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Time variation in an optimal asymmetric preference monetary policy model

Abstract: This paper considers a time varying parameter extension of the Ruge-Murcia's (Ruge-Murcia, F. J. 2003. "Does the Barro-Gordon Model Explain the Behavior of US Inflation? A Reexamination of the Empirical Evidence." *Journal of Monetary Economics* 50: 1375–1390; Ruge-Murcia, F. J. 2004. "The Inflation Bias When the Central Bank Targets the Natural Rate of Unemployment." *European Economic Review* 48: 91–107.) model to explore whether some of the variation in parameter estimates seen in the literature could arise from this source. A time varying value for the unemployment volatility parameter can be motivated through several means including variation in the slope of the Phillips curve or variation in the preferences of the monetary authority. We show that allowing time variation for the coefficient on the unemployment volatility parameter improves the model fit and it helps to provide an explanation of inflation bias based on asymmetric central banker preferences, which is consistent across subsamples.

Keywords: asymmetric preferences; conditional unemployment volatility; time varying parameter.

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1 Introduction

In an important contribution to understanding monetary theory, Barro and Gordon (1983) described a policy maker who is unable to make long term policy commitments. By focusing on short term results, it is possible that the policy maker pursues policies which create surprise inflation. This inflation bias result has been explored in numerous empirical studies including Ireland (1999), Ruge-Murcia (2003, 2004) and others. Although Ruge-Murcia (2003, 2004) showed that the Barro and Gordon style inflation bias is not supported by the data, he did show that an inflation bias arising from asymmetric monetary authority preferences is.

One concern with many of the empirical results, is that the parameter estimates reported in the literature show a wide range of values with some showing that the asymmetric preference bias hypothesis disappearing over certain sample periods.¹ This paper considers a time varying parameter extension of the Ruge-Murcia (2003, 2004) model to explore whether some of the variation in parameter estimates seen in the literature could arise from this source. A time varying value for the unemployment volatility parameter can be motivated through several means. It could arise through a common hypothesis that there is variation in the slope of the Phillips curve and/or it could arise because of time variation in both the preferences of the monetary planner and the political pressures she faces over time. We show that allowing time variation for the coefficient associated with the conditional unemployment volatility significantly improves the model fit.

Following Ball and Mazumder (2011), we tie the magnitude of the coefficient variation to the variation in the inflation shocks. We show that for larger coefficient variation settings the mean value of the coefficient decreases. Although this implies that the mean value of the coefficient becomes statistically insignificant, it is important to stress that it does not imply that the asymmetric monetary planner

¹ In a similar framework where central bank preferences depend on output gap volatility instead of unemployment volatility, Surico (2007) finds that output gap volatility is not important in characterizing inflation bias during the Great Moderation period of 1982:4–2002:3.

preference hypothesis suggested in Ruge-Murcia (2003) is invalid because the time varying parameter is still tied to the conditional unemployment variance. What our results do show is that time variation could account for the range of parameter estimates tied to the conditional unemployment volatility found in the literature.

The rest of the paper is organized as follows. Section 2 provides a brief description of the empirical model with asymmetric central banker preferences suggested by Ruge-Murcia (2003) and the simple extension of considering a time varying coefficient associated with the conditional unemployment volatility. Section 3 shows the estimation results and discusses their implications. Finally, Section 4 concludes.

2 The empirical model

We begin with a brief review of the asymmetric preference model suggested by Ruge-Murcia (2003, 2004). This model begins with a short run Phillips curve given by

$$u_t = u_t^n - \lambda(\pi_t - \pi_t^e) + \eta_t,$$

where u_t is observed unemployment at time t , u_t^n is the natural rate of unemployment at time t , π_t is the inflation rate at time t , π_t^e is the public's forecast of inflation at time t constructed at time $t-1$, and η_t is a supply disturbance.

The natural rate of unemployment evolves over time according to

$$\Delta u_t^n = \psi - (1 - \delta)u_{t-1}^n + \theta \Delta u_{t-1}^n + \zeta_t,$$

where ζ_t is serially uncorrelated and normally distributed with mean zero and standard deviation σ_ζ . Note that when $\delta=1$, the model imposes a unit root process for unemployment while when $\delta \neq 1$, there is no stochastic trend.

