

# Testing Commitment Models of Monetary Policy: Evidence from OECD Economies

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## **Abstract**

Inflation rates in many OECD countries was low in the 1960s, rose for a time before peaking in the 1970s or early 1980s, and then fell back to initial levels. This paper shows that a simple time inconsistency model of monetary policy does not explain OECD inflation outcomes, except in the U.S. The hypothesis that time inconsistency mattered only in earlier decades fits the data no better than the baseline model. We find some, albeit limited support for a model in which inflation spills over from the U.S. into other countries as a result of exchange rate targeting.

# 1 Introduction

A key feature of inflation in many industrialized economies in recent decades was the substantial run-up of inflation in the late 1960s and 1970s, followed by an equally substantial dis-inflation in the 1980s and 1990s. In the U.S., the period of high inflation is sometimes referred to as the Great Inflation, and has been described as “the greatest failure of American macroeconomic policy in the post war period” (Mayer 1999). A substantial body of recent research attempts to explain the rise and fall of inflation in the U.S., but has paid little attention to the international dimension of the issue to date. The similarity in the behavior of trend inflation,<sup>1</sup> however, suggests that a good explanation of long run inflation outcomes ought to apply across OECD countries.

In this paper, we ask whether time inconsistency models of monetary policy based on the framework of Kydland & Prescott (1977), and Barro & Gordon (1983), can explain inflation trends across OECD economies. Ireland (1999) finds that the Kydland-Prescott, Barro-Gordon (KPBG) model, extended to allow for a time varying NAIRU, is consistent with the U.S. data. As the model is general enough to encompass the institutional arrangements across OECD countries, it is natural to ask whether the success of the model in matching U.S. outcomes extends to an explanation of the common trend in international inflation. Although the basic KPBG framework is a well known and influential model in macroeconomics, there has been relatively little empirical testing of the framework. Furthermore, policy insiders have questioned the relevance of these models, arguing that the time inconsistency story is a poor representation of policymakers’ behavior.<sup>2</sup> Thus, an assessment of the empirical performance of the time inconsistency mechanism for inflation outcomes adds to our understanding of both the causes of historical inflation trends and the relevance of a

well established class of macroeconomic models.

Ireland (1999) observes that in the KPBG framework, as the NAIRU rises, central bankers increase their attempts to drive unemployment down, which leads to increasing inflation. Essentially, any long run trend in the NAIRU is reflected in long run inflation and unemployment trends. If the NAIRU is  $I(1)$ , then both inflation and unemployment inherit the non-stationarity of the NAIRU and, according to the model, must be cointegrated. This insight provides the basis for statistical tests of the model, which we apply to OECD data. Our first pass at the data is to apply the insight of Ireland to quarterly data from 13 OECD countries going back to 1964. Our results suggest that inflation and unemployment are not cointegrated in OECD countries, with the sole exception of the U.S..

An obvious problem with this first pass approach is that a number of key model parameters are unobserved and there is no strong reason to believe that these parameters have remained constant throughout the period of interest. Thus, the failure of the time inconsistency model to fit the data may be due to a failure of the assumption that the model parameters were unchanged over the estimation period. We investigate whether allowing for structural breaks significantly improves the empirical performance of the model, using the Gregory-Hansen test for cointegration in the presence of a possible structural break in the cointegrating relationship. Our results imply that allowing for time varying model parameters in this way does not overturn the conclusion that the time inconsistency framework does not fit the data for most OECD countries.

An important variant of the KPBG story is the widespread view that policy makers targeted unattainable unemployment rates in the 1960s and 1970s, but, perhaps due to advances in economic theory, became more cautious about trying to use monetary policy to offset high unemployment rates in more recent decades.<sup>3</sup> If correct,

this variant of the theory implies that inflation and unemployment should be cointegrated in the first part of the sample, when policy makers were still treating the Phillips curve as an exploitable relationship. In the latter portion of the sample, when policy makers learned not to target unemployment rates below the NAIRU, the time inconsistency model no longer describes the behavior of central bankers, so inflation and unemployment ought not to be related at all.

We test this variant of the basic KPBG model by re-estimating the model using only data from the first part of the sample. Surprisingly, given the prevalence of this hypothesis among macroeconomists, the time inconsistency account fits little better when estimated just on data from the 1960s and 1970s than it does on data over the whole sample. The main problem is that the theory can only deliver rising inflation in the presence of increases in the NAIRU, but outside of the U.S. there is little evidence of a rising NAIRU during the period of rising inflation. We conclude that this variant of the theory finds little more support than the baseline model.

A final variant on the basic framework allows that inflation may have spilled over from the U.S. to other countries, due to a dislike on the part of monetary authorities in smaller countries of large nominal exchange rate movements with respect to the U.S. In this case, an increase in the U.S. NAIRU drives up U.S. inflation, which forces foreign monetary authorities to allow domestic inflation to rise, so as to avoid an appreciation of the domestic currency. We extend the baseline model to incorporate exchange rate targeting, and show that U.S. unemployment rates enter as an additional cointegrating variable into the smaller country's inflation and unemployment relationship.

