.20) | p=0.63 |
| Canada | $3.41^b$ | $0.92^a$ | -0.17 | $0.35^c$ | 1.42 | 7.16 |
| | (1.02) | (0.11) | (0.12) | (0.20) | (0.36) | p=0.52 |
| Denmark | 3.46 | $1.23^a$ | $-0.24^b$ | 0.87 | 1.73 | 3.72 |
| | (5.15) | (0.12) | (0.11) | (1.25) | (2.20) | p=0.88 |
| Finland | 4.09 | $1.26^a$ | $-0.25^a$ | -0.21 | 0.56 | 6.36 |
| | (3.76) | (0.05) | (0.05) | (0.71) | (1.80) | p=0.61 |
| France | 0.83 | $0.80^a$ | 0.07 | 0.74 | $2.82^f$ | 7.03 |
| | (1.94) | (0.20) | (0.13) | (0.52) | (0.91) | p=0.53 |
| Germany | $2.59^b$ | $1.32^a$ | $-0.38^a$ | $0.83^b$ | 0.86 | 5.29 |
| | (1.19) | (0.08) | (0.07) | (0.39) | (0.34) | p=0.73 |
| Ireland | $2.84^b$ | $0.94^a$ | -0.15 | 0.03 | 1.98 | 10.07 |
| | (1.17) | (0.09) | (0.09) | (0.14) | (0.42) | p=0.26 |
| Italy | -3.88 | $0.74^a$ | $0.19^a$ | 2.16 | $2.34^e$ | 5.13 |
| | (4.29) | (0.08) | (0.07) | (1.44) | (0.63) | p=0.74 |
| Japan | -6.06 | $1.37^a$ | $-0.37^a$ | 2.18 | 3.17 | 6.27 |
| | (11.41) | (0.10) | (0.10) | (2.96) | (3.98) | p=0.46 |
| Netherlands | 2.30 | $1.35^a$ | $-0.38^a$ | 1.32 | 0.85 | 6.52 |
| | (2.73) | (0.07) | (0.06) | (1.11) | (1.03) | p=0.64 |
| Spain | $5.20^b$ | $0.93^a$ | -0.11 | 0.21 | 0.99 | 4.76 |
| | (1.97) | (0.11) | (0.15) | (0.42) | (0.34) | p=0.78 |
| Sweden | 5.75 | $1.00^a$ | -0.03 | 2.86 | 0.77 | 6.71 |
| | (4.28) | (0.10) | (0.07) | (5.69) | (0.84) | p=0.57 |
| Switzerland | 1.67 | $1.08^a$ | -0.14 | $0.75^b$ | 0.66 | 7.77 |
| | (1.11) | (0.10) | (0.10) | (0.35) | (0.40) | p=0.46 |
| U.K. | $3.96^a$ | $1.08^a$ | $-0.16^a$ | 1.04 | 0.83 | 7.89 |
| | (1.42) | (0.07) | (0.06) | (0.87) | (0.27) | p=0.44 |
| U.S. | $3.88^a$ | $1.00^a$ | -0.13 | $0.74^b$ | $1.74^e$ | 7.23 |
| | (1.51) | (0.06) | (0.05) | (0.34) | (0.37) | p=0.51 |

a, b, c: statistically significantly different from zero at the 1%, 5% and 10% level, respectively

d, e, f: statistically significantly greater than one at the 1%, 5% and 10% level, respectively

