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Systematic Monetary Policy and the Forward Premium Puzzle*

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Abstract

Is systematic monetary policy a driver of the forward premium puzzle, i.e. the tendency of high interest-rate currencies to appreciate, thus strongly violating Uncovered Interest Parity (UIP)? We address this question by studying a battery of monetary policy rules in a small open economy that is subject to stationary but persistent domestic and foreign shocks. Each rule leads to model-implied UIP violations, which we derive analytically and then calibrate numerically. Our key finding is that only a forward-looking rule based on CPI inflation can account for frequently observed strong UIP violations.

Keywords: Monetary policy rules, uncovered interest parity, exchange rate dynamics, inflation-targeting.

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1 Introduction

The currency-risk premium has long eluded the profession. In a classic paper, Eugene Fama (1984) formulated the risk-premium puzzle: higher-interest rate currencies, i.e. those that come with a premium, tend to *appreciate*. Fama's finding, which has become a stylized fact of empirical international finance, has the Uncovered Interest Parity (UIP) condition in reverse. Predicated on risk neutrality of traders, UIP states that the short term interest-rate differential between two currencies should equal the expected *depreciation* rate of the higher-interest rate currency.

We argue that systematic monetary policy can help explain the currency-risk premium puzzle. In support of our argument, we report evidence that changes in the degree of UIP violations coincide with changes to monetary policy regimes. Then, on the basis of a standard parsimonious model, we characterize the exogenous shocks and monetary policy rules that deliver strong UIP violations in small open economies.

In the empirical part of this paper, we study the degree of UIP violations in six small open economies – Australia, Canada, New Zealand, Sweden, Switzerland, and the United Kingdom – and in the euro area. Adopting the Fama (1984) regression structure, we estimate a linear relationship between (i) the depreciation of the nominal exchange rate of the currency in a particular economy with the US dollar, and (ii) the corresponding interest differential. In contrast to previous studies, we allow for a structural break around the time when explicit inflation targeting was announced by the central bank in charge of a given currency. Our economy-specific sample periods include such announcements in Australia, Canada, Sweden, Switzerland, and the United Kingdom.

We find ample evidence of strong UIP violations, in accordance with Fama (1984) and subsequent related research (e.g. Burnside et al. (2009a)).[1] All of the seven exchange rates we consider go through a period over which the estimate of the slope coefficient in the regression is negative, lying between -0.5 and -3.2. In six of these cases (the exception being the New Zealand dollar), we reject the null hypothesis that the slope coefficient is equal to 1: i.e. we reject UIP.

[1] On the widespread violation of UIP see also Engel (1996), Bansal and Dahlquist (2000), Chinn and Meredith (2004, 2005) and Evans (2011).

Moreover, there is evidence that the degree of UIP violation changes with the monetary policy regime, which we believe is a novel finding. In particular, UIP violations evolve from being initially weak in Canada, or non-existent in Sweden, to being quite pronounced after the adoption of explicit inflation targeting in these economies. By contrast, UIP was violated only prior to explicit inflation targeting in the United Kingdom.

These findings lead us to investigate whether there is a *causal* link from monetary policy to UIP violations. After all, central banks are key players on foreign exchange markets and are in a position to influence the relative dynamics of interest and exchange rates. McCallum (1994) has already explored this avenue of research. He has shown that, in the presence of risk averse investors that demand an *exogenous* currency-risk premium, systematic monetary policy could explain UIP violations.

We build on this idea but, in contrast to McCallum (1994), study how alternative monetary policy rules affect the degree of UIP violation when the currency risk premium is *endogenous*. Our analysis relies on a reduced-form model of a small open economy that is affected by two mutually independent exogenous shocks, each driven by a stationary AR(1) process: a *domestic shock*, to the natural real interest rate (NIR), and a *foreign shock*, which we introduce through the real exchange rate (RER). A dynamic IS equation and an open-economy New Keynesian Phillips curve govern the transmission of these shocks to the endogenous inflation and output gap.

Using the nominal interest rate as a tool for implementing *monetary policy rules*, the domestic central bank sets the interest differential in our model. We consider six rules divided in two groups. In each group, there is a rule that maintains the inflation rate at a target level (a "strict rule"), a rule that responds to the current level of inflation and the output gap (a "Taylor-type rule"), and a rule responding to the expected level of these variables (a "forward-looking rule"). The three rules in one of the groups strictly target or respond to the inflation of *domestic* goods' prices (henceforth, domestic inflation); whereas those in the other group target or respond to *CPI* inflation (which reflects the prices of both domestic and imported goods).

Each policy rule gives rise to a *model-implied* degree of UIP violation. A particular rule generates a mapping from the exogenous shocks to the policy interest rate, and this mapping drives a particular relationship between the nominal exchange rate depreciation and the interest rate differential. This relationship, which embeds an endogenous currency risk premium, determines the slope coefficient in the Fama regression and, thus, the degree of UIP violation. Quite naturally, changes in the currency risk premium would lead to changes in the quantity of the domestic currency demanded by private agents. We keep private currency demand in the background, however, assuming that it is always accommodated by the central bank and, thus, does not influence the dynamics of the interest differential and the exchange rate.

For a policy rule to match strong empirical UIP violations, it has to give rise to a negative slope coefficient in the Fama regression. In other words, the rule has to call for lowering the interest rate in the face of higher expected nominal depreciation of the domestic currency. That said, a nominal depreciation is simply the sum of CPI inflation and RER depreciation. And a consensus in the monetary policy literature, confirmed in our data as well, is that well-behaved monetary policy rules raise the interest rate in the face of higher expected inflation. Thus, to match strong empirical UIP violations, a well-behaved policy rule needs to call for lowering the interest rate in the face of higher expected real depreciation of the domestic currency. And the resulting negative covariance between the interest differential and expected RER change needs to overwhelm the positive covariance between the interest differential and expected inflation. When this holds, we say that the policy rule meets the *sign condition*.

We are interested in policy rules that match not only the negative sign of the slope coefficient in the Fama regression but also the absolute value of this coefficient. The larger is this absolute value, the stronger is UIP violation. And the absolute value of the slope coefficient is large when the absolute value of the covariance between the rule-implied interest differential and expected nominal depreciation is large relative to the volatility of the interest differential. When this holds, we say that the policy rule meets the *magnitude condition*.

Under each of the six policy rules we consider, we derive the implied slope coefficient in the Fama regression as an explicit function of the model parameters. To assign a numerical value to such a coefficient, we calibrate the model as follows. First, we employ consensus values for deep parameters, corresponding to the interest elasticity of aggregate demand and the degree of nominal rigidities. Second, in the case of Taylor-type and forward-looking rules, we adopt standard assumptions about the responsiveness of the policy rate to (expected) inflation and output gap. Third, using data associated with the seven currencies above, we estimate parameters corresponding to (i) openness to trade and (ii) the AR(1) processes of the NIR and RER with the US dollar.

We find that the three policy rules related to domestic inflation cannot account for UIP violations. The reason is twofold. First, such rules do not allow the policy rate to react to foreign shocks, which violates the sign condition. Combined with a positive relationship between the domestic interest rate and CPI inflation, this leads to a slope coefficient in the Fama regression that is close to unity. Thus, domestic inflation-based rules are consistent with UIP holding over our sample period in New Zealand, prior to inflation targeting in Sweden, and during inflation targeting in the United Kingdom.

Turning to the CPI-based policy rules, we find that all three satisfy the sign condition.[2] Having resorted to numerical analysis in order to dissect this result, we conclude that it hinges on two effects of the foreign, i.e. RER, shock. First, as implied by the open-economy New Keynesian Phillips curve, a positive RER shock puts upward pressure on contemporaneous CPI inflation. Second, because of mean reversion, a positive RER shock leads to a negative expected change of the RER in the next period. The first effect triggers monetary tightening under the strict and Taylor-type rules. And the resulting positive change to the interest differential is then associated, because of the second effect, with a negative expected RER change. Thus, a key requirement of the sign condition is satisfied. Similar logic applies to the forward-looking rule, provided that enough persistence translates upward pressure on contemporaneous CPI inflation into upward pressure on expected CPI inflation next period.