We test for cointegration between domestic inflation, and domestic and U.S. unemployment. The results do suggest that there is a cointegrating relationship, with a positive sign, between domestic inflation and U.S. unemployment in two thirds of

the OECD countries in our sample. Thus, the evidence is consistent with an account in which time inconsistency problems in U.S. monetary policy caused inflation in the U.S. to rise in the 1960s and 1970s, and that this inflation spilled over into other OECD countries via exchange rates. The empirical behavior of foreign monetary authorities with respect to their domestic unemployment rates, however, does not conform to the predictions of the KPBG model.

Our results suggest that if pre-commitment issues in the conduct of monetary policy are part of the explanation of the rise and fall of inflation, it is not likely via the standard “inflationary bias” mechanism. Time inconsistency, however, may affect monetary policy for reasons other than “inflationary bias”. Examples include the so called “stabilization bias”, where the inability to commit prevents the central banker from stabilizing expectations of inflation, thus affecting the transition path of inflation between steady states (Svensson 1997, Clarida, Gali & Gertler 1999, and Woodford 1999), and the hypothesis that central bank preferences are asymmetric, which implies that changes in the volatility of inflation and/or unemployment affect equilibrium inflation (Cukierman 1999, Ruge-Murcia 2002, and Doyle & Falk 2006). Whether time inconsistency has been an important driver of inflation trends via one of these other channels remains a question for future research.

Our paper is related to the literature investigating the causes of the ‘Great Inflation’ in the U.S.. This literature includes Clarida, Gali, & Gertler (2000), who argue that the rise of inflation in the late 1960s and early 1970s was due to mistakes made by monetary policy authorities. Explanations of this behavior include the possibility that the Fed conducted otherwise correct monetary policy using bad data (Orphanides 2002, 2003), that the Fed was learning about key parameters of the economy as it went along (Sargent 1999, Primiceri 2006), and that the Fed was responding to unfavorable fundamentals, perhaps filtered through the lens of time

inconsistency problems (Ireland 1999). Our paper clearly falls into this last category. The main innovation is the use of the common international experience as a way of disciplining our empirical work.

The paper proceeds as follows. In Section 2 we present a typical model of time-consistent monetary policy with a time varying NAIRU, as well an extension of the model incorporating international spillovers in inflation. In Section 3 we present the results of our econometric tests of the long run restrictions of the baseline model. We also present the results of tests incorporating both structural breaks and international spillovers into the baseline framework. In Section 4 we offer concluding remarks.

## 2 A Time Inconsistency Model of Monetary Policy

In this section we present a version of the simple KPBG time inconsistency model of monetary policy. The model begins with an expectations augmented short run Phillips curve:

$$u_t = u_t^n - \alpha(\pi_t - \pi_t^e), \quad (2.1)$$

where  $\pi_t$  is the rate of inflation in period  $t$ ,  $\pi_t^e$  represents household's expectations of period  $t$  inflation,  $u_t$  is the rate of unemployment, and  $u_t^n$  is the NAIRU.

The policy maker's preferences over inflation and unemployment take the form of a loss function:

$$L(\pi_t, u_t) = 1/2 [ b \cdot \pi_t^2 + (u_t - k \cdot u_t^n)^2 ], \quad (2.2)$$

where  $b$  represents the central bank's distaste for inflation and  $k \cdot u_t^n$  represents the central bank's target level of unemployment. We assume that  $k < 1$ , in which case the central bank targets an unemployment rate below the NAIRU.

The policy maker chooses a target rate of inflation,  $\pi_t^p$ , so as to minimize the expected loss. The policy maker takes expected inflation as given when choosing  $\pi_t^p$ ,

though, in fact, agents have rational expectations, implying that  $\pi_t^e = \pi_t^p$ . Actual inflation in each period equals the policy maker's inflation target, plus some control error  $\eta_t$ :

$$\pi_t = \pi_t^p + \eta_t, \quad (2.3)$$

where  $\eta_t$  is a serially uncorrelated random variable with mean zero and standard deviation  $\sigma_\eta$ .<sup>4</sup>

Taking the first order condition of the policy maker's problem and applying rational expectations gives

$$\pi_t^p = \pi_t^e = \alpha \left( \frac{1-k}{b} \right) E_{t-1} u_t^n. \quad (2.4)$$

In order to enable the model to address trends in inflation, we incorporate a time varying NAIRU into the model as follows:

$$u_t^n = u_{t-1}^n + \epsilon_t, \quad (2.5)$$

where  $\epsilon_t$  is a serially uncorrelated random variable with mean zero and standard deviation  $\sigma_\epsilon$ . Under this assumption, equilibrium inflation and unemployment are given by

$$u_t = u_{t-1}^n + \epsilon_t - \alpha \eta_t \quad (2.6)$$

and

$$\pi_t = \alpha \left( \frac{1-k}{b} \right) u_{t-1}^n + \eta_t. \quad (2.7)$$

Separately, 2.6 and 2.7 show how inflation and unemployment inherit a unit root from the underlying process for the NAIRU. Combined, these equations give

$$\pi_t - \alpha \left( \frac{1-k}{b} \right) u_t = -\alpha \left( \frac{1-k}{b} \right) \epsilon_t + \left( 1 + \alpha^2 \left( \frac{1-k}{b} \right) \right) \eta_t. \quad (2.8)$$

Thus, if the NAIRU follows a unit root, inflation and unemployment must be cointegrated. It is this implication of the model that we test empirically in Section 3.

