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Can Monetary Policy Cause the Uncovered Interest Parity Puzzle?

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Can Monetary Policy Cause the Uncovered Interest Parity Puzzle?[1]

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Abstract: Using a typical open macroeconomic model, we show that the UIP puzzle becomes more pronounced when the monetary policy rule is stricter against inflation. To determine the empirical validity of our model, we examine (the Taylor-rule-type) monetary policy rules and the slope coefficient in the regression of future exchange rate returns on interest rate differentials before and after the recent global financial crisis. We find that all economies that reduced the reaction of the policy interest rate to inflation in response to the crisis have positive slope coefficients in the UIP regressions after the crisis. Iceland has put greater weight on inflation in the policy rule after the crisis, and the UIP puzzle has become more severe there after the crisis, which is also consistent with our model. Moreover, economies for which we cannot find clear break evidence for the reaction to inflation in the monetary policy rule do not show a clear directional change in the slope coefficient of the UIP regression.

Key Words: Interest rate, Exchange rate, Monetary policy rule, Uncovered interest parity

JEL Classification: F31, F41, F47

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1. Introduction

Although the relationship between the exchange rate and interest rate differentials imposed by the uncovered interest rate parity (UIP) is widely used as a key assumption in theoretical models of international finance, few empirical studies have succeeded in providing supportive evidence for the UIP relation. Most studies have employed the predictive regression of future exchange rate change on the interest rate differential to examine whether the slope coefficient in the predictive regression is equal to unity, as implied by the theoretical UIP. However, the average of the estimated slope coefficients from 75 published studies is -0.88 according to the survey by Froot and Thaler (1990). Other surveys, such as Isard (1995) and Lewis (1995), report similar results for the UIP.

This drastic failure of the UIP has generated extensive studies to explain what makes the exchange rate deviate from the UIP. Fama (1984) emphasizes the role of a volatile risk premium to resolve the UIP puzzle. Chinn and Frankel (2002) estimate highly positive slope coefficients for some currencies that depreciated during the 1992 ERM (exchange rate mechanism) crisis. Lothian and Wu (2011) examine ultra-long time series data and find that severe violations of the UIP are observed only when the sample period is dominated by the 1980s. Chinn and Frankel (2002) and Lothian and Wu (2011) thus argue that the failure of the UIP can be considered as a Peso problem. Ito (1990) and Elliott and Ito (1999) report that expectations formed by traders in the currency market do not satisfy rationality; they have wishful expectations instead, suggesting that the failure of the rational expectation hypothesis is a reason for the failure of the UIP.

This study also attempts to propose an explanation for why an appealing UIP relation is not observed in the reduced-form predictive regression. That is, we consider the monetary policy rule to fight inflation as a main source for the deviation of the UIP in empirical studies. In fact, McCallum (1994) shows that when central banks adjust the interest rate gradually to resist rapid movement in the exchange rate, the negative relation between future change in the exchange rate and the interest rate differential can be observed in the reduced form regression. Since our study relates the monetary policy rule to the UIP puzzle, it is similar to McCallum (1994), but is different in the sense that our theoretical model is based on a more typical open macroeconomic model consisting of the UIP, the expectations-augmented Phillips curve relation, an open economy IS relation, and the Taylor-rule-type monetary policy rule. Since the interest rate differential and the exchange rate are simultaneously determined by the system of equations mentioned above, the relation between these two variables in the reduced-form predictive regression seems contradictory to the UIP in our model even without adding the exchange rate to the monetary policy rule.[2]

Stochastic simulations of the model are conducted to generate artificial data for exchange rates and interest rate differentials. Then, the reduced-form predictive regression of simulated exchange rate returns is run on simulated interest differentials to replicate UIP tests in empirical studies. In our simulations, we have varied values for the interest rate response to inflation in the monetary policy rule, and find that the estimated slope coefficients of artificial interest rate differentials depend on the values of the interest rate reaction coefficient in the monetary policy rule and the volatility of exchange rate shock. More specifically, as the value of the interest rate reaction to inflation in the monetary policy rule increases (i.e., as the central bank puts greater weight on inflation), the estimated slope coefficient is more likely to

[2]McCallum (1994) includes the exchange rate in the monetary policy rule, but Mark and Wu (1997) report that the reaction of the interest rate to the exchange rate is small and insignificant. In fact, our theoretical model for the simulation is closer to that in Chinn and Meredith (2004) than that in McCallum (1994).

become negative. In addition, as the volatility of the exchange rate shock rises, the estimated slope coefficient is more likely to become negative. The result can be interpreted as follows: when a temporary exchange rate market shock causes the exchange rate to depreciate, the inflation rate in the domestic country rises, which induces the central bank to raise the interest rate according to the monetary policy rule. In the next period, as the temporary shock disappears, the exchange rate appreciates, but the interest rate has already risen in the previous period. This mechanism will become more pronounced as the central bank puts greater importance on inflation and as the exchange rate shock becomes more dominant among other shocks. We ascertain this implication via not only simulations but also comparisons between the empirical results before and after the recent global financial crisis which caused a break in the monetary policy regimes of many advanced economies. Therefore, we argue that the negatively estimated slope coefficient in the reduced form regression is the consequence of indirect interaction between interest rates and exchange rates.