That said, only the forward-looking CPI rule satisfies the magnitude condition as well. The strict and Taylor-type rules deliver volatile interest rate differentials that depress the absolute value of the model-implied slope coefficient in the Fama

[2] Wide-spread violations of UIP are thus consistent with the finding of Engel (2009, 2011) that central banks typically base adjustments to the policy rate on analyses of CPI, as opposed to domestic, inflation.

regression. A less volatile interest rate differential under the forward-looking rule leads to a model-implied regression slope of around -2. Thus, this rule accounts for the strong UIP violations in Australia, Canada, Switzerland and the euro area over the entire corresponding sample periods. It also accounts for strong UIP violations prior to inflation targeting in the United Kingdom and during inflation targeting in Sweden.

We organize the rest of this paper as follows. In the next section, we briefly review the related literature. Then, in Section 3, we summarize the forward premium puzzle and provide empirical evidence of UIP violations. In Section 4 we present our model. We derive analytically the model-implied violations of UIP in Section 5 and quantify and numerically analyze these results in Section 6. We conclude with Section 7. Most of the mathematical derivations are in Appendices A-D at the end of the paper.

2 Literature review

This paper belongs to the growing literature on the impact of monetary policy rules on nominal exchange rates.[3] On the empirical side, this literature has studied the in- and out-of-sample properties of model-implied exchange rate time series. By contrast, we examine whether changes of the monetary policy regimes in advanced small open economies are associated with changes in the degree of UIP violation. On the theoretical front, the related literature has focused on deriving the short-term exchange rate dynamics implied by different policy rules. We combine such dynamics with the relationship between macroeconomic shocks and policy interest rates in order to investigate which monetary policy rules can generate the strong UIP violations observed in the data.

Our analysis is in the spirit of Chinn and Meredith (2004), who also relate UIP violations to monetary policy. In contrast to them, we: (i) endogenize the source of UIP violations; (ii) explain such violations on the basis of only two, as opposed to four, exogenous shocks; and (iii) study and compare the implications of alternative monetary policy rules. Interestingly, we find that a Taylor-type rule

[3] Recent contributions include Clarida and Waldman (2008), Engel and West (2006), Engel, Mark and West (2007), Mark (2009), Molodtsova and Papell (2009) and Rogoff and Stavrakeva (2008). These papers are reviewed in Evans (2011).

– the only rule considered by Chinn and Meredith (2004) – fails to account for strong UIP violations.

There are also distinct parallels between our paper and the market microstructure analyses in Bacchetta and van Wincoop (2007, 2010). These articles have found that, on a risk-adjusted basis, investors face limited, if any, arbitrage opportunities in foreign exchange markets. Such findings help to justify our modeling decision to keep the private sector in the background in order to focus on the monetary policy of the central bank.[4]

Finally, by relating the degree of UIP violations to the volatility of the nominal interest rate, we echo arguments that others have made in different contexts. A specific reason why forward-looking policy rules give rise to strong UIP violations in our setup is that these rules generate low interest-rate volatility. This is consistent with the finding of Bansal and Dahlquist (2000) and Frankel and Poonawala (2006) that UIP violations are stronger in advanced economies, as interest rates there are less volatile than in emerging markets. Similarly, Moore and Roche (2012) show that the degree of UIP violation increases when the volatility of money growth is low.

3 Reviewing the forward premium puzzle

This section has two parts. In the first part, we revisit the classic empirical test of uncovered interest rate parity (UIP) and spell out the conditions under which UIP is violated. In the second part, we apply the UIP test to data.

3.1 Fama's test of UIP

Consider a small open economy. Let the nominal exchange rate on date-t, e_t, be the date-t price of a foreign currency in terms of the domestic currency.[5] In addition, let i_t and i^* denote respectively the short-term interest rates on domestic-

[4] The market microstructure literature has offered explanations of the forward premium puzzle that are different from, but not necessarily incompatible with the explanation we provide below. Bacchetta and van Wincoop (2007, 2011), for example, argue that infrequent adjustments to traders' foreign-exchange positions may explain strong UIP violations. Related attempts to resolve the forward premium puzzle are discussed in Burnside, Eichenbaum and Rebelo (2009a,b).

[5] All variables are expressed in logarithmic form.

and foreign-currency bonds that are risk-free in nominal terms, are issued at t and mature at $t+1$. We henceforth assume that $i^* = 0$ and thus refer to i_t as the interest differential.

If UIP holds, the expected exchange-rate depreciation equals the interest differential:

$$E_t \Delta e_{t+1} \equiv E_t \left(e_{t+1} - e_t \right) = i_t \tag{1}$$

Fama (1984) examined UIP's empirical validity by estimating the following regression:

$$\Delta e_{t+1} = \alpha + \beta i_t + u_{t+1} \tag{2}$$

where the residual term, u_{t+1}, is assumed to be independent of any variable observed on or prior to date t. Given rational expectations, UIP then implies that the OLS parameter estimates should not be significantly different from $\alpha = 0$ and $\beta = 1$, respectively. However, Fama (1984) and most of the subsequent empirical studies of UIP report significantly *negative* estimates $\widehat{\beta}$. In other words, contrary to the UIP prediction, countries with higher (lower) interest rates tend to have currencies that subsequently appreciate (depreciate).

Below, we attempt to explain the empirical finding of $\widehat{\beta} < 0$ by modeling systematic monetary policy. To this end, we relax the UIP condition (1):

$$E_t \Delta e_{t+1} = i_t - \xi_t \tag{3}$$

where ξ_t is the premium received by investors holding domestic bonds. If monetary policy creates a systematic link between the interest differential and the currency risk premium – i.e. if $Cov\left(\xi_t, i_t\right) \neq 0$ – then an OLS estimate of the slope coefficient in regression (2) would converge to:

$$\text{plim}\widehat{\beta} = \frac{Cov\left(\Delta e_{t+1}, i_t\right)}{V\left(i_t\right)} = 1 - \frac{Cov\left(\xi_t, i_t\right)}{V\left(i_t\right)} \neq 1 \tag{4}$$

Henceforth, we keep ξ_t in the background and conduct the analysis in terms of $Cov\left(\Delta e_{t+1}, i_t\right)$, $V\left(i_t\right)$ and their drivers.

3.2 Empirical evidence

We investigate violations of UIP in the case of seven currencies: the Australian dollar (AUD), British pound (GBP), Canadian dollar (CAD), Swedish krona (SEK), Swiss franc (CHF), New Zealand dollar (NZD) and the euro. For each of these currencies, we consider the exchange rate with the US dollar (USD) and estimate the Fama regression in (2). We use quarterly series of the spot exchange rates on the 25th of March, June, September and December (or, if a weekend, the last preceding work day). We combine these series with matching series of three-month forward exchange rates, which equal the relevant interest differentials under covered interest parity.[6]

In running each regression, we attempt to identify whether a change in the monetary policy regime coincided with a change in the degree of UIP violations. For the first five currencies above, the available data cover a long period during which the corresponding central bank did not have an explicit inflation target, as well as a long subsequent period of inflation targeting. In these cases, we test whether there is a structural break in the regression within a window of eight quarters centered at the quarter in which inflation targeting was explicitly adopted.[7] And when the statistical test identifies a structural break, we estimate two regressions on two non-overlapping time periods: before and after the break.

3.2.1 Results on UIP violation

The regression results, which we report in **Table 1**, deliver two messages. First, six out of the seven exchange rates have experienced a period of strong UIP violation. In accordance with previous empirical studies, we estimate negative slope coefficients, between -0.5 and -3.2. This range is in line with Engel's (2011) estimates for advanced economies' US-dollar exchange rates over a similar period.

[6] Our dataset stops in 2008:Q3, i.e. just before the global financial crisis affected currency markets in ways that we cannot account for with our model.

[7] We adopt Ball and Sheridan's (2004) convention for the starting date of inflation targeting as the quarter when the central bank announced it began targeting constant inflation. Concretely, this happened in 1994:Q3 in Australia, 1994:Q1 in Canada, 1999:Q1 in Switzerland, 1993:Q1 in the United Kingdom, and 1995:Q1 in Sweden. To the best of our knowledge, the only other paper testing for possible breaks in UIP regressions is Kugler (2000), who analyzes the 1992-3 EMS crisis.

