Cross Country Evidence on Monetary Indeterminacy*

Pierre Chaussé†
Matthew Doyle‡
Jean-Paul Lam§

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Abstract

A leading explanation of long run U.S. inflation trends attributes both the fall in the level and volatility inflation in the early 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 fall of inflation, as well as subsequent low and stable inflation over the same period. In this paper we exploit the international dimension of the fall of inflation and its volatility to investigate the hypothesis that good monetary policy is responsible for recent inflation outcomes. Our results, while not conclusive, suggest that the theory does not provide a compelling explanation of the cross country data.

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

‡Department of Economics, University of Waterloo. Email: m6doyle@uwaterloo.ca.

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

1 Introduction

The new-Keynesian Taylor rule approach toward monetary policy provides the current standard answer to the question "why was inflation high and volatile in the 1970s, but low and stable in the 1980s and 1990s in the U.S.?" According to this theory, U.S. monetary policy accommodated inflationary expectations in the 1970s leading to equilibrium indeterminacy, allowing inflation expectations to become self-fulfilling. In the 1980s, however, monetary policy became more "active" in response to inflationary expectations, thereby removing the indeterminacy and pinning inflation at the Fed's target rate.

While the U.S. has been most widely studied, it is well known that the volatility of inflation fell across OECD countries during the episode known as the Great Moderation. Furthermore, inflation levels across countries are strongly correlated both at trend and business cycle frequencies.[1] Consequently, a good explanation of the level and volatility of inflation ought to be consistent with the cross country character of inflationary episodes. T

In this paper, we ask whether the standard explanation is consistent with the pattern of inflation outcomes in OECD countries, other than the U.S. We approach this question 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 Coibion and Gorodnichenko (2011) by employing a New Keynesian model with trend inflation to determine an individual region of determinacy for each country. We follow Clarida et al. (2000) by using GMM to estimate policy reaction functions[2] across OECD countries and statistically test whether the estimated parameters of each country's policy rule are in the region that the theory suggests.

There are a number of problems with cross country studies of this nature. First, is the need to compress a variety of cross country institutional detail into a single

[1] See Rogoff (2004), Doyle and Falk (2008), Monacelli and Sala (2009) , Ciccaralli and Mojon (2010), and Mumtaz and Surico (2012).

[2] This approach has been criticized by Cochrane (2011) who argues that such methods cannot recover policy rule parameters in New Keynesian models. Simms (2008), however, shows that this critique relies on the assumption that the central bank responds one-to-one to changes in the Wicksellian natural rate of interest. This assumption is not a feature of the policy rules we attempt to estimate.

overarching framework, such as a Taylor type specification. As a practical matter, central bank behaviour is not well described by an interest rate rule across a wide sample of countries going back to the 1960s. Consequently, it is not clear that fitting a conventional looking monetary policy rule to an array of countries is appropriate.

We address this issue by focusing only on the more recent period. 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. 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.[3]

A limitation of this approach is that it represents a relatively weak test of the theory as it means we cannot test for changes in the conduct of monetary policy corresponding to changes in inflationary regimes. Instead we are restricted to the question of whether or not policy was such that equilibrium indeterminacy was eliminated in OECD countries during the period where inflation was low and stable. This means that our approach cannot be readily extended to tests alternate hypotheses that rely on changing monetary policy but work through mechanisms other than indeterminacy.[4]

A second problem concerns data availability.[5] For analyses of U.S. monetary, Greenbook data allows policy research to use inflation and output gap forecasts that were known to the FOMC at of policy formation. The existence of data on the central bank's expectations allows policy rules to be estimated simply, with OLS. Corresponding data does not exist for other OECD countries.[6] As a result, it is necessary to use an alternative econometrics approach, such as GMM, which has been criticized as suffering from

[3]See Clarida et al. (1998), for example.

[4]For example, Sargent (1999), Cogley and J. (2005), Cogley and J. (2001), Ireland (1999), Primiceri (2006) or DiCecio and Nelson (2009).

[5]It is worth noting that we use revised rather than real time data throughout our paper. Orphanides (2001) argues that data misperceptions, rather than bad policy, caused the indeterminacy of monetary policy in the 1970s. Following his work, the U.S. literature on Taylor rules has largely employed real time data. Our goal, however, is not to explain why policy is or is not in the determinacy region, but simply to measure whether it is or is not.

[6]Though Nikolsko-Rzhevskyy (2011) creates pseudo Greenbook forecasts for three countries.

weak identification, low power, and poor small sample properties.[7]

We address the econometric limitations of GMM by employ the bootstrapping procedure of Inoue and Shintani (2006). Although we cannot address the issue of weak identification, we can construct tests and intervals that are more reliable. In fact, intervals based on Wald statistics are not robust to weak identification, because, in that case, the probability of the true intervals to be unbounded is positive, while intervals based on the Wald statistics are bounded with probability 1 (see Dufour (1997)). Bootstrap tests are not only more robust to weak identification, they are also likely to be more reliable in small samples. Furthermore, the determinacy frontier is a nonlinear function of the structural parameters. Bootstrap tests when the null hypothesis is a nonlinear function of the parameters are in general more accurate than tests based on approximations such as the Delta method.

Aside from the cross country focus, our paper is similar in its objectives to work by Lubik and Schorfheide (2004), Boivin and Giannoni (2006), and Coibion and Gorodnichenko (2011) who attempt to identify whether U.S. monetary policy is in the determinacy vs indeterminacy region in different time period. As in these papers, we recognize that determinacy is in principle a property of all the parameters of the model. Our approach is closest to that of Coibion and Gorodnichenko (2011), and in contrast to Lubik and Schorfheide (2004) and Boivin and Giannoni (2006), in that we estimate Taylor rule parameters independently of, rather than jointly with, the other structural parameters of the DSGE model. Relative to these papers, which attempt to measure if and when monetary policy moved from indeterminacy to determinacy, our focus on the more recent period imposes a null hypothesis that monetary policy is in the determinacy region. We simply seek to test this hypothesis.

We find little evidence in favour of the standard explanation for countries other than the U.S. Overall, our evidence is consistent with the findings of much of the existing literature, in that U.S. monetary policy appears to be in the determinacy region in recent decades. However, for other countries the theory fails to fit well. While our results do not constitute a decisive rejection of the theory, we interpret the combination of a weak test with mixed rejections of the hypothesis as suggesting that the theory does not provide a compelling explanation of the fall of inflation observed in most OECD

[7] Mavroeidis (2004), Urquiza (2010).

countries.

The paper proceeds as follows: in Section 2 we briefly present the theoretical model and show how the determinacy region depends on model parameters. In section 3 we outline our estimation procedure. In Section 4 we present our results. Section 5 concludes.

2 Model

We employ a standard New Keynesian model incorporating trend inflation and a unit root for technology. It is, by now, well known that positive trend inflation modifies the Taylor principle in New Keynesian models.[8] Essentially, as the steady state rate of inflation increases, a rise in the relative price distortion drives the steady state level of output down. This turns out to have important implications for monetary policy.

In particular, with trend inflation different from zero, achieving a unique rational expectations equilibrium requires that the central bank respond more strongly to inflation that implied by the Taylor principle, and that the strength of this reaction depends on the level of trend inflation. Furthermore, economic indeterminacy is not a simple function of the central bank's response to inflation. Rather, there exists a region in the parameter space of the model for which determinacy is ensured.