The role of the monetary policy rule in our model to resolve the UIP puzzle can also provide coherent explanations for seemingly unrelated findings in other studies. Bansal and Dahlquist (2000) estimate positive UIP slope coefficients using high-inflation countries. Flood and Rose (2002) report that the UIP appears to hold better during the crisis-strewn 1990s than it did before. Since the weight on inflation is likely to be low in high-inflation countries and crisis-experienced countries, our model predicts that the UIP in the reduced form regression is more likely to hold in these countries. In addition, Chinn and Meredith (2004) argue that the UIP works better with longer-maturity bonds than with short-horizon data. Since the impact of the monetary policy becomes weaker as the bond maturity increases, according to our model, it is natural to observe more supportive results for the UIP when long-horizon data are used.[3]

To present these ideas and evidence, the remainder of the paper is organized as follows: Section 2 briefly presents our model and the simulation method and results. Section 3 provides empirical evidence for implications from our model utilizing the recent global financial crisis. Concluding remarks are offered in Section 4.

2. The Model and Simulation

In order to relate the UIP puzzle to the monetary policy rule, the model employed in this study is similar to the typical open macroeconomic model in Chinn and Meredith (2004) except for long-term expected inflation and the long-term expected interest rate. The model can be described by the following five equations:

$$\hat{\imath}_t = (1 + \phi_\pi)\hat{\pi}_t + \phi_y \hat{y}_t \tag{1}$$

$$\hat{\pi}_t = \beta_\pi \hat{\pi}_{t-1} + (1 - \beta_\pi)\hat{\pi}^e_{t,t+1} + \beta_y \hat{y}_t + \beta_s \Delta(s_t - \hat{p}_t) + \nu_t \tag{2}$$

$$\hat{y}_t = \alpha_s(s_t - \hat{p}_t) + \alpha_r\left(\hat{\imath}_t - \hat{\pi}^e_{t,t+1}\right) + \alpha_y \hat{y}_{t-1} + \varepsilon_t \tag{3}$$

$$\hat{p}_t = \hat{p}_{t-1} + \hat{\pi}_t \tag{4}$$

$$s^e_{t,t+1} - s_t = \hat{\imath}_t - \eta_t \tag{5}$$

[3]In addition to the monetary policy, the volatility of the exchange rate market shock plays an important role in generating the UIP puzzle in our simulation. Since the exchange rate market shock in our model can be interpreted as the risk premium in other studies, the result implies that our model is also related with the literature that emphasizes the role of the risk premium in explaining the UIP puzzle. However, this implication is not tested in the empirical analysis because of the difficulty in quantifying the exchange rate market shock (e.g., the risk premium) with the data.

where i_t is the interest rate, π_t is the inflation rate, y_t is the output gap, s_t is the log exchange rate, p_t is the log price level, superscript e denotes the expectation operator, and ^ denotes a domestic variable relative to the same foreign variable (the same US variable). Equations (1)–(5) reflect the monetary policy rule, the expectations-augmented Phillips curve relation, an open economy IS relation, the price level identity, and the UIP relation, respectively. η_t, ν_t, and ε_t represent the exchange rate market shock, the inflation shock, and the output shock, respectively, and are assumed to be serially uncorrelated and independent of each other. The parameter values for the benchmark case are set as those in Chinn and Meredith (2004) and are summarized in Table 1.

Table 1. Benchmark Parameter Values in Simulations

Parameters	Value
ϕ_π	0.5
ϕ_y	0.5
β_π	0.6
β_y	0.25
β_s	0.1
α_s	0.1
α_r	-0.5
α_y	0.5
σ_η	9.5
σ_υ	1.2
σ_ε	1.8

Note: This table shows the values of the parameters in equations (1)–(5) used in the simulations.

Unlike McCallum (1994), the exchange rate is not included in the monetary policy rule. However, we can show the negative relation between the exchange rate returns and the interest rate differentials in the reduced form regression via the simulation of the model even without the exchange rate in Equation (1).[4] This is possible because the exchange rate, inflation, output, price, and interest rate are determined simultaneously by these five equations. When a temporary shock in the currency market (η_t) causes the exchange rate to depreciate, it also causes domestic inflation to rise. As a result, the central bank raises the interest rate according to the monetary policy rule; however, the exchange rate appreciates back as the temporary shock dies out, which is the UIP puzzle. Since inflation is provoked by an exchange rate shock, the reaction of the interest rate to inflation generates an indirect response of the interest rate to the exchange rate shock, and this reverse direction relation between the exchange rate and the interest rate creates the UIP puzzle phenomenon. According to this explanation, as a central bank puts greater importance on the stabilization of inflation (i.e., as greater values are assigned to ϕ_π by a central bank), the UIP puzzle phenomenon will be more pronounced in the reduced form regression. Hence, we check this possibility by varying values for ϕ_π in our simulations.[5]

[4]When we make the monetary policy react directly to the exchange rate, as assumed in McCallum (1994) in unreported simulations, the negative UIP puzzle relation can be replicated more easily. This result is available upon request. However, the monetary policy rule with the exchange rate in addition to inflation and the output gap might be more realistic for developing countries than for developed ones.

[5]We can consider similar exercises by varying values for ϕ_y in simulations. The simulation results for the decrease in ϕ_y with constant ϕ_π are similar to those from the increase in ϕ_π since both cases result in a relatively higher weight on inflation in the monetary policy rule, which is important for the determination of current and future macro variables. Hence, we do not report the simulation results separately for different values of ϕ_y. We focus on ϕ_π rather than ϕ_y for other reasons as well. The UIP puzzle will be more pronounced with ϕ_π than with ϕ_y because monetary policy might not react immediately to output increases unless there is substantial pressure from inflation. In fact, Schmitt-Grohe and Uribe (2007) demonstrate that the policy interest rate should respond to inflation but not to an output gap in the optimal monetary policy rule. In addition, central

Dynare is used to generate shocks and solve the rational expectation model described in equations (1)–(5). We generate two types of simulated data: 1,000-year data to determine whether the UIP puzzle is indeed a Peso problem and 25-year data to replicate more realistic reduced-form predictive regressions. The Blanchard and Kahn (1980) condition is satisfied to obtain a unique solution, and both ϕ_π and ϕ_y are restricted to be non-negative in simulations.