Second, in the case of three out of the five currencies for which there are long enough time-series data, the degree of UIP violations changes with the monetary policy regime. Namely, explicit inflation targeting goes hand in hand with stronger UIP violations by the exchange rates of the CAD and SEK with the USD. In line with the discussion at the end of Section 3.1, this is due to a lower volatility of the interest differential after the policy regime shift. By contrast, the GBP-USD exchange rate violated UIP only *before* explicit inflation targeting led to a positive co-movement between the interest differential and nominal exchange-rate depreciation.

Interestingly, **Table 1** indicates that the explanatory power of the Fama regression – as captured by $\bar{R}^2$ – is higher than that reported in the literature; for example, Engel (1996) and Evans (2011). One possible reason could be that our results are based on longer time series than those reported in some of the previous studies. Another reason may be rooted in the structural breaks we allow for: the level of $\bar{R}^2$ drops materially in the case of the CAD, GBP and SEK regressions if we estimate them over the entire sample period.

3.2.2 Dissecting the results

To delve into the estimates of the slope coefficient in the Fama regression, we use the identity linking the nominal depreciation rate (e) to the real depreciation rate (q) and inflation (π):

$$\Delta e_t \equiv \Delta q_t + \pi_t \tag{5}$$

to rewrite the first equality in expression (4) as follows:

$$\text{plim}\widehat{\beta} = \frac{Cov\left(E_t \pi_{t+1}, i_t\right)}{V\left(i_t\right)} + \frac{Cov\left(E_t \Delta q_{t+1}, i_t\right)}{V\left(i_t\right)} \tag{6}$$

This leads us to decompose regression equation (2) in two parts:

$$\pi_{t+1} = \alpha_1 + \beta_1 i_t + u_{1,t+1} \tag{7}$$

$$\Delta q_{t+1} = \alpha_2 + \beta_2 i_t + u_{2,t+1} \tag{8}$$

where $\alpha_1 + \alpha_2 = \alpha$, $\beta_1 + \beta_2 = \beta$, and $u_{1,t+1} + u_{2,t+1} = u_{t+1}$. **Tables 1.1** and **1.2** report OLS estimates of the coefficients in these equations. The value of $\widehat{\beta}_1$ is positive for virtually all currencies in our sample, and its 95% confidence interval invariably includes both zero and one. This result is consistent with well-behaved policy rules that call for tightening in the face of higher expected inflation: $Cov\,(E_t\pi_{t+1}, i_t) > 0$. In turn, we obtain that $\widehat{\beta}_2$ is negative and statistically different from unity in virtually all cases. This result, which stems from negative sample estimates of $Cov\,(E_t\Delta q_{t+1}, i_t)$, holds for all currency / time periods for which the estimates of the Fama slope coefficient – i.e. $\widehat{\beta}$ in equation (2) – are statistically different from one (recall **Table 1**).

4 Model

We now describe the analytic environment. First, we introduce the exogenous shocks affecting the modelled economy. Second, we define how these shocks drive inflation and output gap over time. Third, we specify six monetary policy rules that we consider as candidates for explaining the empirical violations of UIP.

4.1 Real uncertainty

There are two exogenous sources of uncertainty, stemming from the *real* exchange rate (RER) and the *natural* real interest rate (NIR). We assume that the RER, q_t, and the *deviation* of the NIR from its long-run mean, r_t^n, follow mutually independent AR(1) processes:[8]

$$r_t^n = \rho\, r_{t-1}^n + \varepsilon_t \tag{9}$$

$$q_t = \mu q_{t-1} + \eta_t \tag{10}$$

where $\rho \in [0, 1)$, $\mu \in [0, 1)$ and, in line with long-run PPP, the steady-state RER is zero, $E\,(q_t) = 0$. In addition, ε_t and η_t are i.i.d. zero-mean variables with

[8] Throughout the paper we abstract from constant terms in the processes of both exogenous and endogenous variables. This tightens the exposition and is without loss of generality in our context, as we focus on the slope coefficient in the Fama (1984) regression, which depends only on second moments.

time-invariant variances $V(\varepsilon)$ and $V(\eta)$, respectively, and $Cov\left(\varepsilon_t, \eta_{t\pm j}\right) = 0$ for any j. Below, we use the ratio of the innovations' variances: $\upsilon \equiv \frac{V(\eta)}{V(\varepsilon)}$.

We think of the NIR and RER as corresponding respectively to domestic and foreign shocks. The NIR is related to the evolution of both real and financial frictions in the economy and has been modeled as an exogenous variable by Benati and Vitale (2007), Laubach and Williams (2003) and Woodford (2003), among others. In turn, in an environment with a perfect exchange rate pass-through (which we assume below), fluctuations in a small open economy's RER are tightly linked to fluctuations in its terms-of-trade. Thus, our assumption of exogenous RER shocks finds indirect support in Kehoe and Ruhl (2008), who derive conditions under which terms-of-trade shocks can be modeled as exogenous. In addition, Mendoza (1995) finds exogenous terms-of-trade shocks to be persistent, in line with our AR(1) specification for the RER.[9]

Note that the process in (10) allows us to rewrite the limiting value of the slope coefficient of the Fama regression in (6) as:

$$\text{plim}\widehat{\beta} = \frac{Cov\left(E_t \Delta e_{t+1}, i_t\right)}{V\left(i_t\right)} = \frac{Cov\left(E_t \pi_{t+1}, i_t\right)}{V\left(i_t\right)} - \frac{(1-\mu)\, Cov\left(q_t, i_t\right)}{V\left(i_t\right)} \qquad (11)$$

We now state two conditions for strong empirical violations of UIP, i.e. values of $\widehat{\beta}$ substantially below zero. The first is a *sign condition*: $\widehat{\beta} < 0$ if and only if a higher interest differential tends to precede a lower nominal depreciation rate: $Cov\left(\Delta e_{t+1}, i_t\right) < 0$. That said, the empirical results in Section 3.2.2 call for well-behaved policy rules, i.e. for $Cov\left(E_t \pi_{t+1}, i_t\right) \geq 0$, which turns out to be a feature of our theoretical model as well. Thus, equation (11) implies that $Cov\left(q_t, i_t\right) > 0$ is necessary for the sign condition to be satisfied.[10] In addition, the second, *magnitude condition* states that: to account for a large absolute value of $\widehat{\beta}$, the absolute value of $Cov\left(\Delta e_{t+1}, i_t\right)$ should be sufficiently larger than the variance of the interest differential, $V\left(i_t\right)$.

[9] Mendoza (1995) also finds evidence that terms-to-trade shocks account for 45–60% of the observed output variability in different country groups.

[10] A positive $Cov\left(q_t, i_t\right)$ would be consistent with a finding of Clarida, Gali and Gertler (1998) that real exchange rate depreciations tend to induce central banks to tighten monetary policy.

4.2 Dynamics in the economy

To specify the evolution of inflation and the output gap over time, we start with a standard New Keynesian Phillips curve and a dynamic IS relationship for a closed economy:

$$\pi_t^D = \kappa x_t + bE_t\pi_{t+1}^D \quad (12)$$

$$x_t = E_t x_{t+1} - \sigma(i_t - E_t\pi_{t+1}^D - r_t^n) \quad (13)$$

where π_t^D is the inflation of the price of domestic goods, henceforth *domestic inflation*, $i_t - E_t\pi_{t+1}^D$ is the ex-ante real interest rate, $\sigma > 0$ corresponds to the representative household's inverse intertemporal elasticity of substitution (equivalently, the interest elasticity of aggregate demand), $\kappa > 0$ reflects the degree of price stickiness, and $b \in (0,1)$ depends on the stochastic discount factor.

Then, we refer to Galí and Monacelli (2005), who assume: (i) a utility function with a constant elasticity of substitution between home and foreign goods; and (ii) constant foreign-currency prices and perfect exchange-rate pass-through.[11] Log-linearization then leads to the following approximate relationship between CPI inflation (π_t), domestic inflation (π_t^D) and the nominal depreciation rate (Δe_t):

$$\pi_t = (1-a)\pi_t^D + a\Delta e_t \quad (14)$$

where $a \in [0,1]$ is the share of domestic consumption allocated to foreign goods, i.e. an index of the economy's *openness*.