The model is otherwise a standard New Keynesian model. A representative consumer chooses consumption of a final good and firm-specific labour supply in each period to maximize the present discounted value of utility, with a discount factor given by β. Utility is separable over consumption and leisure, with log utility for consumption and Frisch labour supply elasticity of η.

The final good is a Dixit-Stiglitz aggregate of a continuum of intermediate goods, each produced by a monopolist using a constant returns to scale production function with firm-specific labour as the only input. The elasticity of substitution of intermediate goods in the aggregator is θ. The aggregate productivity level follows a random walk. Firms follow Calvo (1983) pricing. In particular, it is assumed that in each period firms face a constant probability λ of not being able to change their price. Firms that have the opportunity to change their price in a given period choose the reset price so as to

[8]Hornstein and Wolman (2005), Kiley (2007) Ascari and Ropele (2009).

maximize the expected discounted value of future profits.

There is no investment, government spending or international trade.

The exact specification of the model is that of Coibion and Gorodnichenko (2011). As a full derivation is available elsewhere,[9] we simply present the log-linearized system of equations. The dynamic IS curve is given by:

$$E_t gy_{t+1} = i_t - E_t \pi_{t+1}, \tag{2.1}$$

where gy is the growth rate of output, i is the nominal interest rate, and π is the rate of inflation. All variables are expressed in terms of deviations from the log of their steady state values, denoted by $\overline{GY}$, $\overline{I}$ and $\overline{\Pi}$ respectively.

The other key equation is the reset price b_t chosen by a firm that is able to change its price in period t. The log-linearised optimal reset price is given by:

$$\begin{aligned}(1+\theta\eta^{-1})b_t = (1+\eta^{-1})(1-\gamma_2)\sum_{j=0}^{\infty}\gamma_2^j E_t x_{t+j} + E_t\sum_{j=1}^{\infty}(\gamma_2^j - \gamma_1^j)(gy_{t+j} - r_{t+j-1}) \\ + \sum_{j=1}^{\infty}[\gamma_2^j(1+\theta(1+\eta^{-1})) - \gamma_1^j]E_t\pi_{t+j}, \end{aligned} \tag{2.2}$$

where x is the output gap, defined as the log-deviation from the flexible price equilibrium level of output. Note that $\gamma_1 = \lambda\overline{R}^{-1}\overline{GY}\ \overline{\Pi}^{\theta}$, $\gamma_2 = \gamma_1\overline{\Pi}^{1+\theta/\eta}$, and that $\gamma_2 = \gamma_1$ when trend inflation is zero.

Two other relationships are useful. First is the relationship between the optimal reset price, b_t and the inflation rate, π_t, which takes the form:

$$\pi_t = \left(\frac{1-\lambda\overline{\Pi}^{\theta-1}}{\lambda\overline{\Pi}^{\theta-1}}\right) b_t. \tag{2.3}$$

Higher levels of trend inflation tend to make actual inflation less sensitive to the reset price. This is because, on average, firms that reset their prices set them above the level set by others, which implies that those firms account for a smaller share of expenditures than others.

Finally, given the assumption of a unit root in technology, the relationship between output and the output gap is such that:

$$gy_t = x_t - x_{t-1} + \epsilon_t^a, \tag{2.4}$$

[9] Coibion and Gorodnichenko (2008).

where ϵ_t^a is the innovation to technology at time t. Equations 2.2, 2.3, and 2.4 constitute the new Keynesian Phillips curve, with trend inflation.

2.1 Calibration

The model with trend inflation cannot be solved analytically, hence any results concerning determinacy are necessarily numerical. Here, we follow Coibion and Gorodnichenko (2011) by calibrating the non-policy parameters of the model and estimating the policy rule parameters separately, as opposed to Lubik and Schorfheide (2004) and Boivin and Giannoni (2006) who attempt to estimate the full DSGE model simultaneously with the policy rule. The advantage of this approach is that it is requires us to impose relatively few additional restrictions on the estimation of the Taylor rule. The disadvantage is that it requires us to calibrate some parameter values for the model.

We calibrate the model as follows. The Frisch labour supply elasticity, η, is set to 1, and the discount factor, β, is set to 0.99. We calibrate the steady state growth rate to be 1.5 percent per year, which corresponds to the U.S. per capita GDP rate from 1969 to 2002. We set the elasticity of substitution, θ, to 10, which corresponds to a markup of 11%. This is consistent with estimates in the literature for the U.S. The degree of price stickiness, λ, is set to 0.55, which amounts to firms resetting prices every seven months on average. Again, this is consistent with estimates in the literature concerning the U.S.

2.2 Monetary Policy Rule

To close the model, we need to specify the policy rule of the central bank, which determines the nominal interest rate. The literature has proposed many different versions of the Taylor rule. For our baseline case, we suppose that central bank reaction functions take a generalized Taylor rule form:

$$i_t = \alpha + \rho_1 i_{t-1} + \rho_2 i_{t-2} + (1 - \rho_1 - \rho_2)(\phi_x x_t + \phi_\pi \pi_{t+1}) + \varepsilon_t, \tag{2.5}$$

where α is the, constant, steady state target for the nominal interest rate, ρ_1 and ρ_2 are interest rate smoothing parameters, ϕ_x is the response to the output gap, ϕ_π is the response to expected inflation, and ε_t. This specification allows for interest rate smoothing up to order two, as well as a response to both inflation and the output gap.

This specification of the interest rate rule has been used in several papers in the literature Clarida et al. (1998); Coibion and Gorodnichenko (2008) to describe the behaviour of the central bank. We include interest rate smoothing in the policy rule to reflect the fact that central banks often choose a smooth path for interest rate and that reversals of interest rates are infrequent.

2.3 Determinacy

Given the monetary policy rule from Equation 2.5, and the calibration from section 2.1, we can numerically calculate the region of determinacy in terms of the parameters of the monetary policy rule and trend inflation. Since the determinacy region is a function of several parameters, we present illustrative cross sections. In all cases the determinacy region lies above the frontier and the indeterminacy region lies below it.