Table 4. Iterative GMM, Corvoisier and Mojon sample

| Country | *con* | $\rho_1$ | $\rho_2$ | $\gamma$ | $\beta$ | *J*-stat |
|---|---|---|---|---|---|---|
| Australia | $3.85^a$ | $1.08^a$ | $-0.19^c$ | $1.03^b$ | 0.81 | 5.92 |
| | (0.87) | (0.10) | (0.10) | (0.45) | (0.20) | p=0.66 |
| Canada | $3.33^a$ | $0.89^a$ | -0.15 | $0.34^b$ | 1.46 | 5.34 |
| | (1.17) | (0.12) | (0.12) | (0.17) | (0.39) | p=0.72 |
| Denmark | 4.97 | $1.30^a$ | $-0.31^a$ | 0.57 | 1.10 | 4.31 |
| | (3.15) | (0.10) | (0.10) | (0.80) | (1.25) | p=0.78 |
| Finland | $3.87^a$ | $1.15^a$ | $-0.12^b$ | -0.11 | 0.50 | 6.28 |
| | (1.25) | (0.06) | (0.06) | (0.26) | (0.72) | p=0.62 |
| France | 0.28 | $0.50^a$ | $0.32^a$ | $0.77^b$ | $3.11^d$ | 6.35 |
| | (1.34) | (0.17) | (0.11) | (0.38) | (0.63) | p=0.61 |
| Germany | 2.04 | $1.21^a$ | $-0.26^a$ | $1.18^c$ | 0.80 | 4.77 |
| | (1.66) | (0.09) | (0.09) | (0.64) | (0.43) | p=0.78 |
| Ireland | 1.33 | $1.34^a$ | $-0.56^a$ | -0.00 | $2.56^d$ | 5.83 |
| | (1.51) | (0.11) | (0.11) | (0.16) | (0.52) | p=0.66 |
| Italy | -3.88 | $0.74^a$ | $0.19^b$ | 2.16 | $2.34^e$ | 5.14 |
| | (4.30) | (0.08) | (0.07) | (1.44) | (0.63) | p=0.74 |
| Japan | -7.09 | $1.36^a$ | $-0.37^a$ | 2.59 | 3.84 | 6.32 |
| | (15.62) | (0.10) | (0.10) | (4.21) | (5.74) | p=0.51 |
| Netherlands | 2.00 | $1.38^a$ | $-0.41^a$ | 1.07 | 1.14 | 6.12 |
| | (2.74) | (0.07) | (0.07) | (1.03) | (1.09) | p=0.64 |
| Spain | $4.90^a$ | $1.13^a$ | $-0.36^b$ | 0.07 | 1.09 | 3.99 |
| | (1.46) | (0.14) | (0.15) | (0.29) | (0.24) | p=0.65 |
| Sweden | 5.75 | $1.00^a$ | -0.03 | 2.87 | 0.77 | 6.71 |
| | (4.27) | (0.10) | (0.08) | (5.87) | (0.84) | p=0.57 |
| Switzerland | 1.49 | $1.07^a$ | -0.12 | $0.90^c$ | 0.60 | 7.16 |
| | (1.49) | (0.10) | (0.10) | (0.47) | (0.48) | p=0.52 |
| U.K. | $4.32^a$ | $1.15^b$ | $-0.23^a$ | 0.87 | 0.74 | 7.46 |
| | (1.31) | (0.08) | (0.07) | (0.85) | (0.29) | p=0.49 |
| U.S. | $3.47^a$ | $1.14^a$ | $-0.24^b$ | 0.39 | $2.14^d$ | 4.58 |
| | (1.71) | (0.12) | (0.10) | (0.36) | (0.43) | p=0.73 |

a, b, c: statistically significantly different from zero at the 1%, 5% and 10% level, respectively

d, e, f: statistically significantly greater than one at the 1%, 5% and 10% level, respectively