## 2.1 Extension: Open Economies

In this section we extend the model to allow for the possibility that inflation in one country is transmitted to another country via the effect on exchange rates by assuming that central bankers target changes in nominal exchange rates in addition to domestic inflation and unemployment. The assumption that central banks desire some degree of exchange rate smoothing seems reasonable in light of recent research that documents how infrequently central banks let exchange rates float freely (Calvo and Reinhart 2002, Reinhart and Rogoff 2004).

The model is as before, except that the loss function becomes

$$L(\pi_t, u_t, \Delta e_t) = 1/2[b \cdot \pi_t^2 + (u_t - k \cdot u_t^n)^2 + \phi(\Delta e_t)^2], \quad (2.9)$$

where  $\Delta e_t$  is the change in the log of the nominal exchange rate.

We assume that a simple PPP theory describes exchange rates, so that  $e_t = p_t - p_t^f$ , where  $p_t$  is the log of the domestic price level and  $p_t^f$  is the log of the foreign price level. This implies that

$$\Delta e_t = \pi_t - \pi_t^f, \quad (2.10)$$

where, for simplicity, we assume that the foreign country doesn't target the exchange rate and, therefore, foreign inflation is determined by the baseline model.

The equilibrium outcome of this economy is described by:

$$\begin{aligned} \pi_t - \left[ \frac{\alpha(1-k)}{b+\phi} \right] u_t - \left[ \frac{\phi\alpha^f(1-k^f)}{(b+\phi)b^f} \right] u_t^f &= - \left[ \frac{\alpha(1-k)}{b+\phi} \right] \epsilon_t + \left[ \frac{(1+\alpha)\alpha(1-k)}{b+\phi} \right] \eta_t \\ &\quad - \left[ \frac{\phi\alpha^f(1-k^f)}{(b+\phi)b^f} \right] (\epsilon_t^f - \alpha^f \eta_t^f), \end{aligned} \quad (2.11)$$

where  $f$  superscripts denote the foreign economy.

When the policy maker targets exchange rate movements, it is forced to keep domestic inflation in line with foreign inflation, which depends on the NAIRU of the

foreign country. The result is that domestic inflation, domestic unemployment, and foreign unemployment are cointegrated. The restrictions on the model parameters imply that both  $\alpha(1-k)/(b+\phi)$  and  $\phi\alpha^f(1-k^f)/((b+\phi)b^f)$  are non-negative. Thus, the coefficients on domestic and foreign unemployment in the cointegrating vector must be non-positive.

### 3 Testing the Model

We test the model on quarterly data from a number of OECD countries. Our variables are the civilian unemployment rate (as reported in the OECD's Main Economic Indicators) and the rate of inflation, measured by the quarterly percentage change in the Consumer Price Index (again taken from the MEI).<sup>5</sup> We end up with a sample of 13 countries for which sufficient time series data are available at quarterly frequencies. Due to data availability issues, we use different samples for each country. The samples used are described in column 1 of Table 1.

Equations (2.6) and (2.7) imply that according to the model, both the unemployment rate and inflation rate ought to be unit root processes. Unit root test results are reported in columns 2 and 3 of Table 1. The test we applied was the Augmented Dickey-Fuller test, de-meaning the data first by GLS and then using Ng and Perron's (2001) Modified AIC to select the number of lagged differences to include in the ADF regression.<sup>6</sup> This test seems to have relatively good size and power properties for series like unemployment and inflation that tend to have a large negative moving-average root.

Columns 2 and 3 of Table 1 report the test results for matched sample periods of the unemployment rate and the CPI-based inflation rate for the 13 OECD countries that form our large sample. In no case was it possible to reject the null hypothesis that the series possesses a unit root at the 5% significance level. We can reject the

null at a 10% level for the inflation rate in the U.K and the unemployment rate in the U.S. Based on these results it seems reasonable to treat both the unemployment rate and the inflation rate as unit root processes across all countries and to proceed to test for cointegration.

We begin our estimation by verifying econometrically the observation that the time series pattern of inflation outcomes over the past four decades has been similar across the countries in our sample. We do this by applying Johansen's (1988) maximum likelihood approach to test for cointegration between a country's inflation rate and the U.S. inflation rate. As column 4 of Table 1 shows, the long run trends in inflation rates are very similar across countries.<sup>7</sup> The null hypothesis of no cointegration is rejected at standard significance levels for all countries in our sample except Austria and Japan.<sup>8</sup> These results suggest that trends in inflation in different countries are related. This is an important observation, as it suggests there may be a common explanation of inflation trends in OECD countries.

With these results in hand, we turn to our empirical implementation of the simple model of Section 2, applying Johansen's (1988) maximum likelihood approach to test for cointegration between unemployment and inflation rates. The results are reported in column 1 of Table 2. The first thing to note is that the null hypothesis that inflation and unemployment are not cointegrated is rejected at the 5% level for the U.S. In this sense our results mirror those of Ireland (1999). However, our results differ from his once we look at other countries in our sample. The null of no cointegration between unemployment and inflation cannot be rejected at the 5% level for any of the 13 OECD countries other than the U.S. It can be rejected at the 10% level only for Canada.