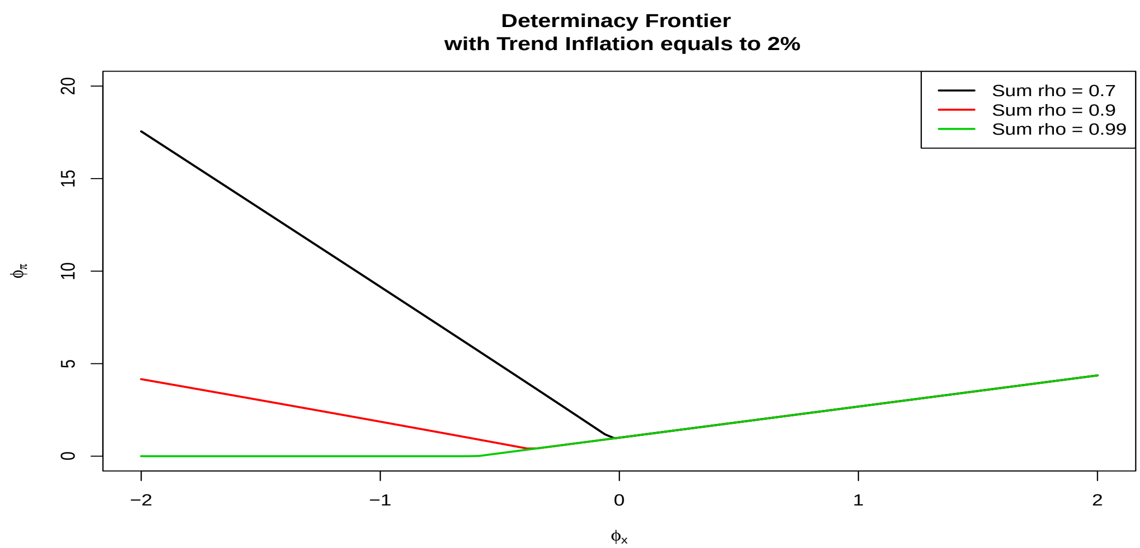
Figure 1 shows the shape of the different frontier in terms of the response coefficients on output and inflation, for a trend inflation rate of 2%. It is apparent from the figure that, given the smoothing parameters and trend inflation, the response to inflation, ϕ_π, necessary to induce determinacy is a non-monotonic function of the coefficient on the output gap, ϕ_x. Furthermore, the figure shows that whether the determinacy is induced or not, given values of ϕ_π and ϕ_x, depends on the amount of interest rate smoothing done, i.e. on $\rho_1 + \rho_2$. As the figure shows, with trend inflation set to 2%, the degree of interest rate smoothing only impacts determinacy for cases where the response to the output gap is negative.

Figure 2 illustrates the effect that changing trend inflation has on the determinacy frontier. Note that, in general, a higher rate of trend inflation makes the determinacy region smaller. That is, for a given response to the output gap, the coefficient on inflation in the policy rule must increase as trend inflation increases. This is the result highlighted by Hornstein and Wolman (2005), Kiley (2007), and Ascari and Ropele (2009). Note that this result generalizes for different degrees of interest rate smoothing.

3 Estimation

We use GMM to estimate Equation 2.5 using the 14×1 vector of instruments $Z_t = \{1, i_{t-1}, ..., i_{t-4}, y_t, y_{t-1}, ..., y_{t-4}, \pi_t, \pi_{t-1}, ..., \pi_{t-3}\}'$. The estimation is therefore

Figure 1: Determinacy frontier for different values of $\rho_1 + \rho_2$



based on the moment conditions:

$$E(g(X_t, \theta)) = E\left[\varepsilon_t(\theta) Z_t\right] = 0,$$

where $X_t = \{i_t, \pi_{t+1}, Z_t'\}'$, and $\theta = \{\alpha, \rho_1, \rho_2, \phi_x, \phi_\pi\}'$. The estimate is

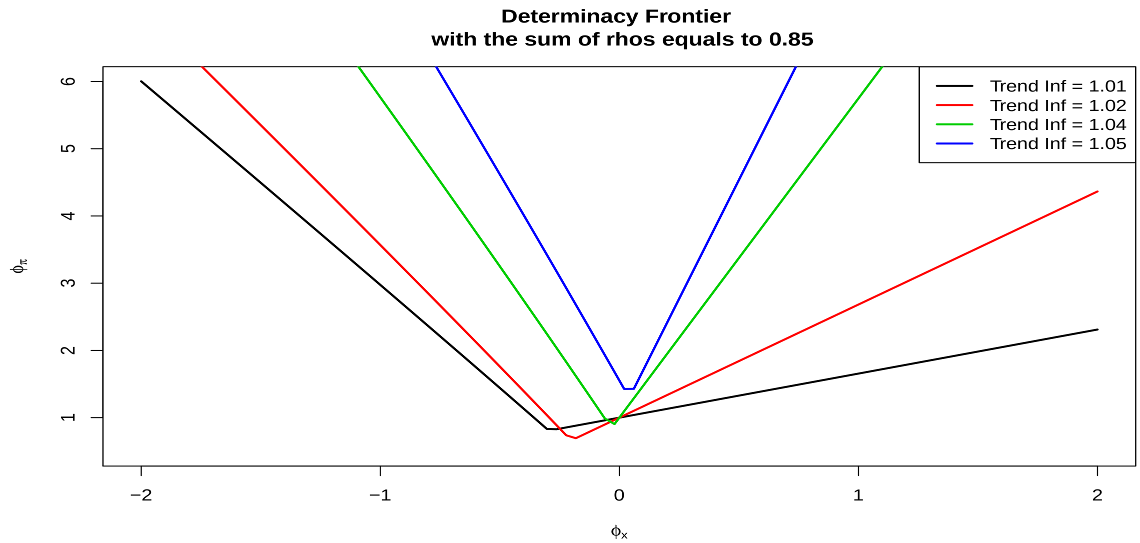
$$\hat{\theta} = \arg\min_{\theta} \bar{g}(\theta)' \hat{\Omega}(\tilde{\theta})^{-1} \bar{g}(\theta),$$

where $\tilde{\theta}$ is a first step estimate, and $\hat{\Omega}(\theta)$ is the HAC estimate of the moment conditions.

Our question of interest is whether the policy parameters ϕ_x and ϕ_π lie within the region of determinacy. According to the theoretical model, however, the frontier of determinacy is both non-linear and a function of $\rho_1 + \rho_2$. The relevant hypothesis test, therefore, is in fact $H0 : \phi_\pi \geq f(\phi_x; \rho_1 + \rho_2)$. It also depends on all parameters of the model including the trend inflation, but since they are not estimated, we consider them as being fixed parameters.

Let $\eta = \phi_\pi - f(\phi_x; \rho_1 + \rho_2)$, so that the null hypothesis becomes $H0 : \eta \geq 0$. Then, we need to obtain a consistent estimate of the variance of $\hat{\eta} = \hat{\phi}_\pi - f(\hat{\phi}_x; \hat{\rho}_1 + \hat{\rho}_2)$. Because of the nonlinearity of the frontier, we can obtain the variance using the Delta

Figure 2: Determinacy frontier for different values of $\bar{\pi}$



method. A first order Taylor expansion gives:

$$\hat{\eta} - \eta \approx (\hat{\phi}_\pi - \phi_\pi) - f_{\phi_x}(\hat{\phi}_x - \phi_x) - f_{\rho_1+\rho_2}[(\hat{\rho}_1 - \rho_1) + (\hat{\rho}_2 - \rho_2)],$$

where f_{θ_i} is the derivative of the frontier with respect to the parameter θ_i. Since there is no closed form representation of the frontier, we use a two-sided finite difference to compute the derivatives. We therefore approximate the variance by:

$$\hat{\sigma}^2_{\hat{\eta}} \approx \begin{pmatrix} 0 & -f_{\hat{\rho}_1+\hat{\rho}_2} & -f_{\hat{\rho}_1+\hat{\rho}_2} & -f_{\hat{\phi}_x} & 1 \end{pmatrix} \hat{V}(\hat{\theta}) \begin{pmatrix} 0 \\ -f_{\hat{\rho}_1+\hat{\rho}_2} \\ -f_{\hat{\rho}_1+\hat{\rho}_2} \\ -f_{\hat{\phi}_x} \\ 1 \end{pmatrix}$$

where $\hat{V}(\hat{\theta})$ is the GMM estimate of the variance of $\hat{\theta}$. This results in the test statistic:

$$\hat{\tau} = \frac{\hat{\phi}_\pi - f(\hat{\phi}_x, \hat{\rho}_1 + \hat{\rho}_2)}{\hat{\sigma}_{\hat{\eta}}}. \tag{3.6}$$

Since the Delta method produces a consistent estimate of the standard error whenever the coefficient estimates are consistent, $\hat{\tau}$ is asymptotically $N(0, 1)$.