Actual inflation for the period is assumed to be simply determined as the sum of i_t , the interest rate chosen by the monetary authority and a control error, ε_t , so that

$$\pi_t = i_t + \varepsilon_t, \quad (1)$$

where ε_t is serially uncorrelated and normally distributed with mean zero and standard deviation σ_ε . Define ξ_t to be the 3×1 vector that contains the model's structural shocks at time t . It is assumed that ξ_t is serially uncorrelated, normally distributed with zero mean, and (possibly) exhibiting conditional heteroskedasticity,

$$\xi_t | I_{t-1} = \begin{bmatrix} \eta_t \\ \zeta_t \\ \varepsilon_t \end{bmatrix} | I_{t-1} \sim N(0, \Omega_t), \quad (2)$$

where Ω_t is a 3×3 positive-definite variance-covariance matrix. The conditional heteroskedasticity of ξ_t relaxes the more restrictive assumption of constant conditional second moments and captures temporary changes in the volatility of the structural shocks.

The policy maker selects i_t in an effort to minimize a loss function that penalizes variations of unemployment and inflation around target values according to

$$\left(\frac{1}{2} \right) (\pi_t - \pi_t^*)^2 + \left(\frac{\phi}{\gamma^2} \right) (\exp(\gamma(u_t - u_t^*)) - \gamma(u_t - u_t^*) - 1),$$

where $\gamma \neq 0$ and $\phi > 0$ are preference parameters, and π_t^* and u_t^* are desired rates of inflation and unemployment, respectively. As in Ireland (1999) and Ruge-Murcia (2003), it is assumed that π_t^* is a constant denoted

by π^* . The linex function characterizing unemployment allows for asymmetric preferences on unemployment by assigning different weights depending on the sign of unemployment deviations from the target.² In particular, for $\gamma > 0$ positive unemployment deviations from the target are weighted more than negative ones in the monetary authority's loss function. Also notice that the asymmetric loss function nests the symmetric (quadratic) loss function whenever γ goes to zero. Thus, the presence of asymmetric central bank preferences over unemployment can be detected by running a test on whether γ is significant.

The unemployment level targeted by the central banker is proportional to the natural rate value according to

$$u_t^* = kE_{t-1}u_t^n \quad \text{for } 0 < k \leq 1. \tag{3}$$

After some algebra involved with solving the optimization problem faced by the central banker under discretion and applying a first order Taylor series expansion to linearize the first order condition, one can arrive at the two key econometric equations given by

$$\pi_t = a + bE_{t-1}u_t + c\sigma_{u,t}^2 + e_t, \tag{4}$$

where a is a constant intercept, $b = \phi\lambda(1-k) \geq 0$, $c = \frac{\phi\lambda\gamma}{2} \geq 0$, and e_t is a reduced form disturbance and

$$\Delta u_t = \psi - (1-\delta)u_{t-1} + \theta\Delta u_{t-1} + \zeta_t + \eta_t - \lambda\varepsilon_t + \delta(\lambda\varepsilon_{t-1} - \eta_{t-1}) + \theta(\lambda\Delta\varepsilon_{t-1} - \Delta\eta_{t-1}). \tag{5}$$

Equations (4) and (5) constitute the system of equations estimated by Ruge-Murcia (2003, 2004). Stationary and nonstationary versions of the model can be investigated by placing different restrictions on δ . When $\delta = 1$, equation (5) implies that u_t is an *ARIMA*(1,1,2) process, while for $\delta < 1$, u_t is an *ARIMA*(2,0,2) process.