Given the close economic connections between OECD countries the errors in our country, estimations may be contemporaneously correlated. Typically, panel cointe-

gration tests take advantage of panel asymptotics (large  $T$  and large  $N$ ) to test the null that the  $N$  regression residual processes are all unit root processes, but do not allow for cross-sectional dependence across units. Westerlund (2005), however, has proposed a pair of residual-based panel cointegration tests that allow for the existence of common time effects and, possibly, more general forms of cross-sectional correlation structures.

We constructed Westerlund's VRG test-statistic for two cases: first, for the seven countries for which have data back to 1964:1,<sup>9</sup> and second, for all 13 countries over the largest possible common sample (1970:1-2002:2). Under the null hypothesis that none of the inflation-unemployment pairs are cointegrated, the (appropriately standardized) VRG statistic is asymptotically standard normal. Under the alternative hypothesis that some or all of the pairs are cointegrated, the test is a one-sided test in the "negative direction." That is, we reject  $H_0$  at the 5% level, for example, if the test statistic is less than -1.65. The calculated value of the VRG statistic is -0.41 when calculated for the seven countries with the longest data series, and 3.07 when applied to the sample with all of the countries. Consequently, the null of no cointegration is not rejected. We conclude that the conclusion that inflation and unemployment are not cointegrated in OECD countries is robust to allowing for correlation of error terms across countries.

The results presented in this section suggest that the simplest time inconsistency model of inflation does not provide an adequate explanation for the pattern of rising and then falling inflation rates observed in OECD economies over the past four decades. In subsequent sections we explore the possibility that extensions of the baseline model provide a better explanation of the data.

### 3.1 Time Varying Parameters

The previous test is a joint test of the model’s mechanism, plus the questionable assumption that the various parameters of the model have remained constant over the relevant time horizon.<sup>10</sup> Not accounting for regime shifts in the cointegrating relationship may explain why we did not find more evidence of cointegration between inflation and unemployment.

In this section, we address this problem by allowing for structural breaks in the cointegrating relationship between inflation and unemployment. We apply Gregory & Hansen’s (1996) residual-based test for cointegration, which allows for the possibility of a regime shift under the alternative of cointegration, although the timing of the regime shift is not known. The null hypothesis is that the series are not cointegrated. The particular version of the Gregory-Hansen test we apply 1) uses their ADF\* statistic, 2) applied to the residuals from their “regime shift” or “C/S” cointegrating regression, which allows for a change in the intercept and/or slope coefficient, and 3) restricting the possible break point to lie within the middle 70% of the sample period. Following Gregory and Hansen, the lag lengths for the ADF-regressions were selected by choosing the smallest lag length for which the t-statistic on the last lag included was less than two in absolute value, with the maximum lag length being equal to eight. Asymptotic critical values were obtained from Table 1 in Gregory and Hansen (1996) using the row  $m = 1$ , ADF\*, C/S. Note that Monte Carlo simulations by Gregory and Hansen (1996) indicate that the actual size of the test in finite sample applications tends to be larger than the nominal size implied by the asymptotic critical values.

The results are reported in columns 2 and 3 of Table 2, where column 2 corresponds to the case where unemployment is used as the dependent variable, and column 3 to

the case where inflation is used as the dependent variable. The results suggest that there is some evidence that Denmark, Norway, and the UK can be added to the set of countries (i.e., the U.S. and Canada) for which inflation and the unemployment rate are cointegrated. On the whole, however, the conclusion that the data do not support the prediction of the model remains unchanged.

It is worth observing that any parameter change that would cause a break in the cointegrating vector would also alter the variance of the error terms (see equation 2.8). Neither the Gregory-Hansen test nor other existing tests for structural breaks in cointegrating regressions account for breaks in the error variance. Under the null of no break, the error variances do not change, thus the size of the test should be unaffected. At issue is the power of the test. As long as there is a break in the cointegrating regression (as opposed to just a break in the error variances) the Gregory-Hansen test should have decent power, although, as it ignores the variance shift, it will not be efficient.

### **3.2 The Triumph of the Natural-Rate Theory?**

In this section we examine the possibility that inflation and unemployment are cointegrated over the first part of the sample, but not the rest. This case, which is not covered by our previous tests, is of particular interest, as it corresponds to a commonly held view among macroeconomists: in the 1960s and early 1970s central bankers attempted to maintain low unemployment rates by trying to exploit the observed Phillips curve relationship, but due to refinements in economic theory, today's central bankers understand that there exists a 'natural' rate of unemployment and do not attempt to push unemployment below this sustainable level. The result of these advances has been an elimination of inflationary bias and time inconsistency problems in monetary policy.

Intuitively, this version of the story asserts that time inconsistency problems were responsible for the rise of inflation in the 1960s and 1970s. The fall of inflation in the post 1980 period, however, occurred after central bankers learned how to avoid time inconsistency problems, possibly as a result of increased independence from political pressures, increased attention to their reputations as inflation fighters, or by tying domestic monetary policy to another central bank that has achieved such independence or reputation (for example, European countries pegging their currencies to the Deutschmark in order to import the Bundesbank's credibility).