The poor performance of GMM in small samples suggests that the test should be performed using bootstrap methods. In response, we apply the bootstrap procedure of Inoue and Shintani (2006). Our approach takes advantage of the fact that the asymptotic distribution of $\hat{\tau}$ is pivotal, meaning that it does not depend on any unknown parameters. Given some regularity conditions, bootstrap tests based on asymptotically pivotal statistics benefit from asymptotic refinement, meaning they approximate the finite sample distribution better than its asymptotic counterpart.

Inoue and Shintani (2006) propose an overlapping block bootstrap of $X_t' = \{i_t, \pi_{t+1}, Z_t'\}$, with an optimal weighting matrix which produces asymptotic refinement. According to their approach, the length of each block, l, is selected using the method of Andrews (1991) for truncated kernels. If T/l is not an integer, the length of the last block is simply adjusted. This gives $b = \lfloor T/l \rfloor$ different blocks, where $\lfloor x \rfloor$ is the nearest integer less or equal to x. The j^{th} block, for $j = 1, ..., (b-1)$, is obtained by drawing N_j from the uniform distribution over $\{0, 1, ..., T-l\}$, and forming the matrix $\{X_{N_j+1}, ..., X_{N_j+l}\}'$. The b^{th} block is $\{X_{N_b+1}, ..., X_{N_b+(T-(b-1)l)}\}'$, where N_b is drawn from the same distribution.

In order to estimate the distribution of $\hat{\eta}$ under the null, we need the moment conditions, using the bootstrap probability measure, to be exactly satisfied at $\eta = \hat{\eta}$. Therefore, the bootstrap estimates are obtained by centering the sample moments around μ^* defined as:

$$\mu^* = \frac{1}{T-l+1} \sum_{t=0}^{T-l} \frac{1}{l} \sum_{i=1}^{l} g(X_{t+i}, \hat{\theta})$$

Let $g(X_{ik}^*, \theta)$ be the 14×1 vector of the i^{th} moment observation that belongs to the k^{th} block, for $k = 1, ..., b$. The author propose to use the following estimate of Ω:

$$\hat{\Omega}^* = \frac{1}{T} \sum_{k=1}^{b} \sum_{i=1}^{l_k} \sum_{j=1}^{l_k} [g(X_{ik}^*, \tilde{\theta}^*) - \mu^*][g(X_{jk}^*, \tilde{\theta}^*) - \mu^*]'$$

where $l_k = k$ for $k = 1, ..., (b-1)$, $l_b = T - (b-1)l$, and $\tilde{\theta}^*$ is a first step estimate using a first step weighting matrix, which could be the identity matrix or the inverse of $Z'Z$ as long as it is positive definite and that it is the same used to obtain the $\hat{\theta}$. The estimate is then obtained as follows:

$$\theta^* = \arg\min_{\theta} [\bar{g}(\theta)^* - \mu^*]' \hat{\Omega}^{*-1} [\bar{g}(\theta)^* - \mu^*]$$

with

$$\bar{g}(\theta)^* = \frac{1}{T}\sum_{k=1}^{b}\sum_{i=1}^{l_k} g(X_{ik}^*, \theta)$$

We can then compute the bootstrap statistic as:

$$\tau^* = \frac{[\phi_\pi^* - f(\phi_x^*, \rho_1^* + \rho_2^*)] - [\hat{\phi}_\pi - f(\hat{\phi}_x, \hat{\rho}_1 + \hat{\rho}_2)]}{\sigma_{\hat{\eta}^*}}.$$

where the standard error is obtained the same way we computed $\hat{\sigma}_{\hat{\eta}}^2$:

$$\sigma_{\hat{\eta}^*}^2 = \begin{pmatrix} 0 & -f_{\rho_1^*+\rho_2^*} & -f_{\rho_1^*+\rho_2^*} & -f_{\phi_x^*} & 1 \end{pmatrix} \hat{V}(\theta^*) \begin{pmatrix} 0 \\ -f_{\rho_1^*+\rho_2^*} \\ -f_{\rho_1^*+\rho_2^*} \\ -f_{\phi_x^*} \\ 1 \end{pmatrix}$$

If we compute B different τ_i^*, the p-value of the bootstrap test is be:

$$Pv^* = \frac{1}{B}\sum_{i=1}^{B} 1\{\hat{\tau} > \tau_i^*\}$$

where $1\{\}$ is the indicator function. In other words, we reject if the probability of being below $\hat{\tau}$ using the distribution of τ^* is small.

The necessity of our bootstrapping approach can be confirmed by an examination of the estimated distribution of $\hat{\tau}$. Figures 3 and 4 show the estimated distribution of $\hat{\tau}$ for the countries in our sample, and compare these with the asymptotic N(0,1) density. It is clear that the true distribution deviates substantially from a standard normal distribution, and is often skewed either to the left of right. Running a test based directly on the test statistic given in Equation 3.6 would therefore lead to poor results.

4 Data and Results

4.1 Data

We use data from the OECD's Main Economic Indicators for a sample of twelve OECD countries. Since historical, consistent GDP data is unavailable for a number of countries, we follow Ciccaralli and Mojon (2010) and use Industrial Production as a measure of output. We create an output gap series by applying a Band Pass filter to the data. [QUESTION: what is the parameterization of the filter?] For inflation, we use the annualized percentage change in the CPI. For the short term nominal interest rate we use a mix of overnight commercial paper rates, depending upon what is available for each country.

4.2 Baseline Results

This section presents the results for our baseline specification. We experiment with several alternate specifications in subsequent sections.

We use a sample of data from 1979Q4 to 1998Q4. The start date corresponds to the appointment of Paul Volker as Chairman of the Federal Reserve system in the U.S. and is often used as a potential break point in U.S. monetary policy.[10] The latter date corresponds to the last period of national currencies for those countries joining the Euro.

The determinacy frontier for each country is calculated using the model calibration presented in Section 2.1. For each country, we use 2% inflation as our measure of trend inflation. This approach associates trend inflation with the target rate of inflation of the central bank. Data on central bank inflation targets dating back the 1980s are generally unavailable. However, inflation rates stabilized somewhere in the are of 2% for many OECD countries by the 1990s. For our baseline results, we assume that 2% represents a common inflation target for OECD central banks.[11]

The interest smoothing parameters ρ_1 and ρ_2 are estimated country by country. In order to produce sensible estimates, we restrict $\rho_1 + \rho_2$ to be less or equal to 0.98. Theoretically, it does not make sense to allow the sum to be greater than one. Numerically it makes the estimates very unstable because of the weak identification problem that arises when $\rho_1 + \rho_2$ is close to 1. Note also that, according to Andrews (2000), bootstrapping is inconsistent in the case where $\rho_1 + \rho_2 = 1$.

Table 1 presents the results. The first five columns present the estimates for the

[10] We have been experimented with later start dates. While this does not change the qualitative nature of the results overall, we do find that, in contrast to some of the literature, the U.S. coefficient lies outside the determinacy region when the start date is set to 1984Q1 or later.