Finally, using equations (5) and (14), we obtain the open-economy version of (12)-(13):

$$\pi_t = \kappa x_t + bE_t\pi_{t+1} + \frac{a}{1-a}[\Delta q_t + b(1-\mu)q_t] \quad (15)$$

$$x_t = E_t x_{t+1} - \sigma(i_t - E_t\pi_{t+1}^D - r_t^n) \quad (16)$$

[11] Perfect pass-through amounts to the law of one price holding across borders at all times. This assumption allows us to work with the parsimonious dynamic IS specification in equation (16). For the implications of modeling incomplete pass-through see Corsetti, Dedola and Leduc (2008) and Monacelli (2005).

4.3 Monetary policy rules

We study six monetary policy rules, which the central bank implements by setting the nominal interest rate, i_t. A rule is allocated to one of three groups, depending on whether it: (i) attains a strict inflation target, (ii) is of the Taylor type, responding to concurrent inflation and output gap, or (iii) is forward looking, responding to the expected level of inflation and the output gap next period. Within each group, there are two rules: one rule is based on *CPI* inflation and the other on *domestic* inflation. We denote the CPI-based rules with *IT*, *TR* and *FW*, respectively, and the corresponding domestic inflation-based rules with *ITd*, *TRd* and *FWd*.

Under a strict rule, the central bank sets the policy rate so that the targeted inflation rate is zero on each date. Thus, the *IT* rule leads to $\pi_t^{IT} = E_t\left(\pi_{t+1}^{IT}\right) = 0$ at all t. Likewise, the strict domestic inflation-targeting rule *ITd* leads to $\pi_t^{D,ITd} = E_t\left(\pi_{t+1}^{D,ITd}\right) = 0$ at all t. In analyzing each of these rules, we follow Woodford (2003) and assume that the central bank can attain its target after observing the current NIR, r_t^n, the structural parameters b, κ and σ, as well as a.

Under the Taylor-type rules the central bank sets its policy rate as follows:

$$i_t^{TRd} = \phi_\pi \pi_t^{D,TRd} + \phi_x x_t^{TRd} \quad (17)$$

$$i_t^{TR} = \phi_\pi \pi_t^{TR} + \phi_x x_t^{TR} \quad (18)$$

where $\phi_\pi > 1$ and $\phi_x \geq 0$ satisfy the Taylor principle for determinacy. Equations (17) and (18) indicate a neutral policy rate of zero, which is consistent with the specification of the exogenous shocks in (9) and (10) and the assumption of a zero target inflation rate underlying the Phillips curve and IS equations in Section 4.2.

Finally, a central bank that has adopted a forward-looking rule sets the policy rate in response to expected (domestic or CPI) inflation and expected output gap deviations from their (zero) targets:

$$i_t^{FWd} = \varphi_\pi E_t \pi_{t+1}^{D,FWd} + \varphi_x E_t x_{t+1}^{FWd} \quad (19)$$

$$i_t^{FW} = \varphi_\pi E_t \pi_{t+1}^{FW} + \varphi_x E_t x_{t+1}^{FW} \quad (20)$$

where $\varphi_\pi > 1$, $\varphi_x \geq 0$.[12]

Henceforth, we suppress the rule-dependent superscripts of π_t, π_t^D and x_t wherever this does not cause confusion.

5 Model-implied UIP violations

In this section we derive the slope coefficient, $\widehat{\beta}$, that would be obtained from regressing the realized depreciation rate, Δe_{t+1}, on the interest differential, i_t, when a particular monetary policy rule is in place in the modelled economy. Abstracting from estimation noise, we focus on the asymptotic value of $\widehat{\beta}$ in (11). To lighten the notation, we henceforth omit the "hat" when referring to specific model-implied coefficient estimates. Instead, we signal the underlying policy rule with a superscript, e.g. β^{FW}, β^{TRd}, etc.

5.1 Policy rules responding to domestic inflation

None of the domestic inflation-based rules can satisfy the sign condition, implying that neither β^{ITd} nor β^{TRd}, nor β^{FWd} can be negative.[13] To see why, recall the discussion following expression (11), in which we showed that a strictly positive relationship between the policy rate i_t and foreign (i.e. q) shocks is a necessary condition for obtaining a negative slope coefficient of the Fama regression. By equations (12)-(13), however, this condition is violated by the domestic inflation-based rules, which allow i_t to respond *only* to domestic (i.e. r^n) shocks.

To differentiate among β^{ITd}, β^{TRd} and β^{FWd}, we note the following:

- Under strict domestic-inflation targeting, the monetary authority attains $\pi_t^D = 0$. Together with equations (14) and (5), this implies that π_t is driven entirely by q_t. Since i_t^{ITd} is independent of q_t, it then follows that $Cov\left(E_t\pi_{t+1}, i_t^{ITd}\right) = Cov\left(q_t, i_t^{ITd}\right) = \beta^{ITd} = 0$.
- Domestic inflation-based Taylor-type and forward-looking rules, which are well-behaved but respond only to domestic shocks, lead to $\beta^{TRd} = \frac{Cov\left(E_t\pi_{t+1}, i_t^{TRd}\right)}{V\left(i_t^{TRd}\right)} =$

[12] Note that we do not consider exogenous policy shocks. The key conclusions of the analysis would not change if we allowed for such shocks, provided they are independent of all other variables in the model.

[13] The statements in this subsection are substantiated in Appendix A.

$\frac{Cov\left(E_t\pi^D_{t+1},i^{TRd}_t\right)}{V\left(i^{TRd}_t\right)} > 0$ and $\beta^{FWd} = \frac{Cov\left(E_t\pi_{t+1},i^{FWd}_t\right)}{V\left(i^{FWd}_t\right)} = \frac{Cov\left(E_t\pi^D_{t+1},i^{FWd}_t\right)}{V\left(i^{FWd}_t\right)} > 0$. The inequalities in these expressions confirm that the rules are well-behaved.

Domestic inflation-based rules could thus explain some of the empirical results in Section 3.2.1. In particular, they would be consistent with the failure to reject the UIP null for New Zealand (recall **Table 1**). They would also be consistent with the positive slope estimates in the Fama regression for Sweden prior to inflation targeting, and for the United Kingdom during inflation targeting.

5.2 Strict CPI-inflation targeting

Strict CPI targeting satisfies unambiguously the sign condition.[14] To see this, note that attaining $\pi_t = E_t(\pi_{t+1}) = 0$ for all t requires the following policy rate:

$$i^{IT}_t = r^n_t + \frac{a}{1-a}\left(1-\mu+\frac{b(1-\mu)^2+2-\mu}{\kappa\sigma}\right)q_t - \frac{a}{\kappa\sigma(1-a)}q_{t-1} \qquad (21)$$

This policy rate relates positively to the RER: $Cov\left(i^{IT}_t, q_t\right) > 0$. The reason is rooted in equation (15), which implies that an RER depreciation, i.e. a rise in q_t, puts upward pressure on CPI inflation, π_t. In order to keep $\pi_t = 0$ in the face of this upward pressure, the central bank needs to lower the output gap, x_t. By equation (16), the central bank lowers x_t by raising the policy rate, i^{IT}_t. Thus, a positive shock to q_t raises i^{IT}_t.

Then, given that $E_t(\pi_{t+1}) = 0$ delivers $Cov(q_t, E_t(\pi_{t+1})) = 0$ and given $Cov\left(q_t, i^{IT}_t\right) > 0$, equation (11) implies:

$$\beta^{IT} = -\frac{(1-\mu)\,Cov\left(q_t, i^{IT}_t\right)}{V\left(i^{IT}_t\right)} < 0 \qquad (22)$$

[14] The statements in this subsection are substantiated in Appendix B.

5.3 Policy rules responding to CPI inflation

The CPI-based policy rules in equations (18) and (20) imply that the equilibrium inflation and output gap depend linearly on the exogenous variables:[15]

$$\begin{aligned} \pi_t &= \delta^\pi r_t^n + \zeta^\pi q_t + \theta^\pi q_{t-1} \Rightarrow E_t \pi_{t+1} = \delta^\pi \rho r_t^n + (\mu \zeta^\pi + \theta^\pi)\, q_t \\ x_t &= \delta^x r_t^n + \zeta^x q_t + \theta^x q_{t-1} \Rightarrow E_t x_{t+1} = \delta^x \rho r_t^n + (\mu \zeta^x + \theta^x)\, q_t \end{aligned} \quad (23)$$

where the value of the coefficient vector $\{\delta, \zeta, \theta\}^{\pi,x}$ changes with the policy rule.