In Table 2, we examine how the slope coefficient in the predictive regression of the exchange rate returns on interest rate differentials changes as ϕ_π varies from 0.00001 to 1. Table 2 shows that as ϕ_π increases (i.e., as a central bank fights more aggressively against inflation), a negative relation is more likely between the future exchange rate returns and interest rate differentials in the predictive regression. Although the UIP relation does not hold perfectly, it works better with lower values of ϕ_π in the sense that high interest rate differentials tend to predict future depreciation in the exchange rate. The results are consistent with our explanation of the relation between monetary policy and the UIP puzzle.

Furthermore, the comparison of results with the 1,000-year simulated data and the 25-year simulated data shows that the UIP puzzle occurs much more frequently with small samples. However, this does not necessarily imply that the Peso problem explanation resolves the UIP puzzle fully. The UIP puzzle does not disappear completely even with 1,000-year simulated data when ϕ_π is high. For example, when ϕ_π is greater than 0.4, the negative UIP puzzle relation occurs in the simulation regardless of the sample size. The estimated slope coefficients with 25-year data are, on average, insignificant and negatively skewed

banks seem to adjust their reaction to inflation when a crisis occurs. The empirical results in Section 3 show that changes in ϕ_π are more pronounced around the recent global financial crisis.

Table 2. Reactions to Inflation in the Monetary Policy and the Slope Coefficient in the UIP Regression

$\phi_y = 0.1$			$\phi_y = 0.5$		
ϕ_π	β when T=1000	β when T=25	ϕ_π	β when T=1000	β when T=25
0.00001	0.4629 (4.0958)	0.2641 (0.5231)	0.00001	0.7485 (9.4407)	0.5737 (1.2707)
0.2	0.0175 (0.1467)	-0.174 (-0.055)	0.2	0.2371 (1.9454)	0.0574 (0.2153)
0.4	-0.1832 (-1.281)	-0.3539 (-0.2665)	0.4	-0.00065421 (0.0136)	-0.163 (-0.067)
0.6	-0.2888 (-2.0741)	-0.4403 (-0.3842)	0.6	-0.1411 (-1.0515)	-0.287 (-0.2242)
0.8	-0.3454 (-2.5676)	-0.4806 (-0.4573)	0.8	-0.2281 (-1.7349)	-0.3597 (-0.3256)
1.0	-0.374 (-2.8875)	-0.4957 (-0.5044)	1.0	-0.2824 (-2.2017)	-0.4018 (-0.3949)

Note: This table shows how the slope coefficient in the predictive regression of exchange rate returns on interest rate differentials varies as the reaction parameter (ϕ_π) to inflation in the monetary policy changes. Each cell shows the average estimated slope coefficient in the simulations. Numbers in parentheses show the average t-statistics of the slope coefficient in the simulations. The number of simulations is 1,000.

compared with those estimated with 1,000-year data. The reason for this is the small number of observations in the regression with 25-year simulated data.

Finally, the results in Table 2 suggest an explanation about UIP-related findings in previous studies. Bansal and Dahlquist (2000) report favorable evidence for UIP using high-inflation countries, and Flood and Rose (2002) report that the UIP appears to hold better for crisis periods. Since the relative weight on inflation is expected to be low in high-inflation countries and crisis-experienced countries, the results in Table 2 are consistent with these findings. In addition, studies such as Lothian and Wu (2011) observe that the UIP puzzle became severe during the 1980s, when not only the US but also other developed countries increased their response to inflation significantly (see Clarida et al. (1998)). This timing consistency could be further evidence in support of our theory relating the UIP puzzle to inflation targeting policy.

We also conduct simulation exercises to ascertain the role of the volatility of the exchange rate market shock (η_t).[6] As shown in the Appendix for policy and transition functions, a temporary positive currency market shock causes the exchange rate to depreciate and inflation to rise, which induces the central bank to raise the interest rate. As the temporary shock disappears, the exchange rate appreciates back, and the negative relation between the exchange rate changes and the interest rate differentials can be observed in the reduced form regression. Thus, the combination of the monetary policy reaction to inflation and the temporary exchange rate market shock can yield the UIP puzzle. Unlike the exchange rate market shock, however, the other shocks (i.e., the inflation shock (ν_t) and output shock (ε_t)) in our model cannot generate the negative UIP puzzle relation. For example, a temporary and positive inflation shock (or output shock) can make the interest rate rise and the exchange rate appreciate initially. However, as the temporary inflation (or output) shock disappears, the exchange rate is expected to depreciate back to normal, which is consistent with the UIP relation. Since different shocks play different roles in explaining the UIP puzzle, we have varied the relative volatility of these shocks to determine whether our explanation is correct.

Table 3 illustrates how the slope coefficient of the predictive regression responds as the relative volatilities of the three shocks change. Consistent with our explanation, as the

[6]The exchange rate market shock can be regarded as the risk premium.