The coefficient associated with the conditional unemployment volatility, $c = \frac{\phi\lambda\gamma}{2}$, is a function of three deep model parameters characterizing central bank preferences (ϕ and γ) and the slope of the Phillips curve, λ . Any of these parameters may have changed over time. Indeed, the instability of the Phillips curve is a recurrent issue in the macroeconomics literature.³ Moreover, the central bank loss function is likely to change over time for several reasons ranging from different preferences associated with the alternative Fed chairmen and changes in the political pressures a Fed chairman faces as a consequence of a changing economic environment.⁴

In order to study empirically the consequences of a changing economic environment on inflation bias, we augment the model to include the possibility of c changing over time according to

$$c_t - \bar{c} = \rho_c(c_{t-1} - \bar{c}) + w_t.$$

We interpret this as a reduced form arising because of time variation in any of the three deep parameters entering in c . Although it would be preferable to formulate the time variation from the actual deep parameter specification, this is rather complicated because one would need to deal with three different time-varying processes. In addition, if there were three fundamental sources of time variation, it would be difficult to ensure identification of the additional time-varying process parameters with only two observable variables (inflation and unemployment) unless more restrictions, such as those described below, are imposed. That

² The linex function was introduced by Varian (1974) in the context of Bayesian econometric analysis. More recently, Nobay and Peel (2003) introduced it in the optimal monetary policy analysis.

³ For instance, recent empirical evidence [among others, Atkeson and Ohanian (2001), Smets and Wouters (2007), Ball and Mazumder (2011)] suggests that the slope of the Phillips curve has decreased in recent decades.

⁴ See Surico (2007, p. 317) for an excellent review of anecdotal evidence found in the literature of the changing political pressures faced by the Fed over the post-war period.

said, it is rather straight forward to see that our formulation could arise in a setting in which the model is simply augmented to include a time varying ϕ_t .⁵

Following Ball and Mazumder (2011) we tie the variance of w_t to the variance of e_t .⁶ In addition to their constraint that $\sigma_e^2 = 100 \times \sigma_w^2$, we also consider a ratio of 100,000 to show that this value roughly replicates the Ruge-Murcia (2003) estimates. In the tables below, we refer to this ratio of the variances using the notation r (i.e., $r = \sigma_e^2 / \sigma_w^2$). Linking the variance of w_t and e_t is important for two main reasons. First, we view the shocks, w_t , affecting deep parameter stability as relatively small compared to inflation shocks, e_t . Second, from an econometric perspective, since we consider two observables (unemployment and inflation) it seems reasonable to impose some additional restrictions in order to properly identify the parameters describing the three shocks of the model.

3 Estimation results

In this study we consider three sample periods. The first is the full sample period which runs from 1960:1 to 2011:2. Later we undertake a sensitivity analysis which focuses on two subsamples. The first subsample is the period 1960:1–1999:4, which is a popular data interval that has been used by numerous people, including Ireland (1999) and Ruge-Murcia (2003). The second subsample considers the so-called Great Moderation period, when most macroeconomic variables exhibit a much lower volatility than before or after the period. The analysis of the Great Moderation period is of interest because some authors, such as Surico (2007), have challenged the validity of the asymmetric central banker preference hypothesis for this period. We used the interval of 1983:1–2007:2 as our Great Moderation period.⁷

Following Ruge-Murcia (2003) and others, we first perform neglected *ARCH* tests to check whether there is conditional volatility in the unemployment series. Table 1 shows the results of these tests for the three

Table 1 LM tests for neglected *ARCH*.

Squared residuals	Sample period	No. of lags					
		1	2	3	4	5	6
Original	1960:1–2011:2	16.78*	23.84*	24.16*	24.45*	25.16*	25.72*
	1960:1–1999:4	9.65*	15.15*	15.89*	16.21*	16.66*	17.17*
	1983:1–2007:2	0.03	0.43	1.50	2.73	2.82	3.72
Standardized	1960:1–2011:2	0.00	0.17	0.75	1.59	5.06	5.75
	1960:1–1999:4	0.03	0.70	1.03	2.10	4.36	6.30

Notes: The superscript * indicates the rejection of the null hypothesis of conditional homoskedasticity at the 5% and 10% significance levels. LM tests for neglected *ARCH* in the standardized squared residual were not reported for the Great Moderation period (1983:1–2007:2) because there is no evidence of conditional heteroskedasticity in the original squared residuals for this subsample.