The coefficients governing the impact of the NIR, r_t^n, on inflation π_t and the output gap x_t are unambiguously positive irrespective of the policy rule: i.e. $\delta^\pi > 0$ and $\delta^x > 0$. In the light of equations (15), (16), (18), (20) and (23), it then follows that NIR shocks lead to $Cov\left(i_t^{TR}, E_t\pi_{t+1}\right) > 0$ and $Cov\left(i_t^{FW}, E_t\pi_{t+1}\right) > 0$. In other words, the policy rules call for tighter monetary policy when expected inflation increases.

Thus, as anticipated in the discussion following equation (11), a positive covariance between the policy rate, i_t^{TR} or i_t^{FW}, and the RER, q_t, is a necessary condition for obtaining a model-based $\widehat{\beta} < 0$. It turns out that such a covariance does lead to $\widehat{\beta} < 0$ but only on a subset of the parameter space, i.e. only for certain values of the coefficients governing the dependence of π_t and x_t on q_t (i.e. ζ^π, ζ^x, θ^π and θ^x). In the next section we derive this subset numerically.

6 Numerical calibration

We analyze each of the three CPI-based rules in a separate subsection, where we proceed as follows. We start by reporting and interpreting the implied slope coefficient of the Fama regression for benchmark values of the model parameters. We obtain such values for the openness and stochastic-processes parameters – a, μ, ρ and υ – by averaging our currency-specific estimates for the 7 economies in our sample (**Table 2**).[16] We complement these with consensus values for the

[15] The statements in this subsection are substantiated in Appendices C and D.

[16] Some remarks are in order. First, values of μ close to unity are in line with the finding of Benigno (2004) and Engel (2000) that the real exchange rates of advanced economies exhibit extremely high persistence. Second, the average value of our $\sqrt{V(\eta)}$ estimates is comparable to that of Mendoza (1995) for G-7 countries: 4.7%. Third, in order to estimate the parameters

structural parameters b, κ and σ (**Table 3**). Then, we conduct sensitivity analysis to determine which parameter values can change substantially the benchmark conclusions. For this analysis, we express model-implied slope coefficients as functions of each model parameter (**Figures 1A** and **1B**), keeping the values of the other parameters at the levels in **Table 3**.[17]

6.1 Strict CPI-inflation targeting (IT)

Even though a strict CPI-inflation targeting rule satisfies the sign condition (recall equation (22)), it does not satisfy the magnitude condition. As reported in **Table 4**, the implied slope coefficient of the Fama regression is virtually zero. This is because the rule generates an extremely volatile policy rate, i_t^{IT}.

To see in what sense the volatility of i_t^{IT} is high and what the implications of this are, recall equation (21). This equation reveals that the lower the product $\kappa\sigma$, the more strongly i_t^{IT} needs to respond to exogenous shocks in order to keep CPI inflation at zero. Concretely, $V\left(i_t^{IT}\right)$ is of the order of $(\kappa\sigma)^{-2}$, which is roughly $(0.008)^{-2}$ for the adopted parameter values. In turn, since the process of the RER, q_t, does not depend on either κ or σ, it follows that $Cov\left(i_t^{IT}, q_t\right)$ is of the order of $(\kappa\sigma)^{-1}$. Thus, the small value of $\kappa\sigma$ translates into a small absolute value of $Cov\left(i_t^{IT}, q_t\right)/V\left(i_t^{IT}\right)$, which leads to a small absolute value of β^{IT} by equation (22).

This conclusion is robust to changing the model's parameters within a reasonable range. The sensitivity results, reported in **Figures 1A** and **1B**, reveal that β^{IT} is virtually zero over the entire explored region of the parameter space.[18]

ρ and $V(\varepsilon)$, we first need to estimate time series of the natural real interest rate, which are not observed directly. We do so on the basis of three-month treasury rates and the corresponding three-month ex-post inflation rates implied by GDP deflators.

[17] Reassuringly, for all the endogenous variables in our model, we derive well-behaved impulse response functions with respect to NIR and RER shocks. These impulse response functions, which we do not report here to conserve space, are available upon request.

[18] Mathematical derivations in support of statements we make regarding the sensitivity analysis under the IT, TR and FW rules are in Appendices B-D.

6.2 CPI-based Taylor rule (*TR*)

Under the parameter values in **Table 3**, the Taylor rule implies a slope coefficient of the Fama regression, β^{TR}, that is negative but with a small absolute value. Even though $|\beta^{TR}| > |\beta^{IT}|$, the former value is still an order of magnitude smaller than what is necessary in order to account for strong empirical UIP violations (compare **Table 4** with **Table 1**). In other words, the Taylor rule TR fails to satisfy the magnitude condition at the end of Section 4.1.

To examine the dependence of β^{TR} on parameter values, we turn to results of the sensitivity analysis (**Figures 1A** and **1B**). The main messages as regards the sign and magnitude conditions are as follows. First, to satisfy the sign condition introduced in Section 4.1, TR needs to be responsive to foreign shocks (i.e. shocks to the RER, q_t), implying that the economy's openness (a) and the policy rule parameters (ϕ_x and ϕ_π) should be above certain levels. Second, the value of β^{TR} gets further into negative territory as stronger mean-reversion increases the predictability of future changes to RER (i.e. as μ decreases). This is because higher predictability of RER changes, Δq_{t+1}, increases the absolute value of $Cov\left(i_t, E_t\Delta q_{t+1}\right) = -\left(1-\mu\right)Cov\left(i_t, q_t\right)$ in equation (11). Third, the value of β^{TR} also gets further into negative territory as RER shocks increase in importance relative to NIR shocks. This is the case when: (i) the variance of RER innovations increases relative to that of NIR innovations, i.e. υ increases; or (ii) lower NIR persistence depresses the unconditional variance of NIR shocks, i.e. ρ decreases. Fourth, to satisfy the sign condition, TR should not be too closely aligned with CPI inflation (π_t), which implies that the strength of the policy transmission mechanism (as captured by σ) and the openness parameter (a) should be below certain levels. Fifth, to satisfy the magnitude condition, the policy rate should not be too responsive to shocks, which implies that the policy rule parameters (ϕ_x and ϕ_π) should be below certain levels. In the rest of this subsection we flesh out some of these statements.

We start with the U-shaped relationship between β^{TR} and the economy's openness, a. If $a = 0$, the economy is closed and RER shocks are inconsequential. In this case, $\beta > 0$ which simply confirms that TR is a well-behaved policy rule, as discussed at the end of Section 4.1. In turn, if $a > 0$, equation (15) implies that a positive shock to q_t raises the current inflation rate, π_t, which induces an upward revision of the policy rate, i_t^{TR}, by the definition of TR in equation (18). This generates $Cov\left(q, i_t\right) > 0$. And, if a increases from low positive levels, the latter covariance drives β^{TR} into negative territory by equation (11). Importantly, however, equation (15) also implies that a higher a raises the sensitivity of CPI inflation to RER shocks, which, all else equal, raises the volatility of CPI inflation. Thus, if a increases beyond a certain point, the policy rate's response to highly volatile inflation raises $Cov\left(E_t\pi_{t+1}, i_t\right) > 0$, so that the implied slope coefficient increases and eventually re-enters positive territory.

Next, we consider the positive relationship between β^{TR} and the strength of the impact of the NIR on inflation and the output gap, as captured by σ. When σ is large, it translates a given NIR volatility directly in a highly volatile output gap (equation (16)), and indirectly in a highly volatile CPI inflation (equation (15)). Thus, similarly to a large a, a large σ delivers $\beta^{TR} > 0$. To understand the implications of a small σ, consider the extreme case of $\sigma = 0$. In this case, the output gap is not hit by any shocks and thus equals its steady-state level, i.e. $x_t = 0$ at all t. On the basis of equation (15), this leads to a closed-form expression for CPI inflation:

$$\pi_t = \frac{a}{1-a}\Delta q_t \qquad \Rightarrow \qquad E_t\pi_{t+1} = -\frac{a\left(1-\mu\right)}{1-a}q_t \tag{24}$$

Since TR forces i_t^{TR} to rise with π_t, expression (24) implies that $Cov\left(i_t^{TR}, q_t\right) > 0$. Then, by equation (11), we obtain $\beta^{TR} < 0$ for $\sigma = 0$. And by the fact that all relationships are continuous in our model, we obtain that $\beta^{TR} < 0$ for sufficiently low but strictly positive σ.