Table 5. Two Stage GMM, 1979:4 as break date

| Country | *con* | $\rho_1$ | $\rho_2$ | $\gamma$ | $\beta$ | *J*-stat |
|---|---|---|---|---|---|---|
| Australia | $3.82^a$ | $1.05^a$ | $-0.16^c$ | $1.05^c$ | 0.83 | 6.20 |
| | (0.89) | (0.10) | (0.09) | (0.47) | (0.20) | p=0.63 |
| Canada | $4.14^a$ | $0.86^a$ | $-0.12^c$ | $0.35^b$ | 1.17 | 7.91 |
| | (0.69) | (0.09) | (0.07) | (0.15) | (0.14) | p=0.44 |
| Denmark | $4.16^a$ | $1.24^a$ | $-0.29^a$ | 0.62 | 0.56 | 5.55 |
| | (1.15) | (0.09) | (0.09) | (0.53) | (1.25) | p=0.70 |
| Finland | 3.93 | $1.23^a$ | $-0.24^a$ | 0.44 | 0.34 | 7.13 |
| | (3.93) | (0.05) | (0.05) | (1.06) | (0.73) | p=0.52 |
| France | $4.98^a$ | $1.19^a$ | $-0.24^a$ | 0.44 | 0.34 | 7.13 |
| | (1.70) | (0.08) | (0.06) | (0.69) | (0.36) | p=0.38 |
| Germany | $2.59^b$ | $1.16^a$ | $-0.23^b$ | $1.02^b$ | 0.51 | 5.41 |
| | (1.29) | (0.11) | (0.11) | (0.46) | (0.43) | p=0.71 |
| Ireland | $6.81^a$ | $0.90^a$ | -0.06 | $0.42^c$ | 0.56 | 7.36 |
| | (0.90) | (0.09) | (0.09) | (0.24) | (0.10) | p=0.50 |
| Italy | 0.25 | $0.73^a$ | $0.19^a$ | 2.08 | 1.41 | 8.00 |
| | (4.10) | (0.07) | (0.07) | (1.28) | (0.39) | p=0.43 |
| Japan | -5.44 | $1.46^a$ | $-0.47^a$ | 3.91 | 3.81 | 7.09 |
| | (20.96) | (0.10) | (0.09) | (8.67) | (8.80) | p=0.52 |
| Netherlands | 2.48 | $1.42^a$ | $-0.44^a$ | 3.73 | 0.34 | 7.08 |
| | (3.50) | (0.15) | (0.14) | (2.99) | (1.36) | p=0.53 |
| Spain | $4.78^c$ | $0.73^a$ | 0.09 | 0.49 | 1.00 | 6.60 |
| | (2.82) | (0.10) | (0.10) | (0.52) | (0.43) | p=0.58 |
| Sweden | 5.75 | $1.00^a$ | -0.03 | 2.86 | 0.77 | 6.71 |
| | (4.28) | (0.10) | (0.07) | (5.69) | (0.84) | p=0.57 |
| Switzerland | 1.67 | $1.08^a$ | -0.14 | $0.75^b$ | 0.66 | 7.77 |
| | (1.11) | (0.10) | (0.10) | (0.35) | (0.40) | p=0.46 |
| U.K. | $3.81^a$ | $1.05^a$ | $-0.14^b$ | 1.26 | 0.82 | 9.00 |
| | (1.33) | (0.06) | (0.06) | (0.77) | (0.17) | p=0.34 |
| U.S. | 1.78 | $0.95^a$ | -0.09 | $0.70^b$ | $1.54^e$ | 7.33 |
| | (1.19) | (0.06) | (0.06) | (0.29) | (0.27) | p=0.50 |

a, b, c: statistically significantly different from zero at the 1%, 5% and 10% level, respectively

