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Can endogenous monetary policy explain the deviations from UIP?*

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Abstract

The co-movements of nominal exchange rates and short-term interest rates as the economy is hit by shocks is a potential source of *ex post* deviations from uncovered interest rate parity. This paper investigates whether an established model of endogenous monetary policy in an open economy is capable of explaining the exchange rate risk premium puzzle. Time series on interest differentials and exchange rate changes are generated from the Svensson (2000) model. Uncovered interest rate parity is tested on the simulated data and the β -coefficients are investigated. For most realistic choices of parameter values, the β -coefficients are positive but much smaller than the unity value expected from UIP. It is however also possible to obtain large, negative β -coefficients if the central bank is engaged in interest rate smoothing.

Keywords: Monetary policy, Uncovered Interest Parity, Exchange Rate Risk Premium.

JEL classifications: E52, F31, F41.

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1 Introduction

Uncovered interest parity (UIP) implies that the coefficient β from regressing exchange rate changes on lagged interest differentials equals $+1$. Numerous empirical studies have reported β -coefficients that are significantly negative, and large. Several surveys point out -3 to -4 as a typical result, see Engel (1995), or McCallum (1994). Traditional explanations of the deviations from UIP focus on time varying risk premia and/or systematic forecast errors. It is probably fair to say that neither approach has been successful in explaining the empirical failure of UIP. This paper explores a third possibility, namely that the observed negative relationship between short-term interest rates and *ex post* exchange rate changes is a consequence of the response of monetary policy to shocks.

Tentative evidence suggests that the typical finding of a negative β -coefficient is confined to short-term interest rates. The few available studies of UIP for long-term interest rates report β -coefficients that are positive and often insignificantly smaller than unity (Alexius, 2001, Meredith and Chinn, 1998). Short-term interest rates differ from other financial assets in that they constitute the main monetary policy instrument in most industrialized countries with flexible exchange rates. Both interest rates and exchange rates are endogenous variables in open economy macro models with endogenous monetary policy. McCallum (1994), Meredith and Chinn (1998), and Alexius (2000) provide examples of theoretical models where the co-movements of short-term interest rates and exchange rate changes can produce a negative relationship between short-term interest rates and *ex post* exchange rate changes. However, if this was a general result, the mechanism would be

present in standard models of endogenous monetary policy and not just in special cases

In this paper, the Svensson (2000) model of monetary policy in an open economy is used to generate artificial time series on exchange rate changes and interest differentials. The standard UIP test is applied to the resulting data sets and the β -coefficients are collected. The model parameters are then varied in order to (i) investigate what ranges of β -coefficients that emerges for realistic parameter values and (ii) identify the conditions under which large, negative β -coefficients can be obtained from the model.

The idea that the negative co-movements of exchange rates changes and interest rate differentials could be a consequence of the monetary policy response to shocks is originally due to McCallum (1994). He uses a two equation framework, consisting of a UIP relationship and a monetary policy reaction function, to illustrate his point. Meredith and Chinn (1998) incorporate this mechanism into an open economy macro model. They calibrate their model and show that it is capable of generating negative β -coefficients in standard UIP tests. The results in McCallum (1994) and Meredith and Chinn (1998) hinge crucially on existence of substantial shocks to the exchange rate risk premium, which is an exogenous shock in these models. Attempts to model endogenous exchange rate risk premia have however failed to generate a risk premium of the required magnitude. Note that whether the monetary policy response to shocks can explain the observed *ex post* deviations from UIP is a separate issue that can be discussed in isolation from the magnitude of the risk premium shocks. The shocks that monetary policy responds to do not even have to be shocks to the exchange rate risk premium. For instance,

Alexius (2000) finds a negative relationship between interest differentials and ex post exchange rate changes as the variables move in response to *demand* shocks.

This paper contributes to the literature in several ways. If it is the endogenous response of monetary policy to shocks that generates negative β -coefficients in UIP tests, this mechanism should be present not only in models constructed specifically for the purpose of explaining the exchange rate risk premium puzzle, but in open economy macro models in general. The Svensson (2000) model is an open economy version of an established framework for analyzing monetary policy (Rudebusch and Svensson, 1998, Svensson, 1999a and 1999b). Hence, we know that the model is not specially designed to produced the desired negative relationship between exchange rate changes and interest differentials. Second, while Meredith and Chinn (1998) calibrate their model for one particular set of parameter values, the parameters are varied systematically in this paper. The conditions under which the model provides an explanation for the exchange rate risk premium puzzle can then be delineated. In particular, Meredith and Chinn (1998) assign a very high variance to the risk premium shocks, which is a well-known way of obtaining negative β -coefficients in UIP tests (Fama, 1984). Third, while the models of Meredith and Chinn (1998) and Alexius (2001) are postulated *ad hoc*, Svensson (2000) derives the building blocks of his model from microeconomic foundations. Finally, monetary policy in the Svensson (2000) model does not merely follow a rule of thumb but is conducted in a forward looking manner, utilizing all available information to minimize an intertemporal loss function.

The paper is organized as follows. First, the exchange rate risk premium

puzzle is briefly discussed (Section 2) and the Svensson (2000) open economy model is presented (Section 3). The heart of the paper is Section 4, where time series on interest differentials and exchange rate changes are generated from the Svensson (2000) model. The standard UIP test is applied to the simulated data and the β -coefficients are studied for different choices of parameters values. In Section 5, the findings are analyzed in terms of the Froot and Frankel (1989) decomposition of the deviations from UIP into parts due to the presence of risk premia and forecast errors. Section 6 concludes.