The case where there is no inflationary bias corresponds, in the model, to the case where  $k$  takes the value 1. Inspection of Equation (2.8) shows that the implication of  $k = 1$  is that inflation and unemployment are no longer cointegrated. In the absence of inflationary bias, the unemployment rate would still follow the NAIRU, but the policy maker would not be tempted to try to inflate away this employment with surprise inflation. The inflation rate would hover around the policy maker's target, deviating from the target only as a result of the control error. This hypothesis is consistent with the recent experience of a decline in inflation that is not associated with falling unemployment rates in many western European countries. It is less clear, however, whether this account is able to explain the rise of inflation in the 1960s and 1970s.

To test the possibility that time inconsistency is an important consideration in the first but not latter part of the sample, we applied the Johansen cointegration test to data up to 1982:4 for each country.<sup>11</sup> If the story that the KPBG model applies only to the first part of the sample is correct, we would expect to reject the no cointegration null more often using the early sample data than we did when using the full sample. According to the results, presented in Table 3, there is little evidence that inflation and unemployment are cointegrated in the earlier portion of the sample

period. This hypothesis is rejected, in favor of the null of no cointegration, in nine of 13 countries, and only then at the weakest (i.e. 10%) significance level.<sup>12</sup>

It is possible that our failure to reject the null hypothesis of no cointegration here is a result of the low power of standard cointegration tests in the shorter samples. To address this issue, we employ a cointegration test, due to Shin (1994), in which the null hypothesis is that the data are cointegrated. Columns 2 and 3 of Table 3 report the results. The results are somewhat sensitive to whether inflation or unemployment is used as the dependent variable. When we use unemployment as the dependent variable, the null hypothesis of cointegration is largely rejected. When inflation is the dependent variable, the hypothesis of cointegration is rejected at the 5% significance level in four of the 13 countries, and can be rejected at the 10% level of significance in three additional countries. While these results do not constitute overwhelming evidence against the null, when viewed along with the results of the Johansen test they are, in our view, broadly supportive of the notion that inflation and unemployment are not generally cointegrated in the early portion of the sample in OECD countries.

Overall, the results in this section suggest that the empirical shortcomings of the model documented in the previous sections are not solely due to the end of sample data, where one might expect issues related to time inconsistency to play a smaller role in monetary policy. In fact, the model fails to fit the data well, even in the earlier period, where the existence of an inflationary bias is more plausible.

### **3.3 Open Economy Spillovers**

We have seen that while the time inconsistency model fits the U.S. data, it does not do a very good job at explaining inflation in other countries. In this section, we test the version of the model presented in Section 2.1, in which inflationary pressures

generated by a rising U.S. NAIRU spill over into the other OECD economies as a result of monetary authorities' attempts to manage nominal exchange rate movements of their domestic currencies vis a vis the U.S. dollar. Our motivation for using the U.S. as the country of reference is the observation that the U.S. unemployment rate appears to be the only reasonable driver for the observed inflation trends.<sup>13</sup>

The implication of the international spillovers version of the model is that domestic inflation, domestic unemployment, and U.S. unemployment should be cointegrated. According to the results presented in Table 4, these three variables do appear to be cointegrated in many countries. Column 3 presents the Johansen trace statistic for the null hypothesis that there are no cointegrating vectors for domestic inflation, domestic unemployment, and U.S. unemployment in the twelve non-U.S. countries in our sample. The null of no cointegrating vectors is rejected at the 10% level in eight countries, and at the 5% level or greater in six of the 12 countries.<sup>14</sup>

One potential concern here is that our finding of cointegration may be the result of cointegration between domestic and U.S. unemployment rates, while the coefficient on domestic inflation in the cointegrating vector is zero. The theory has nothing to say about whether domestic and foreign unemployment rates ought to be pairwise cointegrated. It predicts cointegration between these three variables due to the response of domestic inflation to both the domestic and foreign NAIRU. If it were the case that the coefficient on domestic inflation in the cointegrating vector found in the data were really zero, then the finding of cointegration would not constitute support for the theory.

We check whether this is the case by applying the ADF residual based test for cointegration. The advantage of this approach is that it allows us to normalize the coefficient on domestic inflation to one. Therefore, a rejection of the null hypothesis of no cointegration using this test implies that the coefficient on domestic inflation

in the cointegrating vector cannot be zero. Column 4 of Table 4 presents the results, which echo the results of the Johansen test in rejecting the null of no cointegration for most countries. This suggests that the coefficient on domestic inflation is not zero, so that the initial finding of cointegration represents support for the theory.

A second potential concern turns out to be more problematic. The open economy version of the KPBG model implies that the coefficients on domestic and foreign unemployment in the cointegrating vector must be non-positive when the coefficient on domestic inflation is normalized to one. Columns 1 and 2 of Table 4 present the estimates of the coefficients on domestic and foreign unemployment rates, respectively, when the coefficient on domestic inflation is normalized to one. While the coefficients on foreign unemployment are negative, as required by the theory, the coefficients on domestic unemployment are uniformly positive. This violates the restrictions of the model and constitutes a point on which the data contradicts the theory.