Table 3.Volatilities of Shocks and the Slope Coefficient in the UIP Regression

Exchange rate market shock (σ_η)					
	$0.5 \times \sigma_\eta$	$0.75 \times \sigma_\eta$	σ_η	$1.25 \times \sigma_\eta$	$1.5 \times \sigma_\eta$
β when T=1000	0.6889 (9.2439)	0.3409 (3.2492)	-0.0795 (-0.5838)	-0.5304 (-3.4103)	-0.9974 (-5.7281)
β when T=25	0.6594 (1.427)	0.2383 (0.4498)	-0.2333 (-0.1551)	-0.7173 (-0.5986)	-1.144 (-0.9139)
Inflation shock (σ_ν)					
	$0.5 \times \sigma_\nu$	$0.75 \times \sigma_\nu$	σ_ν	$1.25 \times \sigma_\nu$	$1.5 \times \sigma_\nu$
β when T=1000	-0.6925 (-4.2511)	-0.3688 (-2.4828)	-0.0795 (-0.5838)	0.1513 (1.2968)	0.3273 (3.1065)
β when T=25	-0.8704 (-0.7149)	-0.5398 (-0.448)	-0.2333 (-0.1551)	0.0192 (0.1389)	0.2168 (0.4239)
Output shock (σ_ε)					
	$0.5 \times \sigma_\varepsilon$	$0.75 \times \sigma_\varepsilon$	σ_ε	$1.25 \times \sigma_\varepsilon$	$1.5 \times \sigma_\varepsilon$
β when T=1000	-0.4423 (-2.9261)	-0.2652 (-1.8519)	-0.0795 (-0.5838)	0.0921 (0.7682)	0.2399 (2.1357)
β when T=25	-0.6306 (-0.5264)	-0.4396 (-0.3577)	-0.2333 (-0.1551)	-0.0225 (0.0855)	0.1654 (0.3346)

Note: This table shows how the slope coefficient in the predictive regression of exchange rate returns on interest rate differentials varies as the volatility of a shock changes, keeping the volatilities of other shocks constant at the level of the benchmark case. Each cell shows the average estimated slope coefficient in the simulations. The numbers in parentheses show the average *t*-statistics of the slope coefficient in the simulations. The number of simulations is 1,000.

volatility of the exchange rate market shock (σ_η) increases while keeping the volatilities of other shocks constant (i.e., as the exchange rate market shock becomes more dominant), the negative slope coefficient is more likely, which implies the UIP puzzle.[7] As the volatility of the inflation shock (or output shock) increases, however, we tend to obtain a positive slope

[7]Even if a large value for the volatility of the exchange rate market shock (or the volatility of the risk premium) is assigned in the simulation, the UIP puzzle cannot be observed without assuming a sufficient reaction of the interest rate to inflation. This simulation result is also available upon request.

coefficient, which implies favorable evidence for the UIP. Hence, the results in Tables 2 and 3 confirm that the combination of the monetary policy reaction to inflation and the temporary exchange rate market shock can explain the UIP puzzle reported in previous empirical studies.

3. Empirical Evidence

We are utilizing the recent global financial crisis in the empirical analysis to examine whether our model relating the monetary policy rule to the UIP puzzle is consistent with the data pattern before and after the crisis. The recent global financial crisis, which was caused by the burst of the housing price bubble in the US, led to severe recessions in many countries. In response to the impact of the global financial crisis, central banks, including the Federal Reserve Bank (FRB) of the US, have implemented an aggressive and expansionary monetary policy, the so-called quantitative easing. We believe that this change in monetary policy stance is reflected as coefficient changes in the monetary policy rule for many economies. Therefore, after conducting structural break tests for the monetary policy rule, we further investigate whether the UIP puzzle has indeed been mitigated in countries that experienced structural breaks in monetary policy rules.

3.1 Data

Since we relate changes in the monetary policy rule resulting from the global financial crisis to changes in the UIP puzzle phenomenon, the Taylor rule and the reduced-form predictive regression for the UIP are estimated. Regarding the estimation of the Taylor rule, data for the policy interest rate, inflation rate, and output gap are required. Policy rates on the last day of each month are taken from Datastream for the policy interest rate, and consumer price index (CPI) series are obtained from International Financial Statistics (IFS). Either industrial production series or unemployment rates are used for the construction of the output gap; these variables are obtained from Datastream. If industrial production or the unemployment rate is not available from Datastream, we refer to IFS for these variables. When we use quarterly data in the analysis, real GDP is utilized in the construction of the output gap, and the data are obtained from IFS.[8] The inflation rate is computed as the percentage change of CPI over the corresponding period of the previous year, and the output gap is calculated by rolling Hodrick–Prescott filtering of industrial production or real GDP data. Unemployment rates are used as a proxy for the output gap for economies that do not have industrial production series. The whole sample period is from January 2000 to December 2012 unless otherwise noted. We focus on analyzing recent data because Bae et al. (2012) report that the monetary policy regime has varied significantly over time. Thus, changes in the monetary policy rule for other reasons could contaminate the coefficient change in the monetary policy rule due to the recent global financial crisis, as the sample period includes periods far away from the crisis.

In order to estimate the UIP relation in the predictive regression, daily interest rates and exchange rates are obtained from Datastream. We utilize daily data for the estimation of the UIP relation in order to have a sufficient sample size for precise estimation. We analyze the Taylor rule and the UIP relation for 12 developed economies. According to the International Monetary Fund's *Annual Report on Exchange Arrangements and Exchange Restrictions 2012* and Roger (2010), Australia, Canada, Iceland, New Zealand, Norway, Sweden, and the UK are classified as inflation-targeting economies, whereas the rest of the

[8]Quarterly data are used in the estimations for Australia and New Zealand because of the availability of CPI data.

economies (EU, Hong Kong, Japan, Singapore, Switzerland, and the US) are non-inflation targeting economies.