⁵ In particular, this occurs when ϕ_t is the only time-varying parameter and follows the same *AR*(1) process as the one describing c_t . This can be seen by looking at the first-order condition of the central bank's optimization problem [see the first order equation on Ruge-Murcia (2003, p. 1380)] and the fact that ϕ_t , being a central bank preference parameter, belongs to the central bank's information set.

⁶ Ball and Mazumder (2011) impose this type of restriction in the estimation of the slope of a backward-looking Phillips curve. Moreover, they impose that the Phillips curve slope follows a random walk process, whereas the coefficient associated with the conditional unemployment volatility is assumed to follow an *AR*(1) process in this paper.

⁷ The exact timing for the Great Moderation period is somewhat debatable. Regarding its beginning, Surico (2007) suggests the fourth quarter of 1982 whereas Smets and Wouters (2007) consider the first quarter of 1984. We choose a starting point between these two. We also choose the second quarter of 2007 as the end of the Great Moderation since after this date the rate of unemployment has shown a sharp positive increase. Currently the rate of unemployment is twice as big as it was in the second quarter of 2007 and is about the same as those seen in the early 1980's. That being said, the estimation results from the Great Moderation period are not sensitive to slight changes in its dating.

sample periods using residuals of an $ARMA(1,2)$ fit to the original unemployment series and then using the residuals of the standardized series which were created by running a first step $GARCH(1,1)$ on unemployment with an $AR(6)$ model for the mean equation. These test statistics have $\chi^2(q)$ distribution where q is the number of lags. Focusing on the full sample results, we see that these neglected $ARCH$ tests reject the null hypothesis of no conditional volatility for the original series and do not reject the null hypothesis of no conditional volatility for the standardized series.

Because the conditional $ARCH$ test results show evidence that unemployment over the full sample exhibits conditional heteroskedasticity, we now undertake an investigation of whether there might be asymmetric policy preferences. Ruge-Murcia (2003, 2004) considered two formulations for the natural rate of unemployment process which lead to two slightly different reduced form specifications. In order to save space, we focus our attention on the nonstationary case, $ARIMA(1,1,2)$.⁸ Table 2 presents estimates, for the whole sample period (1960:1–2011:2) for the Ruge-Murcia model and the time varying models for variance ratios of 100 and 100,000. There were three differences between our estimation procedure and Ruge-Murcia’s, none of which impacted the results. First, we estimated the unemployment conditional variances using the same first step $GARCH(1,1)$ model used to compute the standardized residuals described in the neglected $ARCH$ tests.⁹ Second, our maximum likelihood estimates made use of the Kalman Filter which is well suited for estimating time variation in the parameters.¹⁰ And, third we constrained b to be greater or equal to zero. This constraint is necessary so that the Ruge-Murcia model is properly nested as a special case of the time varying parameter model and thus likelihood ratio statistics can be applied.¹¹

Table 2 Sample 1960:1–2011:2 (number of observations: 206).

Coefficient	Model		
	Ruge-Murcia $0 < k \leq 1$	Time varying	
		$k=1, r=100,000$	$k=1, r=100$
a	3.0593 (0.2003)	3.0650 (0.2000)	3.2273 (0.1895)
b	0.0000 (–)		
c or \bar{c}	6.1857 (2.3425)	6.0313 (2.3381)	1.6090 (2.2067)
ρ_c		0.9769 (0.0023)	0.9773 (0.0036)
Mean log likelihood	–2.3154	–2.3124	–2.2325

Notes to Tables 2–4: Forecasting unemployment using $ARIMA(1,1,2)$. a is the intercept term, b is the coefficient of expected unemployment, c (\bar{c}) is the (mean of the time-varying) coefficient of the conditional variance of unemployment, and ρ_c is the first-order autoregressive parameter of c_t process. Standard errors are in parenthesis. The standard error of b is not reported in some cases because the lower-bound imposed by economic theory on b (i.e. b is non-negative) is binding in those cases.

⁸ The interested reader may check the working paper version of the paper, Cassou and Vázquez (2012), for a comparison of the estimation results under both the stationary and nonstationary specifications of the natural rate of unemployment process. The empirical results are robust across the two specifications and justifies our inclusion of only one here.