[11] We investigate the implications of this assumption in section 4.4 .

Table 1: Taylor Rule Estimates and Determinacy Tests

	α	ρ_1	ρ_2	ϕ_x	ϕ_π	D_{ϕ_π}	P-value	Det. %
Belgium	0.18	0.80	0.13	1.51	1.31	3.54	0.08	0.04
Canada	1.28	0.88	-0.20	0.30	1.32	1.51	0.49	0.25
Denmark	0.05	1.24	-0.27	1.02	0.70	2.72	0.02	0.04
Finland	0.08	1.40	-0.42	0.46	0.59	1.78	0.05	0.12
France	0.45	1.06	-0.16	0.69	0.85	2.16	0.21	0.04
Germany	0.27	1.21	-0.30	0.79	0.67	2.32	0.03	0.00
Ireland	1.43	0.88	-0.09	0.38	0.60	1.64	0.11	0.03
Italy	0.20	0.87	0.05	1.24	1.15	3.08	0.03	0.02
Japan	0.09	1.35	-0.40	0.09	0.29	1.15	0.31	0.23
Spain	0.33	0.64	0.28	1.66	1.18	3.79	0.07	0.09
Switzerland	0.07	1.33	-0.40	0.17	0.93	1.28	0.53	0.22
UK	0.34	1.14	-0.19	2.99	0.33	6.03	0.03	0.00
US	-0.35	0.71	0.16	0.40	2.10	1.68	0.73	0.54

policy rule parameters. The sixth column shows the value of the policy rule coefficient on inflation that would be on the determinacy frontier, given the other estimated coefficients and calibrated parameters.

We provide two pieces of information concerning the statistical likelihood that the estimated policy rule coefficients lie within the determinacy region. The first, reported in the seventh column of table 1, is the p-value for our test of the null hypothesis that the policy rule coefficients are in the determinacy region. Given what we know about the performance of GMM in small samples, formal statistical tests are likely of low power. We supplement the formal p-values by also report the proportion of point estimates from our bootstrap process lies within the determinacy region. This proportion is reported in the final column of the table.

The first thing to note from the table is that, of the 13 estimated policy rule, only the point estimates for U.S. lie in the determinacy region. The hypothesis that policy was in the determinacy region would be rejected at conventional significance levels for a handful of the 12 other OECD countries in the sample, however the power of the test is likely to be quite low. For the majority of the sample (7 of 12 countries) the proportion of points that lie within the determinacy region is below 0.1. Table 1 suggests that the standard explanation, that policy was in the determinacy region in recent decades, is consistent with U.S. monetary policy, but does not fit the OECD countries very well over all.

The main factor driving the difference appears to be a stronger response to inflation in the policy rule, as the smoothing and output gap parameters for the U.S. are not outliers. Relative to the determinacy frontiers in other countries, the U.S. coefficient on inflation would be within the determinacy region for 5 of the 12 other countries.

To give a sense of where our results come from, we present the Bootstrap estimates under the null using a fixed inflation rate of 2% in Figures (5) and (6). We can see that the estimates are mostly centered around the frontier, which is expected since we impose the null. The countries for which they are not centered around the frontier, like Canada, are so because $\rho_1 + \rho_2$ is very volatile across bootstrap samples, which implies a very different frontier for each sample. It seems that when the $\rho_1 + \rho_2$ is not close to 1 in the original sample, as it is the case for Canada, we get a wider range of values. To aid visualization, we also include the Wald confidence ellipse around the point estimate in each figure.[12]

4.3 Alternative Calibration

One of the limitations of our approach, as raised in section 2.1, is that we are forced to calibrate a number of the model parameters and that the calibrated parameters can affect the determinacy region, in some cases by substantial margins. For our baseline case, we use the calibration of Coibion and Gorodnichenko (2011). However, because their model is calibrated to the U.S., some parameterizations may be inappropriate for some of the countries in this section.

In this section we consider alternate, country specific, calibrations for some of the theoretical model parameters from section 2.1. In particular, the markup parameter θ and the Calvo adjustment parameter λ affect the determinacy region significantly and appear to vary across country. We calibrate these parameters country-by-country where possible, using existing data. The remaining model parameters are unchanged from section 2.1. Table 2 presents our alternative calibration.

Table 3 presents the results under the alternative calibration. Column 1 presents the value of the policy rule coefficient on inflation that would be on the determinacy

[12] The Wald confidence interval does not constitute the basis for a formal test in our setting because it does not take into consideration the random aspect of the frontier.

Table 2: Alternate Calibration of θ and λ.

	λ^a	θ^b
Belgium	0.81	6
Canada	0.77^c	6
Denmark	0.81	7.67
Finland	0.82	5
France	0.79	7.67
Germany	0.87	6
Ireland	0.77^c	6^d
Italy	0.89	6
Japan	0.74	5
Spain	0.83	6.56
Switzerland	0.77^c	6^d
UK	0.79	7.67
US	0.60	7.67

[a] Monthly frequencies converted to average duration using method from Dhyne et al. (2005), Technical Appendix, section 5. Monthly frequencies from Table 1, Klenow and Malin (2010), unless otherwise indicated.

[b] Data from Martins et al. (1996), unless otherwise indicated.

[c] Euro area average, Table 2, Dhyne et al. (2005)

[d] Euro Area Average from Christopoulou and Vermeulen (2012).

Table 3: Determinacy Tests: Alternate Calibration

	D_{ϕ_π}	P-value	Det. %
Belgium	9.46	0.04	0.01
Canada	2.01	0.29	0.07
Denmark	12.03	0.03	0.02
Finland	2.87	0.13	0.15
France	6.58	0.15	0.00
Germany	13.63	0.04	0.00
Ireland	2.28	0.11	0.01
Italy	33.95	0.04	0.00
Japan	1.14	0.34	0.16
Spain	16.95	0.04	0.01
Switzerland	1.57	0.57	0.09
UK	25.13	0.03	0.00
US	1.53	0.83	0.59

frontier, given the other estimated coefficients and calibrated parameters.[13] Columns 2 and 3 of the table present the p-value of the test and the proportion of bootstrapped point estimates that lie within the determinacy region respectively.

Comparing Tables 3 and 1 reveals that the effect of the alternate calibration is to increase the value of ϕ_π required for determinacy, that is to shrink the region of determinacy, in all countries except for Japan and the U.S. Consequently, as in section 4.2, only the point estimate for the U.S. lies in the determinacy region.

Despite the shrinking of the determinacy region under the alternate calibration, the p-values for the test of determinacy do not change very much, and actually rise for some countries. This is because the slope of the determinacy frontier affects both the numerator of our test statistic, via the vertical distance between the estimate of ϕ_π and the determinacy frontier, and the denominator of the test statistic, via its effect on the covariance matrix. Consequently, the impact of a steeper frontier on the likelihood of rejecting the null hypotheses is ambiguous.

The smaller determinacy region does have an impact on the proportion of bootstrapped point estimates that lie within the determinacy region. Under the alternate calibration, the proportion of bootstrapped point estimates that lie within the determi-

[13]Note that the estimated policy rule parameters do not depend on the calibration and are, therefore, identical to those presented in Table 1.