3.2 Structural Breaks in the Monetary Policy Rule (Taylor Rule)

The commonly used Taylor rule, which originated from Taylor (1993), can be written as follows:

$$i_t^* = (r^* + \pi^*) + \varphi_\pi(\pi_t - \pi^*) + \varphi_y \tilde{y}_t \tag{6}$$

where i_t^* is the annualized policy rate, r^* is the steady-state real interest rate, π^* is the target inflation rate, π_t is the actual annualized inflation rate, and $\tilde{y}_t$ is the output gap. This equation shows that the central bank's policy interest rate reacts to the inflation rate and output gap. When estimating Equation (6) in empirical studies, lagged interest rates are often added to capture the interest rate smoothing behavior in practice. Hence, we estimate the following equation to check whether the monetary policy underwent structural breaks around the global financial crisis:

$$i_t = \gamma_0 + \gamma_1 \pi_t + \gamma_2 \tilde{y}_t + \rho i_{t-1} \tag{7}$$

where i_t is the annualized short-term nominal interest rate and ρ is the interest rate smoothing parameter. By comparing Equations (6)–(7), we can see that $\gamma_0 = (1-\rho)(r^* + (1-\varphi_\pi)\pi^*)$, $\gamma_1 = (1-\rho)\varphi_\pi$, and $\gamma_2 = (1-\rho)\varphi_y$. Hence, we can recover the long-term response of the interest rate to inflation and the output gap, which are φ_π and φ_y, respectively, using estimated parameters.

We conduct the Quandt–Andrews (QA) structural break test[9] for Equation (7) to determine whether monetary policies in developed economies underwent a structural break in response to the global financial crisis. When implementing QA tests, we assume that no break falls within the first 15% or the last 15% of the sample period to leave enough observations before and after an estimated breakpoint. The results are presented in Table 4. The fourth column of Table 4 shows the QA test results under the null hypothesis that γ_1 is constant. Under this null hypothesis, structural breaks are detected for Australia, Canada, the EU, Iceland, New Zealand, Norway, Sweden, the UK and the US. Furthermore, except the EU, estimated breakpoints fall between September 2007 and October 2008, which strongly suggests that these detected breaks in the monetary policy rule should be related with the recent crisis. The estimated breakpoint for the EU is December 2005, which does not appear to be related with the global financial crisis. Interestingly, all economies except the EU and the US where a break for γ_1 is detected, are inflation-targeting economies. This result implies that monetary policies for these economies were strict with inflation before the crisis but became lenient with inflation after the crisis.

In order to check this implication, we examine whether the estimated long-term responses of the target interest rate to inflation (φ_π) have been reduced for economies for which breaks for γ_1 have been detected. Table 5 compares the estimated φ_π before and after the estimated break timings. The absolute values for φ_π during the post-break periods are lower than those during the pre-break periods for all economies examined except Iceland.

[9]The distribution of test statistics and the approximate asymptotic p-values are provided by Andrews (1993) and Hansen (1997), respectively.

Table 4. Structural Break Test for the Monetary Policy Rule

	Economy	Output Gap Proxy	H_0: γ_1 is constant
Inflation targeting economy	Australia	Real GDP	2008.10* (fourth quarter of 2008)
	Canada	Unemployment rate	2007.12***
	Iceland	Unemployment rate	2008.10**[10]
	New Zealand	Real GDP	2008.10*** (fourth quarter of 2008)
	Norway	Industrial Production	2008.10***
		Unemployment rate	2008.10***
	Sweden	Unemployment rate	2008.10**
	UK	Industrial Production	2008.09***
		Unemployment rate	2008.09***
Non-inflation targeting economy	EU	Industrial Production	2005.12**
	Hong Kong	Unemployment rate	No rejection
	Japan	Industrial Production	No rejection
		Unemployment rate	No rejection
	Singapore	Industrial Production	No rejection
	Switzerland	Unemployment rate	No rejection
	US	Industrial Production	2007.09***
		Unemployment rate	2007.09***

Note: This table shows the results of the Quandt–Andrews (QA) structural break tests for $i_t = \gamma_0 + \gamma_1 \pi_t + \gamma_2 \tilde{y}_t + \rho i_{t-1}$, where i_t is the annualized short-term nominal interest rate and ρ is the interest rate smoothing parameter. The output gap is constructed by the use of detrended industrial production, detrended real GDP, or the unemployment rate. "No rejection" means that the null hypothesis is not rejected at the 10% level. When the null hypothesis is rejected, the estimated breakpoint is written with the significance level. "*", "**", and "***" denote that the null hypothesis is rejected at the 10%, 5%, and 1% levels, respectively.

Although negative values for φ_π are estimated for Norway and the UK after breaks, they are insignificant, which implies that φ_π has been lowered to a level around zero. Iceland is an exceptional economy in the sense that φ_π has been raised after the break, which implies that policymakers in Iceland have put greater importance on inflation after the crisis. The

[10]The estimated break date is detected from the test using observations after July 2004.