Finally, we dissect the positive relationship between β^{TR} and the RER persistence parameter, μ. As $\mu \to 1$, expected changes in the RER, i.e. $E_t\Delta q_{t+1}$, converge to zero. This means that $Cov\left(i_t^{TR}, E_t\Delta q_{t+1}\right)$ converges to zero and β^{TR} is driven by the responsiveness of the policy rate to expected inflation, $Cov\left(i_t^{TR}, E_t\Delta\pi_{t+1}\right)$. As we have already shown, the latter covariance is positive, implying that β^{TR} moves into positive territory as $\mu \to 1$. As μ decreases from 1, the implied regression slope β^{TR} not only turns negative but increases monotonically in absolute value. To see why, recall that the policy rate co-moves positively with the RER: $Cov\left(i_t^{TR}, q_t\right) > 0$. All else equal, mean reversion in q_t

(i.e. $\mu < 1$) translates $Cov\left(i_t^{TR}, q_t\right) > 0$ into $Cov\left(i_t^{TR}, E_t\Delta q_{t+1}\right) < 0$. And the stronger is this mean reversion, i.e. the lower is μ, the larger is the absolute value of the latter covariance. By equation (6), this explains why lower values of μ push β^{TR} further below 0.

6.3 CPI-based forward-looking rule ($\boldsymbol{FW}$)

For the parameter values in **Table 3**, the forward-looking rule leads to a slope coefficient of the Fama regression, β^{FW}, that is roughly equal to -2 (**Table 4**), in line with strong UIP violations observed in the data (**Table 1**). As in the case of the other two CPI-based rules, we relate this finding to the volatility of the policy rate, $V\left(i_t^{FW}\right)$. FW induces i_t^{FW} to respond to current exogenous shocks only to the extent that they carry information about expected future inflation. And since this information is limited under the high estimated persistence of the exogenous shocks (**Table 2**), the responsiveness of i_t^{FW} to concurrent shocks is low. The resulting low $V\left(i_t^{FW}\right)$ then translates into a high β^{FW} by equation (11).

Turning to the sensitivity analysis, the impact of several parameters on β^{FW} is similar to that on β^{TR}, and for similar reasons. This is the case of: (i) the relative variance of the RER and NIR innovations, υ; the policy coefficients, φ_x and φ_π, which are the forward-looking versions of ϕ_x and ϕ_π; and (iii) the openness parameter a. In addition, high values of the parameters affecting the monetary transmission mechanism, σ, and the persistence of the RER, μ, also have similar effects on β^{FW} and β^{TR}. In the remainder of this subsection, we discuss cases where the results of the sensitivity analysis under FW differ from those under TR (recall **Figures 1A** and **1B**).

First, we explain why a weak transmission from the policy rate to the output gap and inflation – i.e. a low σ – leads to $\beta^{FW} > 0$. This result follows from the inequalities $Cov\left(i_t^{FW}, E_t\pi_{t+1}\right) > 0$ and $Cov\left(i_t^{FW}, q_t\right) < 0$ and equation (11). The first of these inequalities is a direct implication of the definition of FW in equation (20). In turn, when $\sigma = 0$, the second inequality follows from the first and the implication of expression (24) that $Cov\left(E_t\pi_{t+1}, q_t\right) < 0$. By continuity, $Cov\left(i_t^{FW}, q_t\right)$ remains negative for a sufficiently small σ.

The combination of an intermediate σ and a highly persistent but stationary RER – i.e. a high $\mu < 1$ – delivers $\beta^{FW} < 0$. We explain why this is the case in four steps.

1. Starting with $\sigma = 0$, equation (24) and policy rule (20) lead to a policy rate $i_t^{FW} = -\varphi_\pi \frac{a(1-\mu)}{1-a} q_t$.

2. Next, setting $\sigma > 0$ allows i_t^{FW} to affect the inflation rate, π_t. On the basis of equations (15)-(16), step 1 suggests that the impact of i_t^{FW} on π_t introduces a positive dependence of π_t on q_t. In Appendix D we show that this is indeed the case.

3. For a sufficiently persistent q_t – i.e., a sufficiently high μ – the positive dependence of π_t on q_t translates into $Cov\,(q_t, E_t\pi_{t+1}) > 0$ and, by policy rule (20), into $Cov\left(i_t^{FW}, q_t\right) > 0$.

4. $Cov\left(i_t^{FW}, q_t\right) > 0$ leads to $Cov\left(i_t^{FW}, E_t\Delta q_{t+1}\right) = -\left(1-\mu\right) Cov\left(i_t^{FW}, q_t\right) < 0$. However, given (11) and $Cov\left(i_t^{FW}, E_t\Delta\pi_{t+1}\right) > 0$, this leads to $\beta^{FW} < 0$ only if μ is not too close to 1 because $\lim\limits_{\mu\to 1} Cov\left(i_t^{FW}, E_t\Delta q_{t+1}\right) = 0^-$.

Finally, high mean reversion of the RER, i.e. a sufficiently low μ, leads to $\beta^{FW} > 0$.[19] To see why, note that, for the reasons outlined in step 2 above, the benchmark value of σ leads to $Cov\,(\pi_t, q_t) > 0$. When q_t is expected to revert strongly to the mean, the last inequality translates into $Cov\,(E\pi_{t+1}, q_t) < 0$ and, by the definition of FW, into $Cov\left(i_t^{FW}, q_t\right) < 0$. And since $Cov\,(E\pi_{t+1}, q_t) > 0$, we obtain $\beta^{FW} > 0$ by equation (11).

7 Conclusion

The question we asked in this paper is whether systematic monetary policy can be a driver of the strong empirical violations of UIP that were first reported by Fama (1984) and have puzzled economists since then. In other words, can monetary policy explain why researchers often estimate negative slope coefficients when regressing the short-term nominal exchange rate change on the corresponding

[19] Note that the mean reversion of a stationary random variable increases with the forecast horizon. Concretely, if $\mu = 0.95$ at a one-quarter horizon, then it would be equal to $0.95^{20} = 0.36$ at a five-year horizon. As implied by Figure 1B, such a value of μ would lead to $\beta^{FW} > 0$, consistent with findings in Chinn and Meredith (2004, 2005).

interest differential? Our main *empirical* contribution is to provide evidence of the relevance of this question. Concretely, we show that a change in the monetary policy regime coincides with a statistically significant change in the degrees of UIP violation by the exchange rates of the Canadian dollar, British pound and Swedish krona with the US dollar.

The *theoretical* contribution of the paper lies in explaining a *causal* link from rule-based monetary policy to UIP violations. We study the joint determination of exchange and interest rates in a reduced-form New Keynesian model with rational expectations, in which monetary policy responds to exogenous time variation in the natural rate of interest and the real exchange rate. Given our assumption of perfect exchange rate pass-through, shocks to the real exchange rate can be thought of as terms-of-trade shocks. Under each of the six monetary policy rules we consider, we derive a closed-form expression of the degree of UIP violation as a function of model parameters. Then, we show that a negative *sign* of the slope coefficient in the Fama regression cannot arise from policy rules responding only to domestic inflation. And while each of the CPI-based rules can deliver a negative slope coefficient, only the forward-looking CPI-based rule can lead to *magnitudes* of this coefficient that are consistent with strong UIP violations observed in the data.

Our analysis reveals that the degree of UIP violations can be quite sensitive to the persistence and volatility of the real exchange rate and the natural interest rate. By extension, it then follows that accurate estimates of the processes of these two variables are a prerequisite for accurate assessments of the excess returns on investments in a particular currency. What estimates are possible in real time and whether such estimates can support profitable investment strategies are open questions that could lead to natural extensions of the work of Bacchetta and van Wincoop (2007, 2010) on arbitrage opportunities in foreign exchange markets. We leave them to future research.

Appendix A

In this appendix we show that the *TRd* and *FWd* policy rules non-negative regression slopes: $\beta^{TRd} \geq 0$ and $\beta^{FWd} \geq 0$. To lighten the notation, we do not flag the policy rule in a superscript wherever this does not create confusion.