⁹ An advantage of estimating the conditional variance in a first step is that we use the whole sample to estimate it. This feature is important when we perform a sensitivity analysis across subsamples below. For instance, if the estimated conditional variance were estimated using data from the Great Moderation period, this estimated conditional variance is likely to be biased since it only considers data from a period which featured low volatility.

¹⁰ Of lesser note is that we also included $\frac{1}{\sqrt{2\pi}}$ in our likelihood function which Ruge-Murcia (2003) did not. So our likelihoods differ from his by this factor. This omission does not invalidate Ruge-Murcia’s likelihood ratio statistics because the missing factor in both likelihood values cancel.

¹¹ As Table 2 shows, in the sample period from 1960:1 to 2011:2, this non-negativity constraint was binding. We did estimate the model without this constraint and the estimated value was -0.06 with a standard error of 0.10 . Furthermore, the likelihood values are only different in the second decimal place. From these results, we conclude that the constraint is supported by the data.

Column 1 shows the estimation results of the Ruge-Murcia model with no time variation. As in Ruge-Murcia (2003), we cannot reject the null that $H_0: b=0$ for the two unemployment specifications considered. This result means that policy makers seem to target the natural rate of unemployment as suggested by Blinder (1998) and McCallum (1997), among others. Since we cannot reject this null, the time-varying parameter models were estimated using this constraint. Columns 2 and 3 show the estimation results for two alternative values of the variance ratio, r . For large values of r (column 2 with $r=100,000$), which implies a low variation of c_t , the estimation results are quantitatively similar to those found in Ruge-Murcia model. Thus, parameter estimates, standard errors of estimates and model fit measured by the mean log likelihood are almost identical in the two versions. In contrast, column 3 shows that a relatively small variance of the shock affecting c_t (i.e., $r=100$) is enough to improve significantly the fit of the time varying model. Interestingly, the estimated value of \bar{c} is no longer significant. However, this result should not be interpreted as evidence against the asymmetric policy maker preference suggested by Ruge-Murcia (2003) since the time varying parameter, c_t , is still tied to the conditional unemployment variance. This is an important feature to keep in mind in the discussion below. Moreover, the estimation results also show that c_t is highly persistent for any value of r . A highly persistent c_t is consistent with the idea of low frequency movements associated with both central bank preferences and the slope of the Phillips curve.

Because the time varying model nests the Ruge-Murcia model when ρ_c and the parameters characterizing the shock process of c_t (i.e., w_t) are all zero, the restrictions imposed by Ruge-Murcia's model can be tested through a standard likelihood ratio (LR) test, which is distributed as a $\chi^2(3)$.¹² The LR statistic takes the value 34.15, which imply that the restriction of a constant coefficient associated with the conditional unemployment volatility in the Ruge-Murcia model is rejected at any standard significance level.

As noted at the beginning of the section, we next undertake a sensitivity analysis by considering two alternative subsamples. First we will focus on the 1960:1–1999:4 subsample. Referring back to Table 1, note that the second row indicates the presence of conditional heteroskedasticity in the unemployment series as was documented by Ruge-Mucia (2003). Furthermore, the second row of the standardized neglected ARCH tests show that the GARCH(1,1) formulation produces a series whose residuals no longer exhibit ARCH behavior which is again consistent with Ruge-Murcia (2003). Table 3 shows the estimation results for this subsample. Column 1 shows the results from estimation of the Ruge-Mucia model. It shows results similar to those reported in Ruge-Murcia (2003) using our slightly different estimation algorithm. In particular, the estimate of b is not significantly different from zero, suggesting that the central banker indeed targeted the natural rate of unemployment, and the estimate of c is positive and significant, which supports the main hypothesis suggested in Ruge-Murcia (2003) that the US Federal Reserve cares more heavily about positive than negative unemployment deviations from the natural rate.

Table 3 Sample 1960:1–1999:4 (number of observations: 160).