Can these coefficient estimates be reconciled with the theory? One possibility is that, while the point estimates are positive, these estimates are contaminated by statistical noise and the true coefficients are zero. A finding of a zero coefficient on domestic inflation in a country does not invalidate the theory, as one of two special cases may hold: either 1) domestic unemployment does not enter the central bank's loss function (i.e.  $b = 0$ ) for that country, or 2) the central banker in that country does not suffer from inflationary bias (i.e.  $k = 1$ ). Either of these cases would result in a zero coefficient on domestic unemployment in the cointegrating vector.

We can test whether the coefficient on domestic unemployment is zero by testing to see whether or not domestic inflation and U.S. unemployment are cointegrated. If there is cointegration between all three variables, but also cointegration between domestic inflation and U.S. unemployment, then the coefficient on domestic unemployment in the cointegrating vector in the three variable case has to be zero. The

reasoning is as follows: if a linear combination of two I(1) variables ( $\pi^d$ , and  $u^{U.S.}$ ) is stationary, adding a third I(1) variable ( $u^d$ ) to this stationary combination will not generate an I(0) variable, except in the special case where the coefficient on the third I(1) term is zero.

Conversely, if there is cointegration between all three variables, and no cointegration between domestic inflation and U.S. unemployment, then the coefficient on domestic unemployment in the cointegrating variable cannot be zero. In this case, a linear combination of three I(1) variables ( $\pi^d$ ,  $u^d$ , and  $u^{U.S.}$ ) is stationary. If no linear combination of  $\pi^d$  and  $u^{U.S.}$  is stationary, then it cannot be the case that the coefficient on  $u^d$  in the three variable cointegrating vector is zero (because adding zero to a combination of variables that is non-stationary cannot produce a stationary variable).<sup>15</sup>

Columns 5 and 6 of table 4 present the results of Johansen and ADF tests of the null hypothesis that domestic inflation and U.S. unemployment are not cointegrated. The Johansen test rejects the null of no cointegration in four out of 12 cases, and the ADF test rejects the null in six of 12 countries.<sup>16</sup> In nine of 12 at least one of the tests suggests some support for the hypothesis that domestic inflation and U.S. unemployment are cointegrated. Overall these results provide at least some evidence that the coefficient on domestic unemployment is zero.<sup>17</sup>

The results presented in this section provide some support for the view that the KPBG model, extended to allow for international inflation spillovers, is consistent with OECD inflation trends. However, the interpretation of evidence most favorable to the model suggests that the coefficient on domestic unemployment may be equal to zero. This suggests that the time inconsistency problem was not an important driver of monetary policy in OECD countries other than the U.S.

## 4 Concluding Remarks

In this paper, we have shown that a simple time-consistency story does not provide an adequate explanation for the pattern of rising and then falling inflation rates observed in OECD economies over the past four decades. Allowing for underlying structural changes in the economy did not improve the ability of the model to explain long run inflation trends. Furthermore, the widely held hypothesis that time inconsistency was historically an important feature of monetary policy making that is no longer relevant does not seem to be a solution to the empirical shortcomings of the baseline model.

We found some support for the following version of the time inconsistency hypothesis: U.S. inflation was driven by the interaction of time inconsistency problems and a drifting NAIRU in the U.S. The trend in U.S. inflation in turn spilled over into other OECD countries via the exchange rate. However, our results suggest that the behavior of non-U.S. central bankers with respect to domestic unemployment rates is not well described by the time inconsistency mechanism.

## Notes

<sup>1</sup>The commonality of OECD trend inflation rates, observed in a simple plot of inflation rates, is confirmed by statistical tests showing that OECD inflation rates are cointegrated with one another.

<sup>2</sup>Blinder (1997), for example, “firmly believe(s) that this theoretical problem is a nonproblem in the real world,” stating that “during my brief career as a central banker I never once witnessed or experienced this [the inflationary bias] temptation” (p. 13). McCallum (1995) argues that it is “inappropriate to presume that central banks ... repeatedly engage in fruitless attempts to exploit predetermined but endogenous expectations” (p. 209).

<sup>3</sup>Sargent (1999) refers to this account of the rise and fall in inflation as the “the triumph of the natural-rate theory” hypothesis.

<sup>4</sup>The assumption that the central banker controls inflation only imperfectly mainly serves to break the link between the actual unemployment rate and the NAIRU, allowing us to avoid the necessity of identifying all changes in unemployment with changes in the NAIRU.

<sup>5</sup>Our results are 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.

<sup>6</sup> This is the test statistic described in Ng and Perron (2001) with  $p = 0$  and  $\bar{c} = -7.0$ .

<sup>7</sup>In the bivariate setting with both series assumed to be unit root processes, Jo-

hansen's  $\lambda$ -max and  $\lambda$ -trace tests are equivalent, since in both cases the null hypothesis is that the cointegration rank is zero and the alternative hypothesis is that the cointegration rank is one.

<sup>8</sup>The result for Japan is sensitive to the starting date. If we look for cointegration over the years 1974-2003, we reject the hypothesis of no cointegration between U.S. and Japanese inflation at a 1% significant level.