d, e, f: statistically significantly greater than one at the 1%, 5% and 10% level, respectively

Table 6. CUE-GMM, Corvoisier and Mojon sample

| Country | *con* | $\rho_1$ | $\rho_2$ | $\gamma$ | $\beta$ | *J*-stat |
|---|---|---|---|---|---|---|
| Australia | $3.95^a$ | $1.14^a$ | $-0.24^b$ | $1.01^b$ | 0.77 | 5.82 |
| | (0.86) | (0.11) | (0.10) | (0.44) | (0.20) | p=0.66 |
| Canada | $2.79^a$ | $0.84^a$ | -0.16 | $0.28^c$ | $1.64^f$ | 4.89 |
| | (1.00) | (0.13) | (0.12) | (0.14) | (0.33) | p=0.72 |
| Denmark | $5.29^c$ | $0.96^a$ | 0.01 | 0.79 | 1.22 | 3.79 |
| | (2.77) | (0.15) | (0.15) | (0.90) | (1.29) | p=0.88 |
| Finland | $3.34^a$ | $1.14^a$ | $-0.12^c$ | -0.21 | 0.67 | 6.16 |
| | (1.04) | (0.07) | (0.07) | (0.24) | (0.56) | p=0.63 |
| France | 0.65 | -1.27 | $1.60^b$ | $1.12^a$ | $2.93^d$ | 5.37 |
| | (1.16) | (0.88) | (0.62) | (0.35) | (0.47) | p=0.72 |
| Germany | 1.84 | $1.16^a$ | $-0.22^b$ | $1.18^c$ | 0.84 | 4.65 |
| | (1.60) | (0.09) | (0.09) | (0.60) | (0.40) | p=0.72 |
| Ireland | 1.13 | $1.45^a$ | $-0.66^a$ | -0.09 | $2.74^d$ | 5.57 |
| | (1.66) | (0.12) | (0.12) | (0.18) | (0.59) | p=0.66 |
| Italy | -1.29 | $0.70^a$ | $0.22^b$ | 1.86 | $1.93^e$ | 5.03 |
| | (2.90) | (0.07) | (0.07) | (1.20) | (0.44) | p=0.75 |
| Japan | -11.38 | $1.49^a$ | $-0.49^a$ | 5.48 | 7.78 | 6.10 |
| | (13.79) | (0.10) | (0.10) | (5.64) | (9.90) | p=0.46 |
| Netherlands | 2.39 | $1.38^a$ | $-0.41^a$ | -0.17 | 1.63 | 5.79 |
| | (2.91) | (0.07) | (0.07) | (1.04) | (1.24) | p=0.69 |
| Spain | $5.21^a$ | $1.12^a$ | $-0.39^a$ | -0.06 | 1.09 | 3.73 |
| | (1.36) | (0.14) | (0.15) | (0.23) | (0.22) | p=0.68 |
| Sweden | 4.48 | $1.04^a$ | -0.07 | 2.33 | 0.77 | 6.44 |
| | (3.58) | (0.10) | (0.08) | (3.62) | (0.64) | p=0.60 |
| Switzerland | 0.55 | $1.10^a$ | -0.13 | 1.33 | 0.56 | 6.93 |
| | (3.06) | (0.11) | (0.10) | (1.06) | (0.82) | p=0.54 |
| U.K. | $4.94^a$ | $1.27^a$ | $-0.33^a$ | 1.16 | 0.57 | 7.16 |
| | (1.76) | (0.10) | (0.09) | (1.18) | (0.44) | p=0.52 |
| U.S. | $3.40^b$ | $1.21^a$ | $-0.30^a$ | 0.28 | $2.26^d$ | 4.36 |
| | (1.70) | (0.15) | (0.12) | (0.38) | (0.48) | p=0.72 |