2 The puzzle

The standard test of UIP is to regress *ex post* exchange rate changes on lagged interest differentials as in (1) and investigate whether $[\alpha, \beta]$ equals $[0, 1]$. Alternatively, a constant risk premium α is allowed and only the hypothesis that β equals unity is tested.

$$s_{t+1} - s_t = \alpha + \beta (i_t - i_t^*) + \varepsilon_{t+1}. \quad (1)$$

Since expected exchange rate changes are unobservable, (1) is a joint test of the UIP hypothesis $E_t \Delta s_{t+1} = i_t - i_t^*$ and the rational expectation hypothesis $s_{t+1} = E_t[s_{t+1}] + \varepsilon_{t+1}$, where ε_{t+1} is white noise. If $[\alpha, \beta] \neq [0, 1]$ in (1), UIP does not hold or the exchange rate expectations are systematically erroneous. The *ex post* deviation from UIP in (1), ε_{t+1} , can be divided into a risk premium rp_{t+1} and a forecast error ν_{t+1} :

$$\varepsilon_{t+1} = rp_{t+1} + \nu_{t+1}. \quad (2)$$

As shown by Fama (1984), a small (below 0.5) or negative OLS estimate of β in (1) implies that the variance of the risk premium exceeds the variance of the expected exchange rate change:

$$\text{Var}(rp_{t+1}) > \text{Var}(E_t \Delta s_{t+1}) \quad (3)$$

For instance, the typical finding that β equals -3 in (1) requires that the variance of the risk premium is at least four times as large as the variance of the expected exchange rate changes (see Meredith, 2000). This is puzzling because it is difficult to generate risk premia of the required magnitude. Numerous unsuccessful attempts to model a large and variable exchange rate risk premium have been made. Hodrick (1989) provides a survey of this literature.¹ The problem is similar to the equity risk premium puzzle of Mehra and Prescott (1985). For instance, the exchange rate risk premium in a standard consumption capital asset pricing model equals $\gamma \text{Cov}(s_t, c_t)$, where γ is the relative risk aversion coefficient and c_t is period t consumption. Because the variance of consumption is small relative to the deviations from UIP, consumers have to be implausibly risk averse for this class of models to generate substantial risk premia.

Several authors argue that the exchange rate risk premium puzzle is not as puzzling as Fama (1984) and others have made it. For instance, Hodrick and Srivastava (1986) show that the Lucas (1982) general equilibrium model is capable of satisfying (3), i.e. may generate risk premia with a variance that

¹ A possible exception is De Santis and Gerard (1997, 1998), who are able to predict a non-trivial portion of the excess returns in foreign exchange markets using GARCH-models. Other studies have however failed to detect significant exchange rate risk premia in similar models based on time-varying second moments (Giovannini and Jorion, 1989, Alexius and Sellin, 1999)

exceeds the variance of the expected exchange rate changes. Meredith (2000) instead claims that since the empirical failure of UIP is well documented, and it requires that (3) is satisfied, the unobservable exchange rate risk premium *must* be large and highly variable. If models have failed to capture this, it is the models, not the facts, that are to blame. A third possibility is to argue that while it is difficult to generate a large and highly variable risk premium, models of expected exchange rate changes have not necessarily been more successful. Nominal exchange rate changes are notoriously difficult to predict, especially at the short forecasting horizons that match short-term interest rates (Meese and Rogoff, 1983). According to equation (3), the variance of the risk premium has to be large *relative* to the variance of expected exchange rate changes. This condition can be satisfied either if the variance of the risk premium is large or if the variance of expected exchange rate changes is small. If nominal exchange rate changes are completely unpredictable, as Meese and Rogoff (1983) and others claim, the right hand side of (3) is zero and an infinitely small variance of the risk premium is sufficient to satisfy the inequality in (3).

Typically, little is known about expected exchange rate changes. One of the advantages of using a full scale open economy macro model rather than a partial equilibrium finance model is that expected exchange rates are endogenously determined. It is then possible to analyze the variance of expected exchange rate changes relative to the variance of the risk premium and also the covariances between these variables. Froot and Frankel (1989) derive the following decomposition of the *ex post* deviations from UIP into parts attributable to the exchange rate risk premium, β_{rp} , and forecast errors,

β_{fe} :

$$\beta_{OLS} = 1 - \beta_{rp} - \beta_{fe}, \quad (4)$$

where β_{fe} equals minus the covariance of the forecast errors ν_{t+1} and the interest differential:

$$\beta_{fe} = -\text{Cov}(\nu_{t+1}, (i_t - i_t^*)) / \text{Var}(i_t - i_t^*). \quad (5)$$

β_{rp} consists of the variance of the risk premium minus the covariance of the risk premium and expected exchange rate changes divided by the variance of the interest differential:

$$\beta_{rp} = [\text{Var}(rp_{t+1}) - \text{Cov}(rp_{t+1}, E_t \Delta s_{t+1})] / \text{Var}(i_t - i_t^*). \quad (6)$$

All elements of (5) and (6) except the variance of the interest differentials are unobservable. They can however be identified numerically in the simulated data from the Svensson (2000) model. An open economy macro model can in principle generate negative β -coefficients from any of the terms in (5) and (6), for instance a negative covariance between expected exchange rate changes and the risk premium. Partial equilibrium models that focus on the first order conditions of consumers given exogenous stochastic processes for exchange rates and interest rates do not contain such a potential. On the other hand, the variance of the exchange rate risk premium is an exogenous parameter in the Svensson (2000) model. In this paper, we want to investigate the potential of an open economy macro model to generate deviations from UIP through the variance and covariance terms in (3), (5) and (6).

3 The model

The Svensson (2000) model consists of a supply equation, a demand equation, a UIP relationship and a monetary policy reaction function. Following Woodford (1996) and Rotemberg and Woodford (1997), the supply function in (7) is derived from microfoundations in Svensson (2000):

$$\pi_{t+2} = \alpha_\pi \pi_{t+1} + (1 - \alpha_\pi) \pi_{t+3|t} + \alpha_y [y_{t+2} + \beta_y (y_{t+1} - y_{t+1|t})] + \alpha_q q_{t+2|t} + \varepsilon_{t+2} \quad (7)$$

$x_{t+\tau|t}$ denotes the rational expectation of a variable $x_{t+\tau}$ conditional on the information available at t . π_t is the inflation rate, q_t is the real exchange rate defined as $s_t + p_t^* - p_t$ and ε_t is a (cost push) supply shock. y_t is the output gap defined as aggregate demand minus the natural output level, $y_t^d - y_t^n$. According to (7), current inflation is a function of lagged inflation, expected inflation, the output gap, the real exchange rate and the cost push shock. The natural level of output follows an AR(1) process:

$$y_{t+1}^n = \gamma^n y_t^n + \eta_{t+1}^n \quad (8)$$

where η_{t+1}^n is a productivity shock. Hence, there are two different supply shocks, a cost push shock ε_t and a productivity shock η_t^n .