Coefficient	Model		
	Ruge-Murcia $0 < k \leq 1$	$k=1, r=100,000$	$k=1, r=100$
a	2.6972 (0.6790)	3.2326 (0.2317)	3.3309 (0.2253)
b	0.1030 (0.1274)		
c or \bar{c}	7.4892 (2.8490)	8.4740 (2.2414)	4.5748 (2.3213)
ρ_c		0.9644 (0.0070)	0.9664 (0.0078)
Mean log likelihood	-2.3165	-2.3162	-2.2682

¹² The characterization of w_t requires three additional parameters describing the variance of the Choleski decomposition together with ρ_c . But because the restriction imposed by the ratio of variances, r , there are only three more parameters in the time variation model than in the Ruge-Murcia model.

Columns 2 and 3 in Table 3 show the estimation results for the time varying model for the two alternative values of the variance ratio, r . These results are similar to those found for the whole sample. In particular, for a large value of r , the results replicate the results obtained in the constant coefficient model and as r becomes smaller, the model fit improves and the mean of the time varying coefficient, \bar{c} , becomes insignificant. The LR statistic takes the value 15.46. This once again gives support for the time varying coefficient specification suggested in this paper.

Next consider the Great Moderation subsample. Referring back to Table 1, we see that row 3 of the original series tests show that we cannot reject the null hypothesis of homoskedastic errors. In other words, unemployment appears to not exhibit conditional volatility during this period.

Table 4 presents the estimation results for this subsample. Looking at Column 1, and comparing the estimation results reported here with those in the previous tables, we observe that the estimation results for the Great Moderation period stand in sharp contrast with those found for the whole sample and the 1960:1–1999:4 subsample analyzed by Ruge-Murcia (2003). In particular, the parameter b becomes significantly different from zero as if the Fed was targeting an unemployment level below the natural rate of unemployment during this period. More importantly, the conditional unemployment volatility coefficient is close to zero and non-significant. These estimation results are consistent with those reported by Surico (2007), who uses a different model and a different econometric approach. Surico (2007) interprets his results as evidence against the asymmetric central banker hypothesis during the Great Moderation period.

We have an alternative interpretation. The coefficient of the conditional unemployment volatility could be nonsignificant during the Great Moderation period because unemployment volatility, as noted in Table 1, was much smaller during this period than before it. So, even if central bank preferences were asymmetric during this period, the central banker did not need to take much action against positive deviations of unemployment from the target because these positive deviations were few and small. This means that it is hard to identify the parameter characterizing asymmetric preferences (coefficient c in Ruge-Murcia’s model) in a context where deviations of unemployment from the natural rate of unemployment are small. To study this possibility, our time varying model is quite well suited because it allows for the possibility for changes in the non-linear inflation bias function characterized by asymmetric central banker preferences.

Columns 2–4 show the estimation results for the two values of the variance ratio previously considered and a new one, $r=1$, respectively. We consider this additional r value here for two main reasons. First, the model fit significantly improves. Second, this value seems reasonable since the inflation shock volatility has substantially decreased during the Great Moderation which would be consistent with a smaller variation ratio as long as the size of shocks hitting the time varying coefficient has remained constant or has decreased much less than the size of inflation shocks.

Notice that the model fit of Ruge-Murcia’s model, which for this period does not show evidence of asymmetric central banker preferences, is almost identical to the time varying model when $r=100$. These results imply that the interpretation of Surico (2007) of symmetric preferences and the time varying coefficient model suggested in this paper are somewhat observational equivalent. However, two qualifications seem to be in order. First, the interpretation of Surico (2007) on this framework leads to a positive, significant estimate of

Table 4 Sample 1983:1–2007:2 (number of observations: 98).