A.1: Domestic-inflation Taylor-type rules

To derive explicit expressions for domestic inflation and the output gap, we apply the method of undetermined coefficients to two alternative expressions for the ex-ante real interest rate, $r_t \equiv i_t - E_t \pi_{t+1}^D$. We start by considering the policy rule in equation (17) and conjecturing that domestic inflation is a linear function of the natural rate, r_t^n:

$$\pi_t^D = \delta r_t^n \Rightarrow \; E_t \pi_{t+1}^D = \delta \rho r_t^n \tag{A.1}$$

Then, we note that combining equation (12) with expression (A.1) leads to:

$$x_t = \frac{\delta}{\kappa}(1 - b\rho)\; r_t^n \Rightarrow E_t x_{t+1} = \frac{\delta}{\kappa}\rho(1 - b\rho)\; r_t^n$$

Substituting these expressions for x_t and $E_t x_{t+1}$ in equation (13) delivers the first expression for the real rate:

$$r_t = \frac{\delta(1 - b\rho)}{\kappa\sigma}\left(\rho - 1 + \frac{\kappa\sigma}{\delta(1 - b\rho)}\right)\; r_t^n \tag{A.2}$$

For the second expression, we apply the policy rule (17) to the definition $r_t \equiv i_t - E_t \pi_{t+1}^D$:

$$r_t = \delta\left(\phi_\pi - \rho + \phi_x \frac{1 - b\rho}{\kappa}\right)\; r_t^n \tag{A.3}$$

Matching the coefficients of r_t^n in (A.2) and (A.3) yields:

$$\delta = \frac{\kappa\sigma}{(1 - b\rho)(1 - \rho) + \kappa\sigma\left(\phi_\pi - \rho + \frac{1}{\kappa}\phi_x (1 - b\rho)\right)} > 0 \tag{A.4}$$

Then, we note that the sign of β^{TRd} is equal to that of $Cov\,(i_t, \pi_{t+1})$. And since π_t is a linear function of π_t^D and q_t, and i_t is independent of q_t, it follows that $Cov\,(i_t, \pi_{t+1}) = Cov\left(i_t, \pi_{t+1}^D\right)$. Finally, since the sign of $Cov\left(i_t, \pi_{t+1}^D\right)$ is that of

δ, which is positive, we obtain $\beta^{TRd} > 0$.

A.2: Domestic-inflation forward-looking rules

In this appendix we mimic Appendix A.1 in the context of the policy rule in equation (19).

Given equations (12)-(13), the first expression for r_t coincides with (A.2) above. The second expression is the forward-looking analog of (A.3):

$$r_t = \delta\rho \left[(\varphi_\pi - 1) + \varphi_x \kappa^{-1}(1 - b\rho) \right] \; r_t^n \tag{A.5}$$

Matching the coefficients of r_t^n, in (A.2) and (A.5), we obtain:

$$\delta = \frac{\kappa\sigma}{(1 - b\rho)(1 - \rho) + \kappa\sigma\rho \left(\varphi_\pi - 1 + \frac{1}{\kappa}\varphi_x (1 - b\rho) \right)} > 0 \tag{A.6}$$

For reasons similar to those stated in *Appendix A.1*, the last inequality implies $\beta^{FWd} > 0$.

Appendix B: Strict CPI inflation-targeting (*IT*)

In this appendix we rewrite expression (22) explicitly in terms of model parameters. We start by noting that the *IT* rule implies $\pi_t = E_t\pi_{t+1} = 0$, and thus $e_t = q_t$ and $E_t\Delta e_{t+1} = E_t\Delta q_{t+1}$. Then, making use of equation (15) we obtain:

$$\begin{aligned} x_t &= -\frac{a}{1-a}\frac{1}{\kappa}\left(\Delta q_t + b(1-\mu)q_t\right) \\ E_t x_{t+1} &= \frac{a}{1-a}\frac{(1-\mu)(1-b\mu)}{\kappa} q_t \end{aligned} \tag{B.1}$$

In turn, noting that expected domestic inflation now equals

$$E_t\pi_{t+1}^D = E_t\Delta e_{t+1} - \frac{E_t\Delta q_{t+1}}{1-a} = \frac{a}{1-a}(1-\mu)\, q_t \tag{B.2}$$

we use the IS equation (16) to derive the policy rate as a function of exogenous variables:

$$i_t^{IT} = r_t^n + \left(\frac{a}{1-a}(1-\mu) + \frac{a}{\kappa\sigma(1-a)}\left(b(1-\mu)^2 + 2 - \mu\right)\right) q_t - \frac{a}{\kappa\sigma(1-a)} q_{t-1} \tag{B.3}$$

The implied Fama slope coefficient is then defined as:

$$\beta^{IT} = \frac{Cov\left(i_t^{IT}, E_t\Delta e_{t+1}\right)}{V\left(i_t^{IT}\right)} = -\frac{(1-\mu)\,Cov\left(i_t^{IT}, q_t\right)}{V\left(i_t^{IT}\right)} \tag{B.4}$$

$$Cov\left(i_t^{IT}, q_t\right) = \frac{V(\eta)}{1+\mu}\frac{a}{1-a}\left(1 + \frac{b(1-\mu)+2}{\kappa\sigma}\right) > 0 \tag{B.5}$$

$$V\left(i_t^{IT}\right) = \frac{V(\varepsilon)}{1-\rho^2} + \frac{a^2\begin{pmatrix}\kappa\sigma(1-\mu)(2b(1-\mu)+\kappa\sigma+4) \\ +b^2(1-\mu)^3 + 4b(1-\mu)^2 + 5 - 3\mu\end{pmatrix}}{\kappa^2\sigma^2(1-a)^2(1+\mu)} V(\eta) \tag{B.6}$$

Finally, dividing (B.5) by (B.6) and using $\upsilon \equiv \frac{V(\eta)}{V(\varepsilon)}$ yields:

$$\beta^{IT} = -\frac{\frac{1-\mu}{1+\mu}\frac{a}{1-a}\left(1+\frac{b(1-\mu)+2}{\kappa\sigma}\right)\upsilon}{\frac{1}{1-\rho^2} + \frac{a^2\begin{pmatrix}\kappa\sigma(1-\mu)(2b(1-\mu)+\kappa\sigma+4) \\ +b^2(1-\mu)^3 + 4b(1-\mu)^2 + 5 - 3\mu\end{pmatrix}}{(1-a)^2\kappa^2\sigma^2(1+\mu)}\upsilon} < 0 \tag{B.7}$$

Appendix C: Taylor-type policy rules (*TR*)

In this appendix we derive the coefficient vector $\{\delta^\pi, \zeta^\pi, \theta^\pi, \delta^x, \zeta^x, \theta^x\}$ determining the linear mapping from the exogenous variables (r^n and q) to the inflation rate and the output gap under TR. To employ the method of undetermined coefficients, we start with the guess

$$\pi_t = \delta^\pi r_t^n + \zeta^\pi q_t + \theta^\pi q_{t-1} \Rightarrow\ E_t\pi_{t+1} = \delta^\pi \rho r_t^n + \left(\mu\zeta^\pi + \theta^\pi\right) q_t \tag{C.1}$$

which implies, via equations (9)-(10) and (15), that

$$x_t = \delta^x r_t^n + \zeta^x q_t + \theta^x q_{t-1} \Rightarrow \; E_t x_{t+1} = \delta^x \rho r_t^n + (\mu\zeta^x + \theta^x)\; q_t \qquad \text{(C.2)}$$

where

$$\begin{aligned} \delta^x &\equiv \delta^\pi \kappa^{-1}(1-b\rho),\ \theta^x \equiv \kappa^{-1}\left(\theta^\pi + \frac{a}{1-a}\right) \text{ and} \qquad \text{(C.3)} \\ \zeta^x &\equiv \frac{1}{\kappa(1-a)}\{(1-a)\zeta^\pi - a[1+b(1-\mu)] - b(1-a)(\mu\zeta^\pi + \theta^\pi)\} \end{aligned}$$

By equations (18), (C.1) and (C.2) we obtain the policy rate under TR:

$$i_t^{TR} = (\phi_\pi \delta^\pi + \phi_x \delta^x)\, r_t^n + (\phi_\pi \zeta^\pi + \phi_x \zeta^x)\, q_t + (\phi_\pi \theta^\pi + \phi_x \theta^x)\, q_{t-1} \qquad \text{(C.4)}$$