Demand for domestically produced goods is given by

$$y_{t+1} = \beta_y y_t - \beta_\rho \rho_{t+1|t} + \beta_y^* y_{t+1|t}^* + \beta_q q_{t+1|t} - (\gamma^n - \beta_y) y_t^n + \eta_{t+1}^d - \eta_{t+1}^n \quad (9)$$

y_t^* denotes the foreign output gap and η_{t+1}^d is an i.i.d. demand shock. The output gap is a function of the lagged output gap, the foreign output gap, the real exchange rate, the real interest rate, the demand shock and the

productivity shock. This aggregate demand equation is derived from (some) microfoundations in Svensson (2000). The long real interest variable ρ_t is the sum of current and expected future short real interest rates, measured as deviations from the natural real interest rate. The latter is assumed to be constant and normalized to zero:

$$\rho_t = \sum_{\tau=0}^{\infty} (i_{t+\tau} - \pi_{t+\tau|t}) \quad (10)$$

The nominal exchange rate is determined by a UIP relationship:

$$i_t - i_t^* = s_{t+1|t} - s_t + \varphi_t, \quad (11)$$

where φ_t is the shock to the foreign exchange rate risk premium. The risk premium follows an AR(1) process:

$$\varphi_{t+1} = \gamma_{\varphi} \varphi_t + \xi_{\varphi,t+1}. \quad (12)$$

For simplicity, the Foreign country is not modelled as elaborately as the Home country. Foreign output and foreign inflation follow univariate AR(1) processes:

$$\pi_{t+1}^* = \gamma_{\pi}^* \pi_t^* + \varepsilon_{t+1}^* \quad (13)$$

and

$$y_{t+1}^* = \gamma_y^* y_t^* + \eta_{t+1}^*. \quad (14)$$

Furthermore, foreign monetary policy is assumed to be conducted according to a Taylor rule:

$$i_t^* = f_\pi^* \pi_t^* + f_y^* y_t^* + \xi_{it}^*, \quad (15)$$

where ξ_{it}^* is the foreign monetary policy shock. The central bank sets a nominal short-term interest rate to minimize the expected value of the loss function given all information available at t . The period t loss function contains up to four arguments: domestic inflation squared, the output gap squared, the interest rate squared and the change of the interest rate squared. For simplicity, possible target levels for inflation, output and the interest rate are hence set to zero. The weights on up to three of the four arguments can be set to zero, i.e. output targeting, interest rate targeting and interest rate smoothing are possible but not necessary characteristics of the model.

$$L_t = \mu \pi_t^2 + \lambda y_t^2 + \theta i_t^2 + \zeta (i_t - i_{t-1})^2 \quad (16)$$

The central bank minimizes the discounted expected value of L_t , i.e., $E_t \sum_{\tau=0}^{\infty} \delta^\tau L_{t+\tau}$. Its choice of an optimal short interest rate can be formulated as a linear regulator problem. The model can be rewritten in state-space form and solved numerically for specific parameters values using the algorithm presented in Oudiz and Sachs (1985). This is discussed in detail in Svensson (2000).

In the discretionary monetary policy regime under consideration, the forward looking variables $x_t = [q_t, \rho_t, \pi_{t+2,t}]$ are linear functions of the predetermined variables $X_t = [\pi_t, y_t, \pi_t^*, y_t^*, i_t^*, \varphi_t, y_t^n, q_{t-1}, i_{t-1}, \pi_{t+1|t}]$. Hence, $x_t = HX_t$ where H is endogenously determined. The solution to the model

includes, among other things, the optimal central bank reaction function. It shows how the (short-term) interest rate is set as function of the predetermined variables.

$$i_t = fX_t \tag{17}$$

The dynamics of the model imply that it takes time before monetary policy affects the economy. The short interest in period t is included in the real interest variable ρ_t , which affects demand one period ahead. The output gap in $t + 1$ in turn affects inflation in $t + 2$. Hence, the short interest in period t affects inflation with a control lag of two periods.

Given the optimal interest rate rule and the resulting dynamics of the model, it can be used to generate time series on e.g. interest differentials and exchange rate changes.

4 Simulations

The model is solved numerically for each set of parameter values. By feeding independent, normally distributed shocks into (7), (8), (9), (12), (13), (14), and (15), time series on interest differentials and *ex post* exchange rate changes can be generated. For each set of parameter values, 100 samples of 100 observations are created. The standard UIP test in (1) is applied to the simulated data and the average β -coefficients are collected. The purpose of this exercise is two-fold. First, we are interested in what ranges of β -coefficients that emerge from the model given realistic choices of parameter values. Second, since negative and large β -coefficients are observed

Table 1: Parameter estimates and ranges of realistic values

<i>Source</i>	α^π	α^y	α^q	β^y	β^q	β^p	γ_φ	row
LS (2000)	0.6	0.08	0.01	0.8	0.039	0.07	0.8	1
MC (1998)	0.6	0.2	0.1	0.5	0.1	0.5	0.0	2
RS (1998)	-	0.14	-	0.91	-	0.1	-	3
OW (1999)	-	0.31-0.39	-	0.47-0.77	-	0.23-0.4	-	4
BH (1999)	-	-	-	0.8	0.2	0.5	-	5
PS (2000)	-	0.11-0.33	-	0.89-0.94	-	0.1-0.12	-	6
FS (2000)	0.48	0.18	-	-	-	0.06	-	7
GR (2000)	0.71	0.13	-	0.88	-	0.09	-	8
range	0.48-0.71	0.08-0.39	0.01-0.1	0.47-0.94	0.039-0.2	0.06-0.5	0.0-0.8	9

empirically, we want to know what it takes to obtain negative and large β -coefficients from the model.