Table 5. Estimated Long-run Response of the Policy Interest Rate to Inflation (φ_π)

Economy	Output Gap Proxy	Estimated Breakpoint	Jan. 2000–Breakpoint	Breakpoint–Dec. 2012
Australia	Real GDP	2008.10 (fourth quarter of 2008)	0.8446***	0.6773**
Canada	Unemployment rate	2007.12	0.9086**	0.3376*
Iceland	Unemployment rate	2008.10	0.4235***	0.8759***
New Zealand	Real GDP	2008.10 (fourth quarter of 2008)	2.9225	0.0058
Norway	Industrial Production	2008.10	1.4171	-0.1368
	Unemployment rate	2008.10	0.4768	-0.0825
Sweden	Unemployment rate	2008.10	1.0031***	0.3149***
UK	Industrial Production	2008.09	1.2133**	-0.0304
	Unemployment rate	2008.09	0.9700	0.1044
EU	Industrial Production	2005.12	$\hat{\rho} > 1$***	0.1354
US	Industrial Production	2007.09	$\hat{\rho} > 1$***	0.07972
	Unemployment rate	2007.09	0.55015*	0.15819

Note: This table compares the estimated long-run responses of the policy interest rate to inflation in the monetary policy rule before and after the breakpoints. "*", "**", and "***" denote 10%, 5%, and 1% significance levels, respectively.

estimation of φ_π often has a problem for the EU and the US. The smoothing parameter for the first sub-sample is estimated to be larger than unity for the EU and the US with industrial production being used to construct the output gap.[11] However, φ_π is lowered after the break

[11]Rotemberg and Woodford (1999) show that the coefficient on the lagged interest rate exceeding unity makes economy stay on the saddle-path rather than on an explosive path since there is an exponential increase in the

in the US when the unemployment rate is used as a proxy for the output gap.

In summary, the results in Tables 4 and 5 suggest that the monetary policy rule has changed around the crisis period in many advanced economies. These include Australia, Canada, Iceland, New Zealand, Norway, Sweden, the UK and the US (where the crisis originated). This sub-section has also shown that these countries (except Iceland) have reduced weights on inflation in determining the target interest rates in response to the recent crisis.

3.3 UIP Relation

We examine whether the UIP relation underwent a break around the crisis period similarly to the monetary policy rule, as predicted by our model.[12] According to the UIP, interest rate differentials have predictive power for future exchange rate returns. Hence, the following commonly used predictive regression between exchange rate returns and interest rate differentials is run for each economy:

$$s_{t,t+k} - s_t = \alpha + \beta(i_t - i_t^*) + \varepsilon_{t,t+k} \tag{8}$$

where s_t denotes the exchange rate, i_t denotes the domestic economy's interest rate, i_t^* denotes the foreign economy's interest rate, and $\varepsilon_{t,t+k}$ represents the mixture of the risk premium and forecasting errors. When estimating the slope coefficient, β, in the regression,

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interest rate subsequent to a rise in inflation, which indicates a considerable inflation-targeting policy.

[12]We use heteroskedasticity-autocorrelation-consistent (HAC) standard errors when testing the "no break" null hypothesis because of possible heteroskedasticity and autocorrelation in the error terms for the UIP regression.

daily frequency data are used to have sufficient observations. The forecast horizon, k, is set as 22 to represent the number of business days in one month. We employ the Newey and West (1987) standard errors to solve possible problems arising from serial correlations and heteroskedasticities in $\varepsilon_{t,t+k}$. β should be equal to unity under the risk-neutral and rational expectation hypothesis. However, we focus on the sign of β because a positive β implies that interest rate differentials can predict at least the direction of exchange rate changes based on the UIP and because previous studies consider the negative β as the UIP puzzle.

Table 6 shows the estimation results of the slope coefficient in Equation (8) for seven inflation-targeting economies for which breaks for γ_1 have been detected around the crisis. The second column of Table 6 shows the estimated β and the corresponding *t*-statistics based on the Newey–West standard errors using the full-sample observations (January 2000 to December 2012). Only two of the seven economies (Australia and Sweden) have negative slope coefficients, whereas the other economies have positive slope coefficients. None of them has a significant coefficient regardless of the estimated sign. These results may be interpreted as indicating that the UIP puzzle was mitigated during the 2000s. However, when the QA test is conducted for Equation (8), strikingly different results emerge. The null hypothesis for constant α and β can be rejected at the 1% level for all economies. Hence, all economies can be interpreted as having experienced a structural break in Equation (8) during our sample period; the estimated breakpoints are listed in the third column of Table 6. When regression equation (8) is run with observations before the estimated breakpoint, β is estimated as negative for all economies with the exception of Iceland, and three of them (Australia, New Zealand, and Sweden) are significantly negative. This finding is reported as the UIP puzzle in previous studies. The fifth column of Table 6 presents regression results

Table 6. Structural Break Test for the UIP Relation among Inflation Targeting Economies