a, b, c: statistically significantly different from zero at the 1%, 5% and 10% level, respectively

d, e, f: statistically significantly greater than one at the 1%, 5% and 10% level, respectively

Figure 1: Inflation trends - pairwise with the U.S

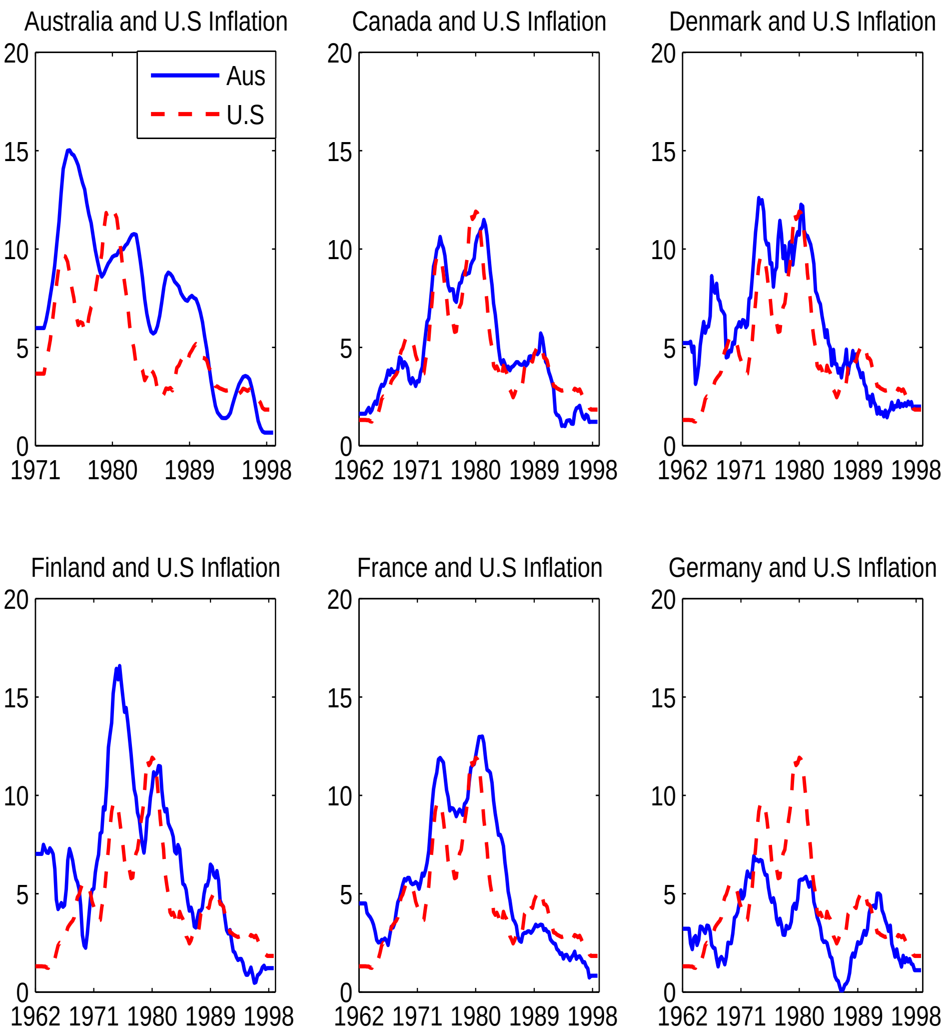

Figure 2: Inflation trends - pairwise with the U.S

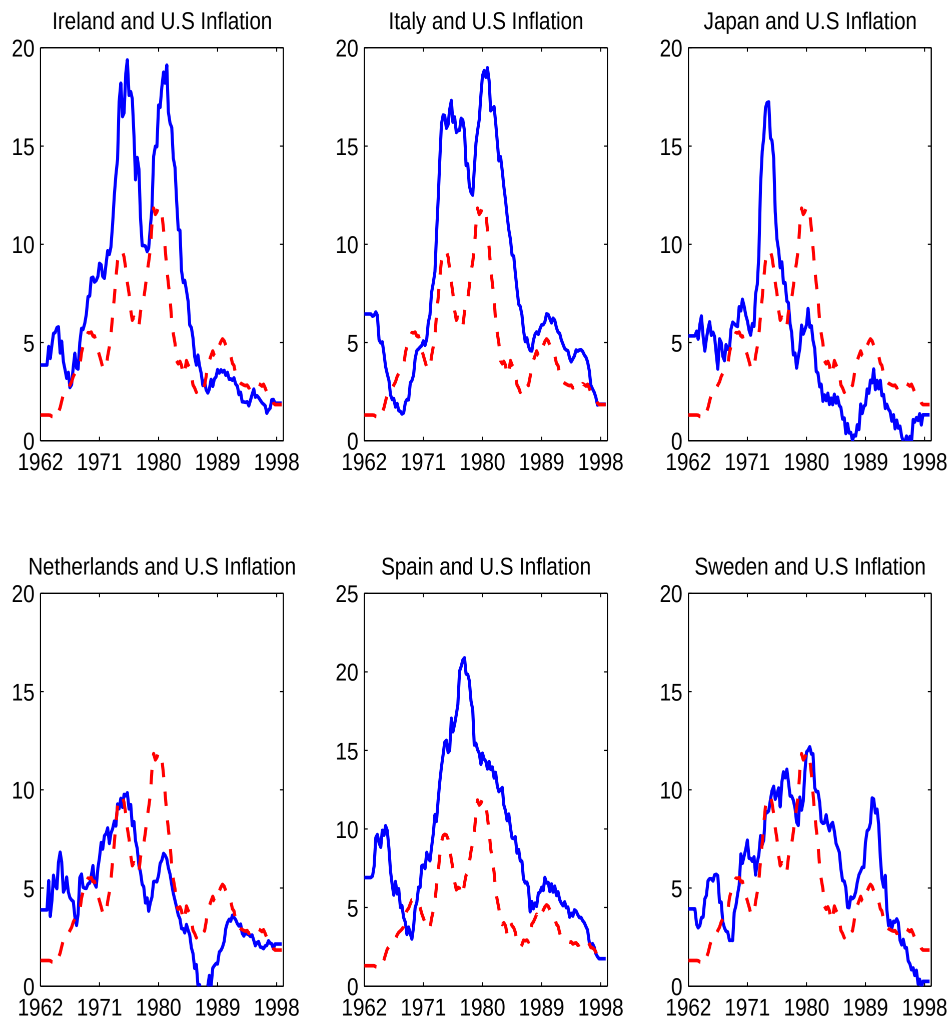

Figure 3: Inflation trends - pairwise with the U.S

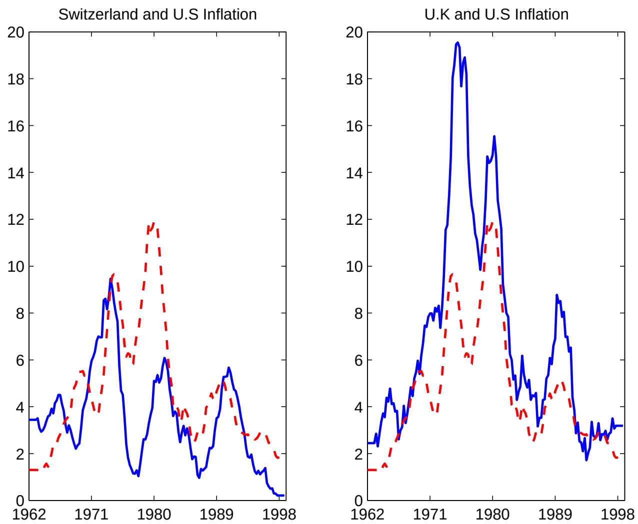

Figure 4: Dynamic correlation with the U.S