4.1 Choosing realistic parameter values

The model parameter vector P contains 20 parameters: $[\alpha_\pi, \alpha_y, \alpha_q, \beta_y, \beta_y^*, \beta_p, \beta_q, f_\pi^*, f_y^*, \gamma_y^n, \gamma_\pi^*, \gamma_y^*, \gamma_\varphi, \sigma_\varepsilon^2, \sigma_n^2, \sigma_d^2, \sigma_{\varepsilon^*}^2, \sigma_{\eta^*}^2, \sigma_{\xi^*}^2, \sigma_{\xi^*}^2]$. Svensson does not estimate the open economy model but selects reasonable parameter values. Table 1 shows his choices of parameter values and estimates from other studies using similar models. The "realistic ranges" in the final row are set from the smallest to the largest value encountered. This is done to avoid complete arbitrariness in the choice of parameter values; the procedure does not aspire to be highly scientific. Alternative parameter values are taken from Meredith and Chinn (1998), (MC in Table 1), Rudebusch and Svensson (1999), (RS), Orphanides and Wieland (1999), (OW), Batini and Haldane (1999), (BH), Smets (2000), (FS), and Rudebusch (2000), (GR).

α_π captures the weight on lagged inflation relative to expected future

inflation in the Phillips curve in (7). The latter enters with a coefficient $1-\alpha_\pi$, i.e. the coefficients on lagged and expected inflation sum to one. The smaller α_π is, the more important are rational inflation expectations or the more forward looking is the Phillips curve. α_π has been set to zero (McCallum, 1997) as well as to one (Ball, 1999; Svensson 1997) within this class of models. Svensson (2000) and Meredith and Chinn (1998) use $\alpha_\pi = 0.6$. Rudebusch (2000) obtains a point estimate of 0.71. The smallest empirical estimate in Table 1, 0.48, is taken from Smets (2000).

α_y is the effect of the output gap on inflation. The largest parameter value here, 0.39, stems from Orphanides and Wieland (2000). They estimate Phillips curves for the United States and the EURO area. The Svensson (2000) choice, 0.08, constitutes the lower boundary for the range of realistic values.

α_q and β_q are the open economy parameters, capturing the effects of the real exchange rate on supply and demand. For the United States, the real exchange rate is often found to have insignificant effects on inflation and output, i.e. α^q and β^q are zero. For other large economies, they are typically significant but small. Meredith and Chinn (1998) set these coefficients to 0.1 based on the IMF's model MULTIMOD for the G7 countries. Svensson uses 0.01 for both α_q and β_q .

The output gap is highly autocorrelated in all studies. Smets (2000) obtain the highest value of β^y , 0.94. The Orphanides and Wieland (1999) value of 0.47 constitutes the lower bound of β^y . It turns out that β^y is one of the most important parameters for the results from the UIP tests. The autocorrelation of the foreign output gap, β_y^* , on the other hand, hardly matters

at all given the rudimentary model of the foreign country. Here, Svensson's value of β_y^* , 0.8, is used throughout.

β^p is the interest rate sensitivity of demand. This is the main channel through which monetary policy affects the economy. The smallest value, 0.06, is taken from Smets (2000), closely followed by Svensson's choice of 0.07. Meredith and Chinn (1998) and Batini and Haldane (1999) use 0.5, which is the upper border of the range of realistic parameter values for β^p .

The foreign parameters, f_π^* , f_y^* , γ_π^* , and γ_y^* , have not been estimated in (these) other studies. f_π^* and f_y^* are the weights on inflation and output in the foreign Taylor rule and γ_π^* , and γ_y^* are the autocorrelation coefficients of foreign supply and demand shocks. As these parameters are inconsequential for the β -coefficients, the Svensson values are used throughout: $f_\pi^* = 1.5$, $f_y^* = 0.5$, $\gamma_\pi^* = 0.8$, and $\gamma_y^* = 0.8$.

γ_φ is the autocorrelation of the shocks to the risk premium. As exchange rate risk premia are unobservable, their autocorrelation is not easily estimated. The upper and lower boundaries are taken from Svensson, who sets this parameter to 0.8, and Meredith and Chinn (1998), who use a value of 0.0. Finally, the model contain seven shocks: A cost-push supply shock ε_t , a productivity shock η_t^n , demand shock η_t^d , a foreign supply shock ε_t^* , a foreign demand shock η_t^* , a foreign monetary policy shock ξ_{it}^* and a risk premium shock ξ_φ . The default value for the variances of all these shocks is 1.0.

It turns out that what matters most for the β -coefficients is the relative importance of the contemporaneous dynamics versus the intertemporal dynamics of the model. The intertemporal dynamics determines the time profile of the effects of lagged shocks. These movements are predictable

given the information in t and therefore necessarily consistent with UIP. The contemporaneous dynamics shows how the variables respond to new shocks. Only the latter, unexpected movements can be inconsistent with UIP. Therefore, large values of the parameters capturing the contemporaneous effects, $\alpha_y, \alpha_q, \beta_\rho$ and β_q , relative to the values of the intertemporal parameters like β^y tends to create a small or negative β -coefficient. A notable feature of the Svensson (2000) choice of parameters is that they constitute the lower bound for the realistic ranges for all the contemporaneous parameters except β^ρ , where his choice 0.07 is slightly higher than Smets' 0.06. Using the Svensson (2000) parameter values as default therefore biases the β -coefficients towards unity. The simulation results would appear much less supportive of the UIP hypothesis if e.g. the Meredith and Chinn (1998) parameter values were used as benchmark.

4.2 Shocks to the exchange rate risk premium

Given previous results in McCallum (1994) and Meredith and Chinn (1998), the endogenous response of monetary policy to risk premium shocks appears to be the primary vehicle for generating negative β -coefficients in the standard UIP test. Therefore, this mechanism is first isolated by feeding only risk premium shocks into the model. The results from this exercise are shown in Table 2. When there is only one shock, its variance does not matter for the results. Hence, $\sigma_{\xi i\varphi}^2$ is set to 1.0 throughout Table 2.