Economy	Full sample	Estimated breakpoint in the UIP regression	Pre-break $\hat{\beta}$	Post-break $\hat{\beta}$	Estimated breakpoint in the monetary policy rule	Pre-break $\hat{\beta}$	Post-break $\hat{\beta}$
Australia	-0.1865 (-0.0774)	2008-06-30***	-5.093 (-2.4409)	38.03491 (3.9452)	2008-10-01	0.0774 (0.0242)	21.2171 (3.0589)
Canada	1.7431 (0.7601)	2007-02-28***	-3.5012 (-1.5775)	21.1451 (3.859)	2007-12-03	-0.3509 (-0.1562)	29.4341 (2.3362)
Iceland	2.1466 (1.1816)	2008-11-05***	7.818 (2.7749)	-1.4879 (-0.7389)	2008-10-01	6.0231 (2.2518)	-0.0793 (-0.0388)
New Zealand	1.1636 (0.4235)	2005-02-21***	-7.9456 (-2.8454)	14.5992 (4.0423)	2008-10-01	-1.0554 (-0.3906)	18.0117 (1.8107)
Norway	0.5562 (0.3785)	2008-07-04***	-1.3301 (-1.0949)	25.66 (3.5346)	2008-10-01	0.0283 (0.0199)	14.3455 (2.0966)
Sweden	-0.0192 (-0.0112)	2008-06-27***	-3.2582 (-2.2729)	22.6924 (3.2013)	2008-10-01	-1.1426 (-0.6458)	15.0969 (2.1337)
UK	1.0447 (0.5208)	2008-07-11***	-2.0943 (-1.1791)	30.9939 (4.3735)	2008-09-01	-0.8189 (-0.4127)	36.6842 (3.6939)

Note: "Full sample" shows the estimated slope coefficient in $s_{t,t+k} - s_t = \alpha + \beta(i_t - i_t^*) + \varepsilon_{t,t+k}$ using the full sample observations. s_t denotes the exchange rate, i_t denotes the domestic economy's interest rate, i_t^* denotes the foreign economy's interest rate, and $\varepsilon_{t,t+k}$ represents the mixture of the risk premium and forecasting errors. The Quandt–Andrews (QA) structural break tests are conducted for the equation. "*", "**", and "***" denote the 10%, 5%, and 1% significance levels, respectively. "Pre-break $\hat{\beta}$" shows the estimated slope coefficient using observations before an estimated breakpoint. "Post-break $\hat{\beta}$" shows the estimated slope coefficient using observations after an estimated breakpoint. The numbers in parentheses show the t-statistics of the slope coefficient based on the Newey–West standard errors. "Estimated breakpoint in the monetary policy rule" shows the estimated breakpoints from the monetary policy rule in Table 4.

estimated with observations after the estimated breakpoints. The estimated coefficients are significantly positive in all economies but Iceland. Iceland is the only country that increased its response to inflation in the Taylor rule after the crisis in Table 5, which implies a more serious UIP puzzle. Although the monetary policy rule in Iceland behaved in opposite directions before and after the crisis, the results for Iceland are also consistent with our model.

Large absolute values for the estimated β are obtained during the post-break periods when interest rate differentials have become substantially less volatile and exchange rate changes have become more volatile. The results in the fourth and fifth columns of Table 6 strongly suggest that the UIP puzzle was not mitigated at all before the break, but disappears after the break when the reaction to inflation in the monetary policy rule is reduced.

In order to further verify the link between the monetary policy rule and the UIP puzzle, each economy's sample period is divided into two sub-sample periods according to the estimated breakpoints in the monetary policy rule of Table 4. The results and the estimated breakpoints are shown in the sixth through eighth columns of Table 6. Interestingly, the estimated breakpoints in the monetary policy rule and in the UIP regression lie within a one-year period for all economies except New Zealand. Similar to the results in the fourth and fifth column, moreover, four of the six economies (excluding Iceland) have a negative β during the pre-break period, whereas all of them have a significantly positive β during the post-break period. The coincidence of the estimated breakpoints from the two independent regressions and the changes in the sign of the estimated β before and after the break suggest that the detected structural breaks in the UIP regression coefficients are related with monetary policy rule changes, as predicted by our model.

We also examine the predictive regression for other economies for which we cannot detect a break for γ_1 around the crisis in the monetary policy rule. Since the US shows a structural break in the monetary policy rule, the UIP regression for other economies might show evidence of structural breaks because the exchange rate is defined as the amount of domestic currency needed to purchase one US dollar. Consistent with this guess, these economies also show evidence of structural breaks, as reported in Table 7, even if the monetary policy rule in these economies has not changed. However, the slope coefficient changes its sign only for the EU and Japan. For other economies, the slope coefficient continues to be negative after the break, probably because the monetary policy rule to inflation for these economies has not been changed.

Since these economies do not seem to change the monetary policy rule in response to the global financial crisis according to the results in Table 4, we also run regressions of Equation (8) for sub-samples using September 2008 as the breakpoint. The selected breakpoint is the time when Lehman Brothers became defunct. As shown in Table 7, we are not able to find such drastic changes in the slope coefficients for three economies (the EU, Hong Kong, and Singapore) before and after September 2008. All five economies show negatively estimated slope coefficients before September 2008, and three of them continue to have negative slope coefficients even after the breakpoint. Although Japan and Switzerland have positive slope coefficients after September 2008, based on the results in Table 4, this change in the sign of the slope coefficient might be due to the change in US monetary policy rather than a change in the policy of Japan or Switzerland.

In summary, the results in Tables 6 and 7 suggest that economies for which we can find structural breaks in the monetary policy rule also experience breaks in the UIP predictive regression. The directions of the changes in slope coefficients are consistent with the prediction of our model. The estimated breakpoints in the predictive regression are quite close to those in the monetary policy rule. However, we cannot find such a directional change in the UIP regression slope coefficients for most economies in which a break in the monetary policy rule has not been detected.