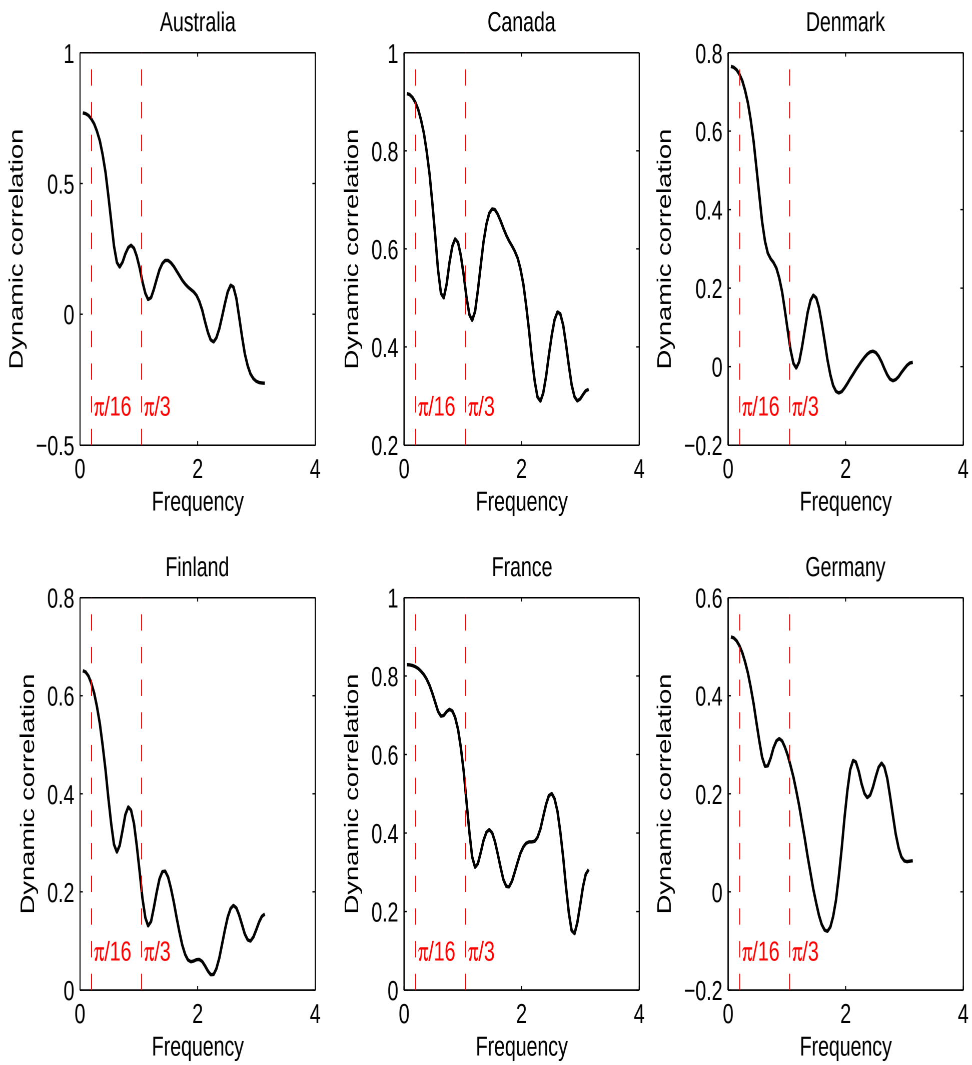

Figure 5: Dynamic correlation with the U.S

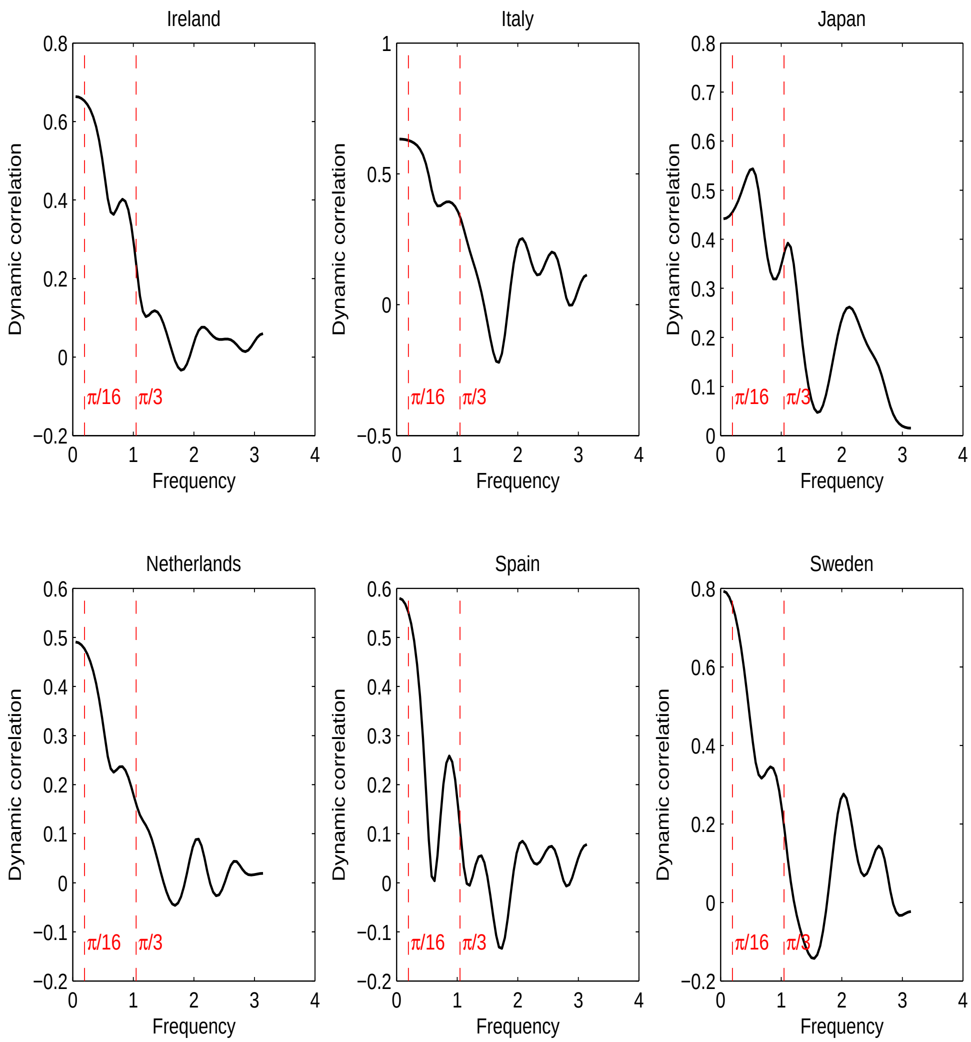

Figure 6: Dynamic correlation with the U.S

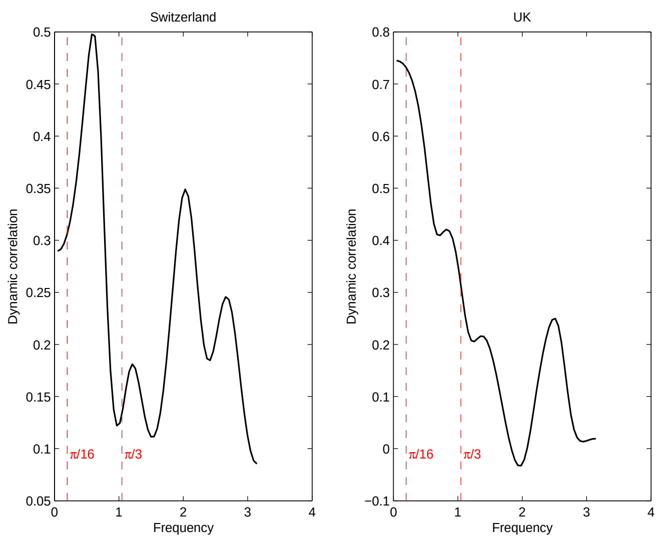

Figure 7: Monte Carlo Densities