The first row in Table 2 contains the parameter values used by Svensson (2000). Given only risk premium shocks, the model generates a very small but positive β -coefficient (0.056). In rows 2 to 11, one parameter at a time

Table 2: β -coefficients given only shocks to the risk premium

β	α^π	α^y	α^q	β^y	β^q	β^p	γ_φ	λ	μ	row
0.056	0.6	0.08	0.01	0.8	0.039	0.07	0.8	0.5	1.0	1 ^a
0.160	0.48	2
0.202	.	0.39	3
0.208	.	.	0.1	4
-0.099	.	.	.	0.47	5
0.153	.	.	.	0.94	6
-0.178	0.2	7
0.051	0.06	.	.	.	8
0.050	0.5	.	.	.	9
0.201	0.0	.	.	10
0.022	1.0	.	11
-0.352	0.6	0.08	0.01	0.47	0.2	0.5	0.8	0.5	1.0	12
-1.981	0.6	0.08	0.1	0.47	0.2	0.5	0.0	0.5	1.0	13
-7.553	0.6	0.08	0.1	0.47	0.2	0.5	0.0	1.0	1.0	14

a These are the Svensson (2000) parameter values.

Dots imply that the Svensson (2000) values are used.

is varied from its smallest to its largest realistic value as defined in Table 1.

Reducing the autocorrelation of demand, β^y , to 0.47 reduces the β -coefficient to -0.099 (row 2). Increasing β^y to 0.94 increases the β -coefficient to 0.153 (row 3). Hence, the β -coefficient appears to be decreasing in the autocorrelation of output in this case.

In row 4, the relative weight on lagged versus expected future inflation in the Phillips curve, α^π , is reduced to 0.48. This increases the β -coefficient to 0.16. Rows 5 and 6 show, perhaps surprisingly, that the magnitude of the effect of monetary policy on demand, β^p , has only a small and unsystematic effect on the results from the UIP tests. Increasing β^p to 0.5 or decreasing it to 0.06 result in small reduction in the β -coefficient to 0.050 and 0.051. In row 7, the effect of the output gap on inflation, α^y , is increased to 0.39.

This increases the β -coefficient to 0.202.

Next, we come to the open economy parameters α^q , the effect of the real exchange rate on inflation, and β^q , the effect of the real exchange rate on demand. Variations in these two parameters turn out to have major effects on the results from the UIP tests. Increasing α^q to 0.1 results in a β -coefficient of 0.208. The same operation for the effect of the real exchange rate on demand, β^q , yields the first negative β -coefficient, -0.065, in row 8.

Rows 10 and 11 show that reducing the autocorrelation of the exchange rate risk premium to 0 increases the β -coefficient to 0.208 and increasing the relative weight on output stabilization, λ , from 0.5 to 1.0 decreases the β -coefficient to 0.022.

The findings in rows 1 to 11 of Table 2 suggest that a smaller autocorrelation of output and a larger effect of the real exchange rate on demand are the main measures that result in substantial reductions of the β -coefficient from the UIP tests. Furthermore, a smaller weight on lagged versus expected future inflation in the Phillips curve, a smaller effect of the real exchange rate on inflation, a larger effect of the real interest rate on demand and higher relative disutility of output variations also appear to decrease the β -coefficient. Several of these effects have to be combined to get β -coefficient that are negative and large. Combining three of them, reducing the autocorrelation of demand, increasing the interest sensitivity of demand and the real exchange rate sensitivity of demand results in a coefficient of -0.274 (row 12). Row 13 demonstrates how to obtain the smallest value of β given parameters values within the realistic ranges as defined in Table 1. It turns out to be -1.75. Increasing the relative weight on output variability in the central bank's loss

function, λ , to 1.0 quickly pulls the β -coefficient down further to -7.233 (row 14). Hence, when only shocks to the risk premium are fed into the model, it can easily generate large and negative β -coefficients. This result is consistent with the original McCallum (1994) model.

4.3 Demand and supply shocks

Next, the model is simulated with only demand shocks (Table 3). Given the Svensson parameter values in row 1 of Table 3, a β -coefficient of 0.972 is obtained. The same procedure is followed here as for the shocks to the risk premium, i.e. one parameter at a time is varied between the highest and lowest realistic values from Table 1. Then, all effects in favor of small or negative β -coefficients are combined to see how small values the model is capable of generating.

Changing the autocorrelation of output, β^y , from 0.47 to 0.94 increases the β -coefficient to 0.978 and 0.991, respectively (rows 2 and 3). Lowering β^y always decreased the β -coefficients in Table 2. Hence, different mechanisms appear to be at work in response to different types of shocks. Lowering α^π to 0.48 decreases β slightly to 0.971 (row 4). Changing the effect of the interest rate on demand, β^r , to 0.06 and 0.5 increases β to 0.975 and 0.997, respectively (rows 5 and 6). Increasing the effect of the output gap on inflation, α^y , increases the β -coefficient to 0.975 (row 7). As in Table 2, increasing the effect of the real exchange rate on demand, β^q , decreases the β -coefficient (row 8), while increasing the corresponding effect on supply, α^q , increases it (row 9). Increasing the relative weight on output gap in the loss function decreases β -coefficient very little, at the fourth decimal (row

11).

The main conclusion from Table 3, where only demand shocks are fed into the model, is that the data are consistent with UIP. The β -coefficients in the first column deviate only marginally from unity.

Table 3: β -coefficients given only demand shocks

β	α^π	α^y	α^q	β^y	β^q	β^r	γ_φ	μ	λ	row
0.972	0.6	0.08	0.01	0.8	0.039	0.07	0.8	1.0	0.5	1 ^a
0.971	0.48	2
0.975	.	0.39	3
1.038	.	.	0.1	4
0.978	.	.	.	0.47	5
0.991	.	.	.	0.94	6
0.961	0.1	7
0.975	0.06	.	.	.	8
0.997	0.5	.	.	.	9
0.953	0.0	.	.	10
0.972	1.0	11
0.928	0.5	0.1	0.005	0.8	0.164	0.018	0.8	1.0	0.5	12

a These are the Svensson (2000) parameter values.

Dots imply that the Svensson (2000) values are used.