Table 4: Determinacy Tests: Trend Inflation Equals Average Inflation

	D_{ϕ_π}	P-value	Det. %
Belgium	6.66	0.05	0.02
Canada	2.56	0.26	0.02
Denmark	6.39	0.03	0.05
Finland	3.79	0.08	0.28
France	4.90	0.15	0.01
Germany	3.00	0.04	0.00
Ireland	4.21	0.13	0.00
Italy	21.59	0.04	0.00
Japan	1.13	0.31	0.22
Spain	24.86	0.07	0.02
Switzerland	1.45	0.54	0.21
UK	24.29	0.03	0.01
US	2.95	0.50	0.28

nacy region fall for all countries except Finland and the U.S. This proportion lies below 0.1 for 10 of the 13 countries in the sample. Overall, the use of the alternate calibration does not alter the findings of baseline results substantively.

4.4 Trend Inflation

One issue with the results presented in the previous section is that they associated the trend inflation rate with the inflation target of the central bank. While this is reasonable, an alternate interpretation is that the trend inflation rate represents the actual average rate of inflation in the economy, as opposed to some hypothetical target to which the central bank aspires. In some of the countries in our sample inflation transitioned from a higher to a lower level throughout the course of our sample. Under this interpretation that trend inflation should correspond to the actual trend inflation rate rather than some target rate. In this section, we re-run our baseline specification using the average inflation rate in the sample as our measure of trend inflation.

Table 4 presents the results when trend inflation is set equal to average inflation. Column 1 presents the value of the policy rule coefficient on inflation that would be on the determinacy frontier, given the other estimated coefficients and calibrated parameters. Columns 2 and 3 of the table present the p-value of the test and the proportion of bootstrapped point estimates that lie within the determinacy region respectively.

The use of average inflation as the trend, rather than a 2% target, causes the value of ϕ_π required for determinacy, to rise for all countries other than Japan and the U.S. As was the case with the alternate calibration presented in section 2, the shrinking of the determinacy region does not strongly impact the p-values, but causes the proportion of bootstrapped point estimates that lie within the determinacy region fall for 10 of the 13 countries. As was the case with the alternate calibration, the use of trend rather than target inflation does not alter the findings of baseline results substantively.

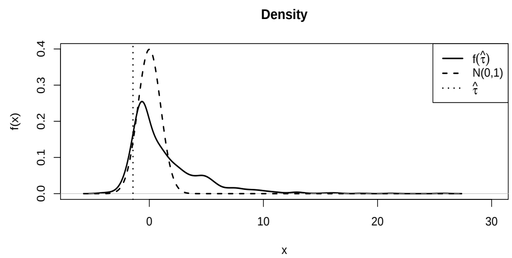
5 Conclusion

Our paper has contributed to the literature on the causes of the episode of low and stable inflation shared by many OECD countries in recent decades. Previous work in this area can be grouped into explanations relying on changes in economic, usually monetary, policy, versus explanations driven by changes in external macroeconomic conditions or shocks. Loosely termed, the good policy and the good luck schools of thought.

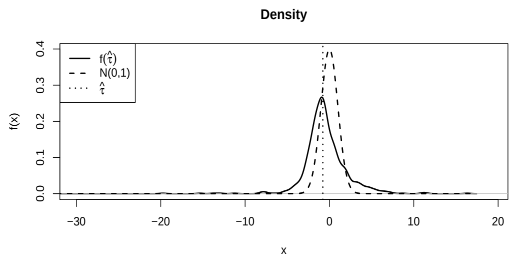
The new-Keynesian Taylor rule explanation, that improved outcomes resulted when monetary policy parameters shifted into a region that induced economic determinacy rather than indeterminacy, is a prominent version of the good policy hypothesis. Our paper investigated the question of whether this hypothesis is consistent with the experiences of a number of OECD countries, all of which have went through similar inflation episodes in recent decades.

Given data limitations, as well as the fact that our approach is only applicable under the null hypothesis that economic outcomes are determinate, we are only able to examine the second part of the explanation: that more recent monetary policy is consistent with economic determinacy. In particular, we do not attempt to investigate whether monetary policy actually changed across inflation episodes.

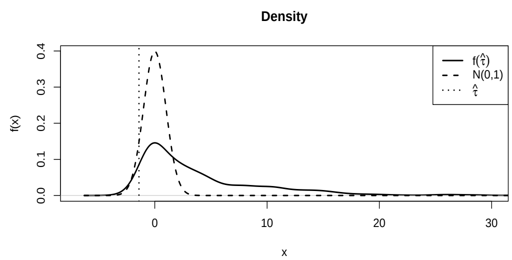
Despite the relative weakness of the test we apply, we find the evidence overall suggestive of the view that the theory is not consistent with the cross country data.



(a) Belgium



(b) Canada



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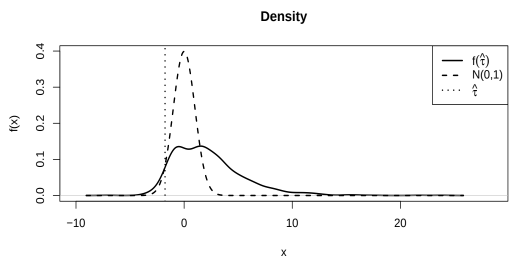
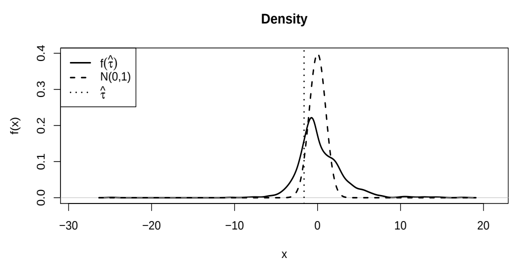


(d) Finland



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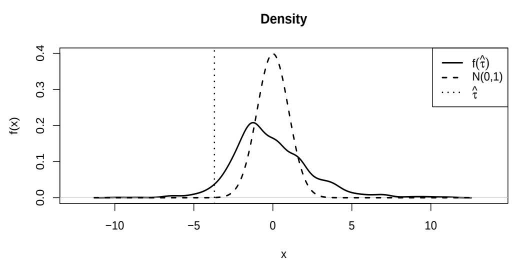

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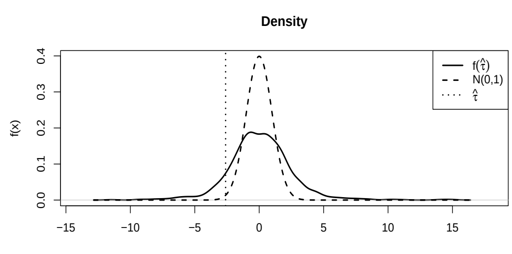
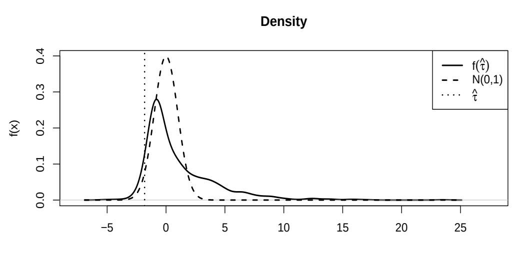
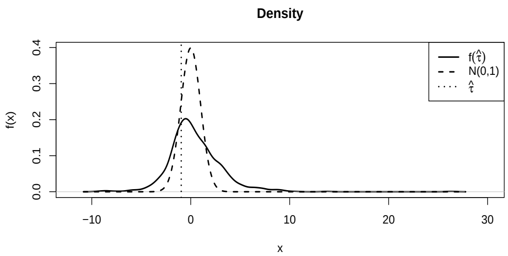