<sup>9</sup>Austria, Canada, Finland, Italy, Japan, the United Kingdom, and the United States

<sup>10</sup>A number of authors in the literature have investigated the possibility that the slope of the short run Phillips curve,  $\alpha$ , has shifted in recent decades due to factors such as increased openness (Romer 1993, Campillo & Miron 1997, Lane 1997, and Temple 2002) or changes in average inflation rates (Lucas 1972, Ball, Mankiw & Romer 1989 and Akerlof, Dickens & Perry 1996). Alternately, it has been suggested that the central banker's preference parameters,  $k$  and  $b$ , may have shifted, perhaps due to increases in central bank independence (Alesina & Summers 1993, Campillo & Miron 1997, Temple 1998 and Brumm 2000). These factors, and others like them, represent potential sources of parameter changes that might affect the cointegrating vector given by (2.8).

<sup>11</sup>We chose this date as it is widely considered to correspond to a regime change in the conduct of U.S. monetary policy (Clarida, Gali, and Gertler 1999)). Our results are robust to alternate choices.

<sup>12</sup>To control for the possibility of contemporaneously correlated country errors, we also applied the Westerlund test described in section 3.2 to the early sample data. For the sample of 7 countries for which data goes back to 1964:1, the test statistic is

0.56, while for the sample using countries going back to 1970:1, the test statistic is 3.07. In neither case does the tests statistic fall below the 5% critical value of -1.65, implying that we cannot reject the null that the inflation-unemployment pairs are not cointegrated.

<sup>13</sup>In practice, European countries may have been more interested in managing exchange rates relative to the Deutschmark, for example, but German unemployment does not rise and fall in such a way as to make this a reasonable explanation of the rise and fall of inflation.

<sup>14</sup> Three of the countries for which there is no evidence of cointegration between these three variables are Germany, Japan, and the UK. It is perhaps not surprising that the model doesn't fit for these countries, given the stature of their currencies and the fact that the small open economy assumption on which the model is based is perhaps less applicable.

<sup>15</sup>Note that this logic implies that the coefficient on U.S. inflation cannot be zero, because the results of previous sections established that domestic inflation and domestic unemployment are not cointegrated in the countries in our data set.

<sup>16</sup>The discrepancy between the two sets of results may be due to differences in power, though power comparisons have yielded mixed results (Boswijk & Frances 1992, Haug 1996, and Pesavanto 2004).

<sup>17</sup>Our test of a zero coefficient on domestic unemployment assumes that there is a unique (up to a scalar multiple) cointegrating vector between domestic inflation, domestic unemployment, and foreign unemployment. If there are multiple cointegrating vectors, a finding of a zero coefficient on domestic unemployment would apply only to one of these cointegrating vectors. A Johansen test of the null hypothesis of one or

fewer cointegrating vectors rejects the null in favor of multiple cointegrating vectors at the 5% level in only two countries: Australia and Italy.

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Table 1. Data Overview

Country	Sample	Unit Root Tests <sup>a</sup>		Cointegration Test <sup>b</sup>
		$\pi$	$u$	Johansen $\lambda$ -max
Australia	1966:3-2003:3	-1.56	-0.55	15.86**
Austria	1964:1-2003:3	-1.11	0.28	8.85
Canada	1964:1-2003:3	-1.55	-0.75	25.09***
Denmark	1970:1-2003:3	-0.79	-0.77	20.92***
Finland	1964:1-2003:3	-1.01	-0.55	16.04**
France	1967:4-2003:3	-1.10	-0.13	19.69***
Germany	1964:1-2003:3	-1.34	-0.40	14.35*
Italy	1964:1-2003:3	-1.39	-0.29	19.99***
Japan	1964:1-2003:3	-0.94	0.69	7.84
Norway	1964:1-2002:2	-1.26	-0.94	13.25*
Sweden	1970:1-2003:3	-1.18	-1.01	17.00**
United Kingdom	1964:1-2003:3	-1.66*	-1.29	16.01**
United States	1964:1-2003:3	-0.93	-1.78 *	N.A.

Notes: \* denotes reject  $H_0$  at the 10% level, \*\* at the 5% level, and \*\*\* at the 1% level.

a. For the unit root tests, the test statistic is the test statistic in Ng and Perron (2001). Asymptotic critical values for the statistic are: -1.62 (10%), -1.98 (5%), -2.58 (1%).  $H_0$ : The series possesses a unit root.  $\pi$  is inflation and  $u$  is unemployment.

b. Test of the null hypothesis that a country's inflation rate is not cointegrated with U.S. inflation. For the cointegration tests, the sample period is 1964:1-2003:3, except for Denmark, which is 1967:1-2003:3. As unit root tests suggest that our time series are  $I(1)$ , for all bivariate cases we report Johansen's  $\lambda$ -max test statistic, which tests the null of no-cointegration against the alternative that the two series are cointegrated. All tests are performed with unrestricted intercepts and no deterministic time trend component in the VECM. The null distribution of the test statistic assumes that the VECM errors are serially uncorrelated. We selected a relatively long lag length (8 lags) in the VECM in the hope of alleviating any serial correlation problems. Critical values are obtained using the procedure of MacKinnon, Haug, and Michelis (1999). Critical values for the statistic are: 12.30 (10%), 14.26 (5%), 18.52 (1%).