The β -coefficients when only the two types of supply shocks are fed into the model are similar to those found in Table 3, i.e. the simulated data on exchange rate changes and interest differentials are consistent with UIP. Hence, it is primarily the endogenous response to exchange rate risk premium shocks that generates *ex post* deviations from UIP in the Svensson (2000) model.

4.4 The full model I: Baseline objective function

Next, the full model is simulated by generating independent normally distributed random values for all seven shocks. Table 5 shows the resulting β -coefficients for different choices of parameter values. Now, the relative variances of the shocks matter. In the benchmark case, all shocks have unity variances.

Table 4: β -coefficients for the full model

β	α^π	α^y	α^q	β^y	β^q	β^r	γ_φ	μ	λ	row
0.685	0.6	0.08	0.01	0.8	0.039	0.07	0.8	1.0	0.5	1 ^a
0.588	0.48	2
0.824	.	0.39	3
0.936	.	.	0.1	4
0.454	.	.	.	0.47	5
1.016	.	.	.	0.94	6
0.665	0.1	7
0.722	0.06	.	.	.	8
0.755	0.5	.	.	.	9
0.967	0.0	.	.	10
0.630	1.0	11
0.281	0.48	.	.	0.47	0.1	12
-0.931	0.1	0.08	0.01	0.1	0.2	0.1	0.9	1.0	1.0	13
-0.082	○	○	○	$\text{Var}(\varepsilon_t)$	$= 0.1$	○	○	○	○	14
0.249	○	○	○	$\text{Var}(\eta_t^d)$	$= 0.1$	○	○	○	○	15
0.488	○	○	○	$\text{Var}(\eta_t^s)$	$= 0.1$	○	○	○	○	16
-1.196	○	○	○	$\text{Var}(\xi_{\varphi,t})$	$= 10$	○	○	○	○	17

a These are the Svensson (2000) parameter values.

Dots imply that the Svensson (2000) values are used.

For the parameter values used by Svensson (2000), the β -coefficient is 0.685. This established model of endogenous monetary policy in an open economy hence generates considerable deviations from UIP. The β -coefficient

in the benchmark case is however positive and larger than typical empirical findings.

As in Table 2, the autocorrelation of the domestic output gap is one of the most important parameters for the results in terms of the β -coefficients from the UIP tests. In row 2, β^y is reduced from the Svensson choice 0.8 to the 0.47 of Orphanides and Wieland (2000). The β -coefficient then falls to 0.454. Increasing β^y to 0.94 increases the β -coefficient to 1.016.

Focusing instead on the effect of lagged versus expected future inflation in the Phillips curve, α^π , decreasing it to 0.48 reduces the β -coefficient to 0.588 (row 4). Rows 5 and 6 again show that changing the effect of the real interest rate on demand only affects the β -coefficient marginally. Decreasing β^ρ to 0.06 from the Svensson choice 0.07 reduces the β -coefficient slightly to 0.722. Increasing it to 0.5 increases the β -coefficient to 0.755. Increasing α^y , the effect of the output gap on inflation, to 0.39 increases the β -coefficient to 0.824.

As in Table 2, the open economy parameters α^q and β^q have large effects on the β -coefficient. Increasing β^q from the Svensson 0.01 to 0.1 reduces the β -coefficient to 0.665 (row 8). The same operation on the supply side, increasing α^q , increases the β -coefficient to 0.936 (row 9). Finally, reducing the autocorrelation of the risk premium shocks, γ_φ , to 0.0 increases the β -coefficient to 0.967 and increasing the relative disutility of output variations, λ , reduces it to 0.630 (row 11).

Again, more than one parameter at a time has to be changed relative to the Svensson (2000) benchmark for small or negative β -coefficients to emerge from the model. In row 12, all the different effects that work in favor

of small β -coefficients are combined, i.e. the β -coefficient is (informally) minimized given the constraint that the parameter values must be within the realistic ranges as defined in the final row of Table 1. This results in a β -coefficient of 0.281, which is obviously small but still positive.

Negative β -coefficients can be generated from the full Svensson (2000) model in several ways. First, values outside the realistic ranges as defined in Table 1 can be assigned to some key parameters. For instance, reducing the autocorrelation of the output gap, β^y , to 0.1, increasing the effect of the real exchange rate on demand, β^q , to 0.2 and setting λ , the relative disutility of output variability in the central bank loss function, to 1.0 yields a β -coefficient of -0.931 (row 13).

A second measure that may result in negative β -coefficients in the full model is to alter the relative variances of the shocks. This is illustrated in rows 14 to 17. The remaining parameters are the same here as in row 12, i.e. the combination of values that result in the smallest β -coefficient given the restriction that they belong to the realistic ranges as defined in Table 1.

In row 14, the variance of the cost push supply shocks is reduced to 0.1. The β -coefficient is then reduced to -0.082. Reducing the variance of the demand shocks to 0.1 instead increases the β -coefficient to 0.249 (row 15). Reducing the variance of the exchange rate risk premium shocks to 0.1 results in a β -coefficient of 1.023 (row 16). When the variance of the risk premium shocks is set ten times as large as the other variances (row 17), we get a β -coefficient that can be classified as negative and large (-1.196 in row 16). Given the high variance of the shocks to the risk premium, the Svensson model is very similar to the Meredith and Chinn (1998) model in this case.

They use almost identical parameter values and obtain a β -coefficient of -0.8 for short-term interest rates.

The main impression from Table 5 is that for realistic parameter values, the Svensson (2000) model typically generates β -coefficients that are below unity but positive. The benchmark model with the Svensson (2000) parameter values can however be considered biased towards the unity β -coefficient implied by UIP. As discussed in Section 4, he uses small values for the parameters working in favor of large deviations from UIP (β^q , β^r) and a large value of β^y , all of which tends to produce data that are consistent with UIP. Negative β -coefficients can be obtained from the model if the relative variances are altered or a very small value is assigned to the autocorrelation of output. The smallest β -coefficient that can be generated using realistic parameter values and unity variances is 0.28.