(g) Ireland

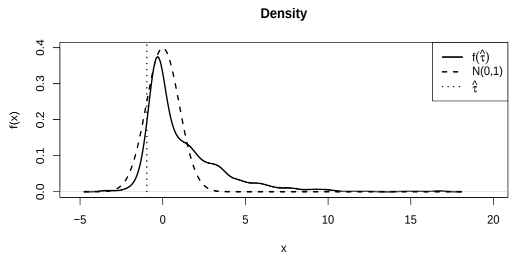


(h) Italy

Figure 3: Estimated density of $\hat{\tau}$



(a) Japan



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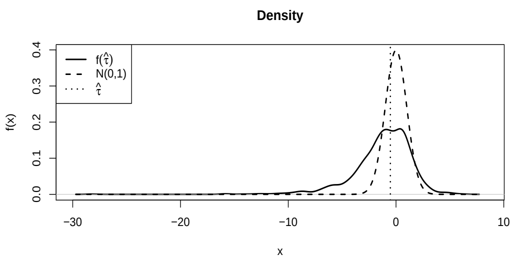

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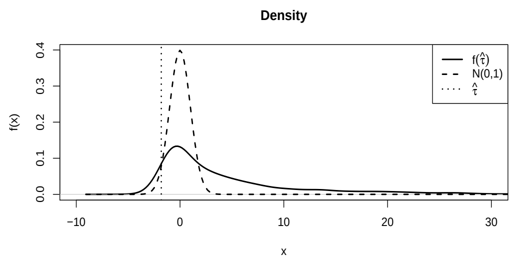

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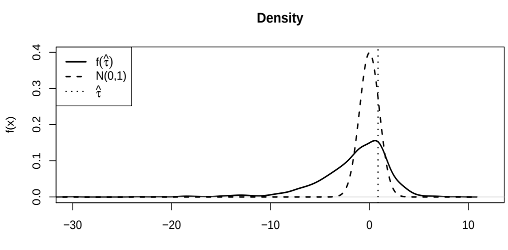
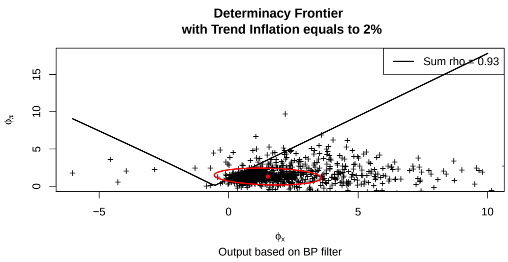
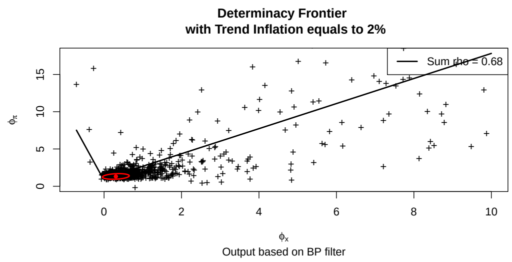


(e) US

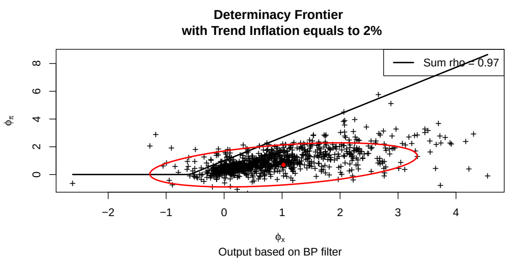
Figure 4: Estimated density of $\hat{\tau}$



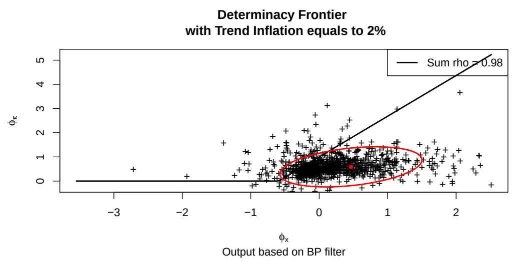
(a) Belgium



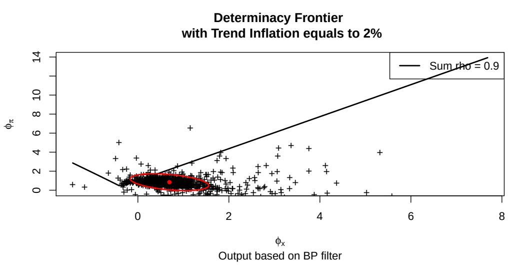
(b) Canada



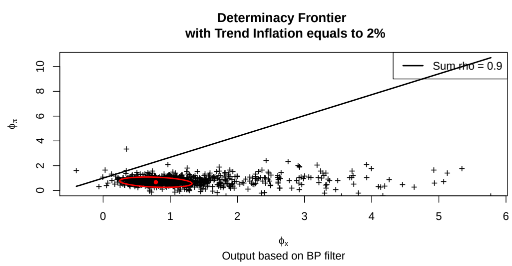
(c) Denmark



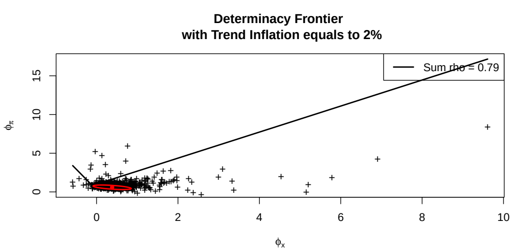
(d) Finland



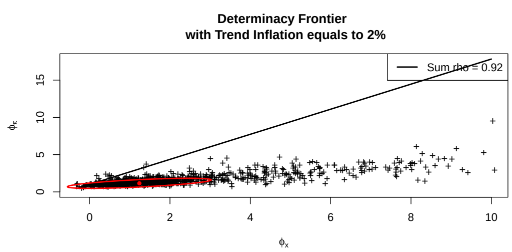
(e) France



(f) Germany

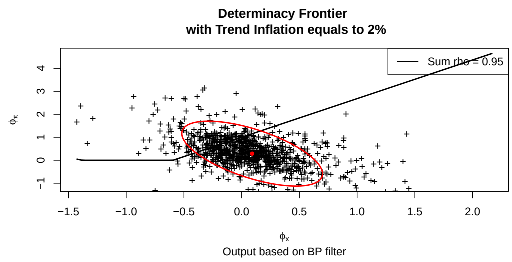


(g) Ireland

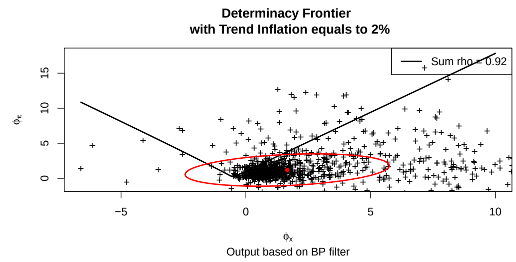


(h) Italy

Figure 5: Wald Confidence Ellipse with Bootstrap estimates under the null



(a) Japan



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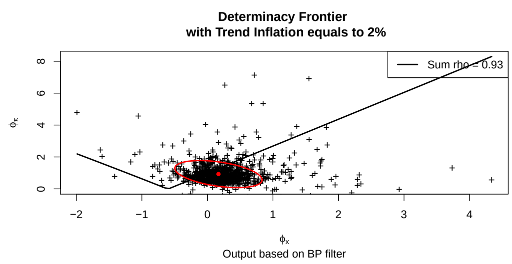