Coefficient	Model			
	Ruge-Murcia $0 < k \leq 1$	Time varying		
		$k=1, r=100,000$	$k=1, r=100$	$k=1, r=1$
a	1.5983 (0.4621)	2.3844 (0.1635)	2.4109 (0.1669)	2.5862 (0.1978)
b	0.1652 (0.0842)			
c or \bar{c}	0.2978 (2.8816)	3.7232 (2.7134)	3.3644 (2.6925)	0.6390 (2.7521)
ρ_c		0.9469 (0.0125)	0.9482 (0.0133)	0.9480 (0.0155)
Mean log likelihood	-1.0275	-1.0348	-1.0215	-0.9310

coefficient b , which implies that the Fed targets an unemployment rate lower than the natural rate and this feature had been questioned by several authors [McCallum (1997) and Blinder (1998) among others]. Second, the time varying specification of asymmetric preferences has the attractive property of being consistent across different samples. This can be seen by comparing the parameter estimates for the $k=1$, $r=100$ models across Tables 1–3. Moreover, if one is ready to accept that the ratio of variance r is much smaller during the Great Moderation than before, the model fit of the time varying model improves significantly. This can be seen in the LR statistic associated with the null hypothesis of a constant value for c which takes the value 18.91.¹³

4 Conclusions

This paper considers a time varying parameter extension of the Ruge-Murcia (2003, 2004) model. A time varying coefficient associated with the conditional unemployment volatility is motivated through several means including variation in the slope of the Phillips curve and/or variation in the preferences of the monetary authority. The estimation results show that allowing time variation for the coefficient on the unemployment volatility parameter in the reduced-form of the model improves the model fit and it helps to provide an explanation of inflation bias based on asymmetric central banker preferences, which is consistent across subsamples.

The contribution of this paper can be understood as an example in which parameter instability is an important issue in the data and one which affects the economic conclusions of models.

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References

- Atkeson, A., and L. E. Ohanian. 2001. "Are Phillips Curves Useful for Forecasting Inflation?" *Federal Reserve Bank of Minneapolis Quarterly Review* 25: 2–11.
- Ball, L., and S. Mazumder. 2011. "The Evolution of Inflation Dynamics and the Great Recession." Paper prepared for the Brookings Panel on Economic Activity, March 2011.
- Barro, R., and D. Gordon. 1983. "A Positive Theory of Monetary Policy in a Natural Rate Model." *Journal of Political Economy* 91: 589–610.
- Blinder, A. S. 1998. *Central Banking in Theory and Practice*. Cambridge: The MIT Press.
- Cassou, S. P., and J. Vázquez. 2012. "Time Variation in an Optimal Asymmetric Preference Monetary Policy." The University of the Basque Country (UPV/EHU), DFAE II WP series, WP2012-08.
- Ireland, P. N. 1999. "Does the Time-Consistency Problem Explain the Behavior of Inflation in the United States?" *Journal of Monetary Economics* 44: 279–291.
- McCallum, B. T. 1997. "Crucial Issues Concerning Central Bank Independence." *Journal of Monetary Economics* 39: 99–112.
- Nobay, R. A., and D. A. Peel. 2003. "Optimal Discretionary Monetary Policy in a Model of Asymmetric Central Bank Preferences." *Economic Journal* 113: 657–665.
- Ruge-Murcia, F. J. 2003. "Does the Barro-Gordon Model Explain the Behavior of US Inflation? A Reexamination of the Empirical Evidence." *Journal of Monetary Economics* 50: 1375–1390.
- Ruge-Murcia, F. J. 2004. "The Inflation Bias when the Central Bank Targets The Natural Rate of Unemployment." *European Economic Review* 48: 91–107.

¹³ The fact that the model empirical fit improves by reducing the variance ratio r during the Great Moderation period is suggestive. That is, r may change over time. In this paper, however, we consider the variance ratio r as a constant parameter during the whole post-war period (1961–2011) and the alternative values assumed for r helps us to investigate whether the test results of the asymmetric central banker preferences are affected by the alternative values assumed for r .

- Smets, F. R., and R. Wouters. 2007. "Shocks and Frictions in US Business Cycles: A Bayesian DSGE Approach." *American Economic Review* 97: 586–606.
- Surico, P. 2007. "The Fed's Monetary Policy Rule and U.S. Inflation: the Case of Asymmetric Preferences." *Journal of Economic Dynamics and Control* 31: 3305–3324.
- Varian, H. 1974. "A Bayesian Approach to Real Estate Assessment." In *Studies in Bayesian Economics in Honour of L. J. Savage*, edited by S.E. Feinberg, A. Zellner, North-Holland, Amsterdam.