Table 7. Structural Break Test for the UIP Relation among Non-inflation Targeting Economies

Economy	Full sample	Estimated breakpoint in the UIP regression	Pre-break $\hat{\beta}$	Post-break $\hat{\beta}$	Pre-2008.09 $\hat{\beta}$	Post-2008.09 $\hat{\beta}$
EU	0.2749 (0.2798)	2005-11-08***	-8.0098 (-2.993)	1.7553 (1.8772)	-1.181 (-0.5536)	-0.2301 (-0.0066)
HK	-0.5119 (-2.8291)	2010-04-06***	-0.5045 (-2.7301)	-9.3224 (-2.6305)	-0.5497 (-2.4335)	-0.7846 (-2.9032)
Japan	-1.7736 (-1.7841)	2008-08-14***	-2.0304 (-1.4916)	50.7603 (6.1352)	-2.5216 (-1.822)	60.5558 (4.2136)
Singapore	-1.677 (-2.0297)	2008-07-04***	-1.8643 (-2.1369)	-21.8715 (-4.1411)	-1.5206 (-1.6947)	-7.2199 (-0.4511)
Switzerland	-1.2001 (-0.6566)	2010-05-18***	-0.3261 (-0.1658)	-178.4274 (-3.9358)	-1.4631 (-0.6762)	3.8344 (0.1682)

Note: "Full sample" shows the estimated slope coefficient in $s_{t,t+k} - s_t = \alpha + \beta(i_t - i_t^*) + \varepsilon_{t,t+k}$ using the full sample observations. s_t denotes the exchange rate, i_t denotes the domestic economy's interest rate, i_t^* denotes the foreign economy's interest rate, and $\varepsilon_{t,t+k}$ represents the mixture of the risk premium and forecasting errors. The Quandt–Andrews (QA) structural break tests are conducted for the equation. "Pre-break $\hat{\beta}$" shows the estimated slope coefficient using observations before an estimated breakpoint. "Post-break $\hat{\beta}$" shows the estimated slope coefficient using observations after an estimated breakpoint. "Pre-2008.09 $\hat{\beta}$" and "post-2008.09 $\hat{\beta}$" represent the estimated slope coefficient before and after September 2008. The numbers in parentheses show the *t*-statistics of the slope coefficient based on the Newey–West standard errors.

4. Conclusion

In this study, we considered the UIP puzzle in terms of the reaction of the policy interest rate to inflation in the monetary policy rule. Our model states that as the central bank puts a greater weight on inflation, we are more likely to observe a negative slope coefficient in the regression of exchange rate returns on interest rate differentials. To evaluate our model, we compared monetary policy rules and the UIP predictive regressions for 12 advanced economies before and after the recent global financial crisis that might have caused central banks to change their monetary policy rules to overcome the impact of the crisis. We found that all inflation-targeting economies except Iceland reduced the reaction of the policy interest rate to inflation in response to the crisis and have positive slope coefficients in the UIP regressions after the crisis. Iceland put greater weight on inflation in the policy rule after the crisis, and the UIP puzzle has become more severe there, which is also consistent with our model. In contrast to inflation-targeting economies, however, among non-inflation targeting economies, we hardly found break evidence for the reaction to inflation in the monetary policy rule nor for the sign changes in the slope coefficient of the UIP regression. These results would seem to warrant a reconsideration of the role of monetary policy in understanding the UIP puzzle.

Finally, although the UIP relation holds in our simulation, the slope coefficient in the UIP regression could be negative depending on the monetary policy rule and the exchange rate market shock. Reduced-form predictive regression seems to have a limit to uncover the UIP relation when the interest rate and exchange rate are simultaneously determined along with other macro variables. Hence, the results in this study also imply that a structural estimation based on a true macroeconomic model should be considered to reveal the UIP relation; this is beyond the scope of this study, and we leave it to future research.

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Appendix. Estimated Policy and Transition Functions

We report the policy and transition functions for the endogenous variables in Section 2, which are generated by Dynare in the state-space form. In the state-space representation, $\hat{\tilde{\imath}}_t$ is the control variable and the state vector is $[\hat{\tilde{y}}_t \;\; \hat{\tilde{p}}_t \;\; \tilde{s}_t \;\; \hat{\tilde{\pi}}_t]'$. The vector of innovations is a 3×1 vector consisting of the exchange rate market shock (η_t), inflation shock (ν_t), and output shock (ε_t). By substituting the state transition equation into the policy function, the full system can be written as

$$\tilde{Y}_t = \begin{bmatrix} 0.36 & -0.04 & 0.04 & -0.26 \\ 0.10 & 1.10 & -0.10 & 0.63 \\ -0.44 & 0.95 & 0.05 & -0.30 \\ 0.10 & 0.10 & -0.10 & 0.63 \\ 0.32 & 0.13 & -0.13 & 0.81 \end{bmatrix} \tilde{S}_{t-1} + \begin{bmatrix} 0.01 & -0.44 & 0.72 \\ 0.07 & 1.05 & 0.19 \\ 0.98 & -0.49 & -0.88 \\ 0.07 & 1.05 & 0.19 \\ 0.12 & 1.35 & 0.65 \end{bmatrix} \begin{bmatrix} \eta_t \\ \nu_t \\ \varepsilon_t \end{bmatrix}$$

Note: $\hat{\tilde{y}}_t$, $\hat{\tilde{p}}_t$, $\tilde{s}_t$, $\hat{\tilde{\pi}}_t$, and $\hat{\tilde{\imath}}_t$ denote the log deviations of the corresponding variables in Section 2 from the steady state. We define $\tilde{S}_t = [\hat{\tilde{y}}_t \;\; \hat{\tilde{p}}_t \;\; \tilde{s}_t \;\; \hat{\tilde{\pi}}_t]'$ and $\tilde{Y}_t = [\tilde{S}_t{}' \;\; \hat{\tilde{\imath}}_t]'$.