4.5 The full model II: Interest rate smoothing

In the previous section, the central bank stabilized inflation and output, i.e. the objective function in (16) had non-zero coefficients on the first two terms only. This is the baseline formulation in Svensson (2000). There is however ample evidence that central banks also stabilize nominal interest rates, implying that the fourth coefficient ζ in (16) is positive as well. It turns out that the β -coefficients in the standard UIP test on simulated data from the Svensson (2000) model change drastically when interest rate smoothing is introduced. Table 5 shows the results from repeating the exercise of the previous section when the central bank not only stabilizes inflation and output but also avoids drastic changes of the nominal interest rate.

Table 5: β -coefficients for the full model

β	α^π	α^y	α^q	β^y	β^q	β^r	γ_φ	ζ	λ	row
0.744	0.6	0.08	0.01	0.8	0.039	0.07	0.8	1.0	0.5	1 ^a
0.461	0.48	1.0	.	2
0.934	.	0.39	1.0	.	3
NC	.	.	0.1	1.0	.	4
0.509	.	.	.	0.47	.	.	.	1.0	.	5
0.835	.	.	.	0.94	.	.	.	1.0	.	6
0.757	0.1	.	.	1.0	.	7
0.837	0.06	.	1.0	.	8
0.765	0.5	.	1.0	.	9
0.994	0.0	1.0	.	10
0.734	1.0	1.0	11
-0.216	0.48	.	.	0.47	0.1	0.5	.	1.0	.	12
-1.033	0.48	0.39	.	0.47	0.1	0.5	0.9	1.0	.	13
-1.247	0.48	0.39	.	0.47	.	0.5	0.9	1.0	.	14
-1.344	0.48	0.39	.	0.47	0.1	0.5	0.9	1.0	1.0	15
-1.840	o	o	o	o	Var($\xi_{\varphi,t}$) =10	o	o	o	o	16

a These are the Svensson (2000) parameter values.

Dots imply that the Svensson (2000) values are used.

Cirkels (o) imply that the values are the same as in the previous row.

NC implies that the model did not converge for this set of parameter values.

Rows one to eleven in Table 5 are similar to the corresponding results in Table 4. Varying one parameter at the time from the Svensson (2000) benchmark values generates β -coefficients between 0.6 and 1.0. The interesting differences between interest rate smoothing and inflation/output stabilization appear in rows 12 to 18. Here, it is clear that as soon as several parameter at a time are altered from the Svensson values to other realistic values, negative β -coefficients emerge. In row 12, four changes are made relative to the Svensson (2000) set of parameter values in row one: The autocorrelation of inflation is decreased to 0.48, the autocorrelation of output is decreased

to 0.47, the effect of the interest rate on output is increased to 0.5, and the effect of the real exchange rate on output is increased to 0.1. This results in a β -coefficient of -0.216. Increasing the autocorrelation of the exchange rate risk premium from 0.8 to 0.9 lowers the β -coefficient further to -1.03 in row 13. In contrast to previous results in this paper, a smaller effect of the exchange rate on output, β^q , produces an even lower coefficient (-1.24 in row 14). Hence, negative and large β -coefficients can easily be generated from the model given realistic parameter values when the central bank engages in interest rate smoothing.

A second difference relative to the results for output and inflation stabilization is that altering the relative variances of the shocks does not have as drastic effects on the β -coefficient. Row 16 shows that it drop moderately from -1.3 to -1.8 when the variance of the exchange rate risk premium is scaled up by a factor ten relative to the variances of the other shocks. Other changes of the variances have even smaller effects on the β -coefficient.

5 Impulse response functions

Why does the Svensson model of endogenous monetary policy in an open economy generate data on interest differentials and exchange rate changes that are inconsistent with UIP? Because we cannot solve the model analytically, it is not possible to analyze the expression for the ex post covariance in response to specific shocks as in McCallum (1994) for shocks to the foreign exchange risk premium or Alexius (2000) for demand shocks. However, the response of interest rates, exchange rate changes, output and inflation to

different shocks for different parameter values shed some light on this issue.

The mechanism behind the negative β -coefficients is similar to what is described by Meredith and Chinn (1998). First, there is a shock to the exchange rate risk premium in (12). Through the modified UIP relationship (11), this depreciates the nominal and hence real exchange rate given that the central bank does not respond by raising the interest rate. However, because the weak exchange rate increases output as well as inflation, the central bank will raise the interest rate. Hence, there will be a depreciation and a positive interest differential in period t . However, in future periods, the exchange rate appreciates again but the interest rate remains high. This is where negative β -coefficients may emerge. Depending on the parameter values, the mechanism can be more or less pronounced.

Figures one to three shows three sets of impulse response functions for the parameter values in row 1 table 4, row 12, table 4, and row 15, table 5.

The pattern is the similar in all three sets of impulse response functions. The exchange rate initially appreciates and then depreciates. In figure 3, however, it appreciates only upon impact in period 0. The interest rate is increased and then falls back. The initial hike is much smaller when the central bank engages in interest rate smoothing. There are generally conflicting movements in output and inflation - output falls and inflation increases. The reason is that the central bank efficiently removes any non-conflicting movements. If both inflation and output tend to increase, the monetary policy is tightened to push both variables back. We observe movements only to the extent that output and inflation move in opposite direction and the central bank is unable to counteract both movements. The negative β -coefficients

stem from the time period when the exchange rate is appreciating while the interest rate is still higher than the foreign interest rate.

6 Conclusions

Tentative evidence indicates that the empirical failure of UIP is confined to short-term interest rates (Alexius, 2001, Meredith and Chinn, 1998). Short-term interest rates differ from other financial assets in that they constitute the main monetary policy instrument in most industrialized countries with flexible exchange rates. According to a relatively unexplored approach to the exchange rate risk premium puzzle, the observed *ex post* deviations from UIP stem from the co-movements of exchange rate changes and interest rate differentials as monetary policy responds to shocks.

This paper investigates whether an established open economy macro model with endogenous monetary policy can explain the exchange rate risk premium puzzle. Data on interest differentials and exchange rate changes are generated from the Svensson (2000) open economy model. UIP is tested on the artificial time series and the resulting β -coefficients are collected. The values of the model parameters are then varied systematically in order to identify the conditions under which the model produces substantial deviations from UIP.