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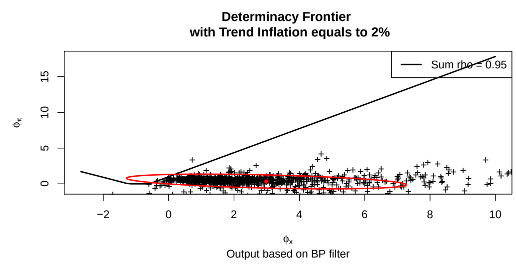

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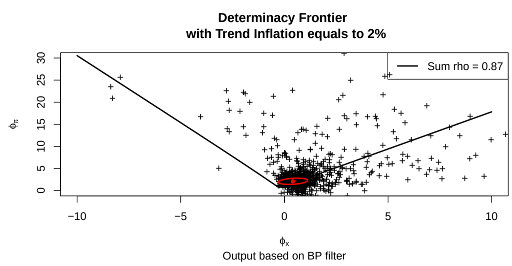

(e) US

Figure 6: Wald Confidence Ellipse with Bootstrap estimates under the null

References

D.W.K. Andrews. Heteroskedasticity and autocorrelation consistent covariance matrix estimation. *Econometrica*, 59:817–858, 1991.

D.W.K. Andrews. Inconsistency of the bootstrap when a parameter is on the boundary of the parameter space. *Econometrica*, 68:399–405, 2000.

G. Ascari and T. Ropele. Trend inflation, taylor pricniple, and indeterminacy. *Journal of Money, Credit and Banking*, 41:1557–1584, 2009.

J. Boivin and M. G. Giannoni. Has monetary policy become more effective? *Review of Economics and Statistics*, 88:445–462, 2006.

G. Calvo. Staggered prices in a utility-maximizing framework. *Journal of Monetary Economics*, 12:383–98, 1983.

R. Christopoulou and P. Vermeulen. Markups in the euro area and the us over the period 1981–2004: a comparison of 50 sectors. *Empirical Economics*, 42:53–77, 2012.

M. Ciccaralli and B. Mojon. Global inflation. *Review of Economics and Statistics*, 92: 524–535, 2010.

R. Clarida, J. Galí, and M. Gertler. Monetary policy rules in practice: Some international evidence. *European Economic Review*, 42:1033–1067, 1998.

R. Clarida, J. Galí, and M. Gertler. Monetary policy rules and macroeconomic stability: Evidence and some theory. *Quarterly Journal of Economics*, 115:147–180, 2000.

J. Cochrane. Determinacy and identification with taylor rules. *Journal of Political Economy*, 119:565–615, 2011.

T. Cogley and Sargent T. J. Evolving post-world war ii u.s. inflation dynamics. *NBER Macroeconomics Annual*, 16:331–373, 2001.

T. Cogley and Sargent T. J. Drifts and volatilities: Monetary policies and outcomes in the post wwii u.s. *Review of Economic Dynamics*, 8:262–302, 2005.

O. Coibion and Y. Gorodnichenko. Monetary policy, trend inflation and the great moderation: An alternative interpretation. NBER Working Paper 14621, 2008.

O. Coibion and Y. Gorodnichenko. Monetary policy, trend inflation and the great moderation: An alternative interpretation. *American Economic Review*, 101:341–370, 2011.

E. Dhyne, L. Álvarez, H. Le Bihan, G. Veronese, D. Dias, J. Hoffman, N. Jonker, P. Lünnemann, F. Rumler, and J. Vilmunen. Iprice setting in the euro area: Some stylized facts from individual country data. ECB Working Paper 524, 2005.

R. DiCecio and E. Nelson. The great inflation in the united states and the united kingdom: Reconciling policy decisions and data outcomes. NBER Working Papers 14895, 2009.

M. Doyle and B. Falk. Testing commitment models of monetary policy: Evidence from oecd countries. *Journal of Money, Credit and Banking*, 40:409–425, 2008.

Jean-Marie Dufour. Some impossibility theorems in econometrics, with applications to structural and dynamic models. *Econometrica*, 65:1365–1389, 1997.

A. Hornstein and A. Wolman. Trend inflation, firm-specific capital, and sticky prices. *Federal Reserve Bank of Richmomd Economic Quarterly*, 91:57–83, 2005.

A. Inoue and M. Shintani. Boostrapping gmm estimators for time series. *Journal of Econometrics*, 133:531–555, 2006.

P. Ireland. Does the time-consistency problem explain the behavior of inflation in the united states? *Journal of Monetary Economics*, 44:219–291, 1999.

M. Kiley. Is moderate-to-high inflation inherently unstable? *International Journal of Central Banking*, 3:173–201, 2007.

P. Klenow and B. Malin. Microeconomic evidence on price-setting. In B. Friedman and M. Woodford, editors, *Handbook of Monetary Economics*. Elsevier, 2010.

T. Lubik and F. Schorfheide. Testing for indeterminacy: An application to u.s. monetary policy. *American Economic Review*, 94:190–217, 2004.

J. Martins, S. Scarpetti, and D. Pilat. Mark-up pricing. market structure and the business cycle. *OECD Economic Studies*, 27:71–105, 1996.

S. Mavroeidis. Weak identification of forward-looking models in monetary economics. *Oxford Bulletin of Economics and Statistics*, 66:609–635, 2004.

T. Monacelli and L. Sala. The international dimension of inflation: evidence from disaggregated consumer price data. *Journal of Money, Credit and Banking*, 41:101–120, 2009.

H. Mumtaz and P. Surico. Evolving international inflation dynamics: world and country-specific factors. *Journal of the European Economics Association*, 10:716–734, 2012.

A. Nikolsko-Rzhevskyy. Monetary policy estimation in real time: Forward-looking taylor rules wihtout forward-looking data. *Journal of Money, Credit, and Banking*, 43:873–897, 2011.

A. Orphanides. Monetary policy rules based on real-time data. *American Economic Review*, 91:964–985, 2001.

G. Primiceri. Why inflation rose and fell: Policymakers' beliefs and us postwar stabilization policy. *Quarterly Journal of Economics*, 121:867–901, 2006.

K. Rogoff. Globalization and global disinflation. In *Monetary Policy and Uncertainity: Adapting to a Changing Economy*, pages 77–122. Federal Reserve Bank of Kansas City, 2004.

T.J. Sargent. *The Conquest of American Inflation.* Princeton University Press, Princeton, New Jersey, 1999.

E. Simms. Identification and estimation of interest rate rules in the new keynesian model. University of Michigan, Working Paper, 2008.

J. Taylor. Discretion versus policy rules in practice. *Carnegie-Rochester Conference Series on Public Policy*, 39:195–224, 1993.

J. I. Urquiza. Do estimated taylor rules suffer from weak identification? University of Houston, Working Paper, 2010.