Table 2. Cointegration Tests: Inflation and Unemployment

Country	Johansen <sup>a</sup>	Gregory-Hansen <sup>b</sup>	
	$\lambda$ -max	u	$\pi$
Australia	8.89	-2.72	-3.36
Austria	11.07	-2.81	-4.53
Canada	14.03*	-4.23	-3.61
Denmark	9.29	-3.11	-4.82*
Finland	7.66	-3.09	-2.77
France	6.99	-2.89	-3.34
Germany	11.59	-4.06	-2.98
Italy	8.77	-2.99	-3.01
Japan	6.73	-5.51**	-4.98**
Norway	8.07	-3.69	-4.75*
Sweden	9.01	-4.56	-4.23
United Kingdom	5.15	-2.85	-3.78
United States	15.15**	-3.22	-4.30

Notes: \* denotes reject  $H_0$  at the 10% level, \*\* at the 5% level, and \*\*\* at the 1% level.

a. See Note *b* to Table 1.

b. The test statistic is the ADF\* statistic in Gregory-Hansen (1996). The null hypothesis is that the unemployment and inflation rates are not cointegrated. The alternative hypothesis is that they are cointegrated with a possible break in the intercept and slope of the cointegrating regression. The asymptotic critical values (from Gregory-Hansen, Table 1,  $m = 1$ , C/S) are: -4.68 (10%), -4.95 (5%), -5.47 (1%). The column labelled *u* reports the results when the unemployment rate is used as the dependent variable. The column labelled  $\pi$  reports the results when the inflation rate is used as the dependent variable.

Table 3. Inflation and Unemployment Cointegration Tests

Country	Johansen Test <sup>a</sup>	Shin Test <sup>b</sup>	
	$\lambda$ -max	$u$	$\pi$
Australia	7.18	1.024***	0.250*
Austria	5.10	0.210	0.167
Canada	7.48	0.293*	0.140
Denmark	12.49*	0.946***	0.061
Finland	7.67	0.725***	0.410**
France	14.00*	0.585***	0.151
Germany	5.59	1.097***	0.347**
Italy	6.34	0.373**	0.363**
Japan	8.93	1.337***	0.222
Norway	12.98*	0.195	0.090
Sweden	12.69*	0.473**	0.253*
United Kingdom	5.79	0.728***	0.482**
United States	6.93	0.348**	0.245*

Notes: \* denotes reject  $H_0$  at the 10% level, \*\* at the 5% level, and \*\*\* at the 1% level.

a. See Note *b* to Table 1.

b. Shin's  $C$ - $\mu$  statistic with  $K = T^{1/3}$  and  $L = 3$ .  $H_0$ : unemployment and inflation are cointegrated. The column labelled  $u$  reports the results when the unemployment rate is used as the dependent variable. The column labelled  $\pi$  reports the results when the inflation rate is used as the dependent variable.

Table 4. Tests for Inflation Spillovers

Inflation, Unemployment, and U.S. Unemployment						
	Coefficient estimates		Tri-variate Model		Bi-variate Model	
	$u^d$	$u^{US}$	$\lambda$ -trace <sup>a</sup>	ADF <sup>b</sup>	$\lambda$ -max <sup>a</sup>	ADF <sup>b</sup>
Australia	0.954	-1.968	28.88*	-3.43**	9.05	-2.66*
Austria	0.502	-0.939	37.18***	-4.01**	11.83	-2.98**
Canada	0.955	-2.012	33.27**	-3.89**	12.29	-2.47
Denmark	0.753	-1.717	27.72*	-3.89**	12.02	-2.56*
Finland	0.536	-1.495	19.56	-2.98	9.29	-2.01
France	0.701	-1.729	35.30**	-4.16***	18.34**	-2.04
Germany	0.348	-0.841	24.82	-3.24*	11.62	-2.25
Italy	0.781	-2.367	37.17***	-3.80**	15.30**	-2.48
Japan	1.459	-1.155	18.14	-3.57**	6.91	-2.95**
Norway	1.379	-1.560	31.07**	-3.46**	12.46*	-2.03
Sweden	0.965	-1.465	39.04***	-4.32***	25.20***	-2.77*
U.K.	1.184	-2.205	24.84	-3.33*	11.54	-2.84**
1%	-	-	35.46	-4.04	18.50	-3.40
5%	-	-	29.80	-3.39	14.26	-2.82
10%	-	-	27.07	-3.06	12.30	-2.49

Notes: \* denotes reject  $H_0$  at the 10% level, \*\* at the 5% level, and \*\*\* at the 1% level. The last 3 rows of the table report the critical values of the relevant test statistics at the 1%, 5%, and 10% levels.

a. Johansen cointegration tests. See Note *b* to Table 1. For the tri-variate case, we report the  $\lambda$ -trace statistic, to allow for the possibility of more than one cointegrating vector.

b. The test statistic reported is the ADF test on the residuals of an OLS estimation of the cointegrating vector. 4 lags of the residual were included in the ADF regression. The null hypothesis is that there is no cointegrating vector.