The β -coefficients that emerge from the Svensson (2000) model are typically smaller than the unity coefficient expected from UIP, but positive. For the benchmark parameter values used by Svensson (2000), the β -coefficient is 0.68. However, the Svensson (2000) choices of key parameters are frequently

small relative to other estimates. The model produces more substantial deviations from UIP if larger values are assigned to some of the parameters. The smallest β -coefficient obtained given parameter values within the range of previous empirical estimates (as well as unity variances for all shocks) is 0.28.

Negative β -coefficients can be obtained from the benchmark Svensson (2000) model either by changing the variances of the shocks from their unit values or by choosing small values for a few key parameters that control the intertemporal dynamics of the model. If the variance of the cost push supply shocks is reduced to 0.1, a β -coefficient of -0.08 emerges. Reducing the variance of the demand shocks as well as the cost push supply shocks lowers it further to -0.19.

In the benchmark Svensson (2000) model, the central bank cares about stabilizing inflation and the output gap. If it also stabilizes the nominal interest rate, the mechanism that may generate negative β -coefficients in the standard UIP test is enhanced. With interest rate smoothing, negative and large β -coefficients (around -1.3) emerge from the model given parameter values that are within the realistic ranges. Altering the relative variances of the shocks or choosing parameter values outside of the realistic ranges however does not result in substantial additional reductions of the β -coefficient. For instance, increasing the variance of the exchange rate risk premium ten-fold only brings it down to -1.8.

Most attempts to explain the exchange rate risk premium puzzle employ partial equilibrium, microeconomic finance models. Exchange rates and interest differentials can however rewardingly be analyzed as two endogenous

variables within an open economy macro model. This paper demonstrates that the co-movements of interest differentials and exchange rate changes as they respond to shocks can in principle create negative β -coefficients in standard UIP tests.

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Figure 1 a and b: Impulse responses to a unit shocks to the exchange rate risk premium, the Svensson benchmark model. Parameter values from Table 4, row 1.

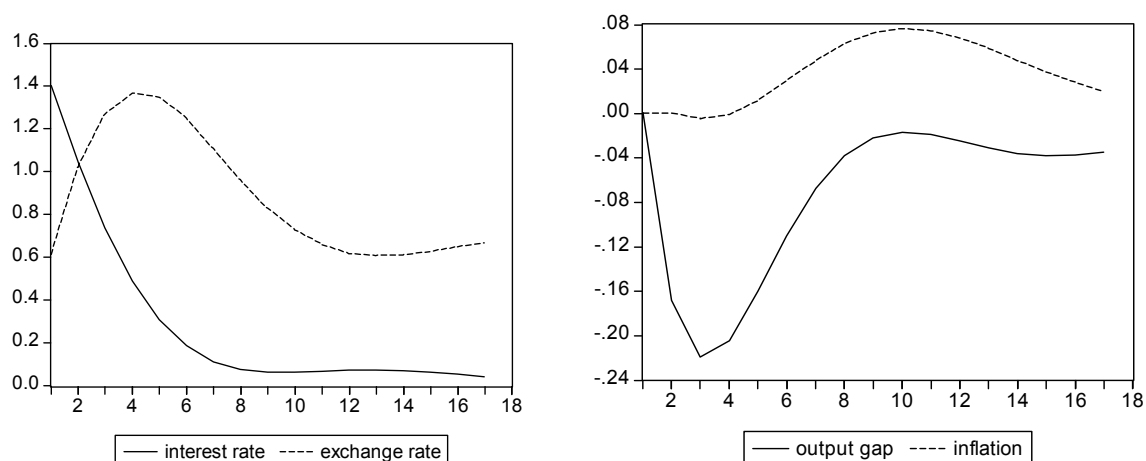


Figure 2 a and b: Impulse responses to a unit shocks to the exchange rate risk premium, small β -coefficient (0.28). Parameter values from Table 4, row 12.

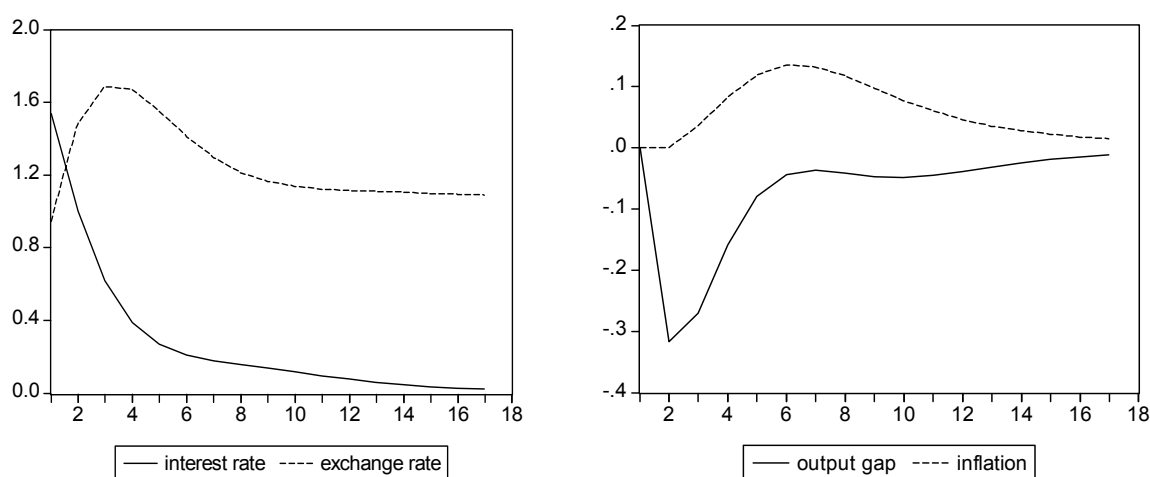


Figure 3 a and b: Impulse responses to a unit shocks to the exchange rate risk premium. Interest rate smoothing, the β -coefficient is -0.21 . Parameter values from Table 5, row 12.