Then, by equations (16), (C.1) and (C.2), we obtain an alternative expression for the policy rate:

$$\begin{aligned} i_t^{TR,FW} &= \frac{E_t \Delta x_{t+1}}{\sigma} + E_t \pi_{t+1}^D + r_t^n \\ &= \frac{E_t \Delta x_{t+1}}{\sigma} + E_t\left(\pi_{t+1} - \frac{a}{1-a}\Delta q_{t+1}\right) + r_t^n \\ &= \sigma^{-1}[-(1-\rho)\,\delta^x r_t^n + (\theta^x - (1-\mu)\,\zeta^x)\; q_t - \theta^x q_{t-1}] \qquad \text{(C.5)} \\ &\quad +(\delta^\pi \rho + 1) r_t^n + (\mu\zeta^\pi + \theta^\pi)\; q_t + \frac{a}{1-a}(1-\mu)\, q_t \end{aligned}$$

Note that the functional form of this expression is the same under both the *TR* rule and the *FW* rule (considered in *Appendix D*). That said, the values of some parameters governing the policy rate – i.e. δ^π, ζ^π and θ^π – can change with the policy rule. These parameters depend on the policy response coefficients, which change from ϕ_π and ϕ_x under TR to φ_π and φ_x under FW. The same comment applies to all expressions in *Appendices C* and *D* that incorporate δ^π, ζ^π or θ^π.

Matching coefficients in (C.4) and (C.5):

- r_t^n terms:

$$\delta^\pi = \left[\phi_\pi - \rho + \frac{1-b\rho}{\kappa}\left(\phi_x + \frac{1-\rho}{\sigma}\right)\right]^{-1} \tag{C.6}$$
$$\delta^\pi > 0 \text{ because } \phi_\pi > 1,\ \rho \in (0,1),\ b \in (0,1)$$

- q_{t-1} terms:

$$\theta^\pi = -\frac{a}{1-a}\frac{1+\phi_x\sigma}{1+\sigma(\phi_x+\kappa\phi_\pi)} \tag{C.7}$$
$$\theta^\pi < 0 \text{ because } a \in (0,1)$$

- q_t terms:

$$\zeta^\pi = \frac{\left(\phi_x + \frac{1-\mu}{\sigma}\right)\left(\frac{b\theta^\pi}{\kappa} + \frac{a}{1-a}\frac{1+b(1-\mu)}{\kappa}\right) + \theta^\pi + \frac{a(1-\mu)}{1-a} + \frac{\frac{a}{1-a}+\theta^\pi}{\kappa\sigma}}{\phi_\pi - \mu + \left(\phi_x + \frac{1-\mu}{\sigma}\right)\frac{1-b\mu}{\kappa}} \tag{C.8}$$
$$= \frac{a}{1-a}\frac{\left(\begin{array}{c}\frac{1}{\kappa}\frac{(1+\sigma\phi_x)((\phi_\pi-1)\kappa+(1-b)\phi_x)}{\sigma\phi_x+\kappa\sigma\phi_\pi+1} \\ +(1-\mu)\left(\frac{b(1-\mu)}{\kappa\sigma} + \phi_x\frac{b}{\kappa} + 1 + \frac{(1+\sigma\phi_x)(1-b)+\kappa\sigma\phi_\pi}{\kappa\sigma(\sigma\phi_x+\kappa\sigma\phi_\pi+1)}\right)\end{array}\right)}{\phi_\pi - \mu + \left(\phi_x + \frac{1-\mu}{\sigma}\right)\frac{1-b\mu}{\kappa}}$$
$$\zeta^\pi > 0 \text{ because } \phi_\pi > 1,\ \mu \in (0,1),\ b\,(0,1),\ a \in (0,1)$$

Finally, equations (5), (6), (9), (10) and (C.1)-(C.8) yield:

$$\beta^{TR} = \frac{\begin{array}{c}\frac{\delta^\pi\rho(\phi_\pi\delta^\pi+\phi_x\delta^x)}{1-\rho^2} \\ +(\mu(\zeta^\pi+1)+\theta^\pi-1)\left(\begin{array}{c}\frac{\mu(\phi_\pi(\zeta^\pi\mu+\theta^\pi)+\phi_x(\zeta^x\mu+\theta^x))}{1-\mu^2} \\ +(\phi_\pi\zeta^\pi+\phi_x\zeta^x)\end{array}\right)\upsilon\end{array}}{\begin{array}{c}\frac{(\phi_\pi\delta^\pi+\phi_x\delta^x)^2}{1-\rho^2} \\ +\left(\frac{(\phi_\pi(\zeta^\pi\mu+\theta^\pi)+\phi_x(\zeta^x\mu+\theta^x))^2}{1-\mu^2} + (\phi_\pi\zeta^\pi+\phi_x\zeta^x)^2\right)\upsilon\end{array}} \tag{C.9}$$

Since $\delta^\pi > 0$ and $\delta^x > 0$, the last expression reveals directly that $\beta^{TR} > 0$ for small enough υ. In addition, it can be shown that there exists $\mu^h > 0$ such that $\beta^{TR} < 0$ for $\mu \in (0, \mu^h)$.

Appendix D: Forward-looking policy rules (*FW*)

In determining the coefficient vector $\{\delta^{\pi}, \zeta^{\pi}, \theta^{\pi}, \delta^{x}, \zeta^{x}, \theta^{x}\}$ under the FW rule, we parallel Appendix C. By equations (20), (C.1) and (C.2) we obtain:

$$i_t^{FW} = (\varphi_\pi \delta^\pi + \varphi_x \delta^x)\, \rho r_t^n + [\varphi_\pi(\mu\zeta^\pi + \theta^\pi) + \varphi_x(\mu\zeta^x + \theta^x)]q_t \tag{D.1}$$

Matching coefficients in (C.5) and (D.1) yields

- r_t^n terms:

$$\begin{aligned} \delta^\pi &= (\kappa\sigma)[\rho\kappa\sigma\,(\varphi_\pi - 1) + \varphi_x \sigma\rho\,(1 - b\rho) + (1-\rho)\,(1 - b\rho)]^{-1} \\ \text{where } \delta^\pi &> 0 \text{ because } \varphi_\pi > 1,\, b \in (0,1)\,,\, \rho\,(0,1)\,,\, a \in (0,1) \end{aligned} \tag{D.2}$$

- q_{t-1} terms:

$$\theta^\pi = -\frac{a}{1-a} < 0 \tag{D.3}$$

- q_t terms:

$$\begin{aligned} \zeta^\pi &= \frac{a}{1-a}\frac{(\varphi_\pi - \mu)\,\kappa\sigma + (1-b\mu)\,(\varphi_x\mu\sigma + 1 - \mu)}{(\varphi_\pi - 1)\,\mu\kappa\sigma + (1-b\mu)\,(\varphi_x\mu\sigma + 1 - \mu)} \\ \text{where } \zeta^\pi &> 0 \text{ because } \varphi_\pi > 1,\, \mu \in (0,1)\,,\, b \in (0,1)\,,\, a \in (0,1) \end{aligned} \tag{D.4}$$

The corresponding values of δ^x, ζ^x and θ^x under the FW rule are then obtained by combining (D.3) and (D.4) with the definitions in (C.3).

Finally, just as in *Appendix C*, we obtain

$$\beta^{FW} = \frac{\frac{\delta^\pi\rho^2(\varphi_\pi\delta^\pi + \varphi_x\delta^x)}{1-\rho^2} + \frac{(\mu(\zeta^\pi+1)+\theta^\pi-1)(\varphi_\pi(\mu\zeta^\pi+\theta^\pi)+\varphi_x(\mu\zeta^x+\theta^x)\,)\upsilon}{1-\mu^2}}{\frac{\rho^2(\varphi_\pi\delta^\pi+\varphi_x\delta^x)^2}{1-\rho^2} + \frac{(\varphi_\pi(\mu\zeta^\pi+\theta^\pi)+\varphi_x(\mu\zeta^x+\theta^x)\,)^2\upsilon}{1-\mu^2}} \tag{D.5}$$

Since $\delta^\pi > 0$ and $\delta^x > 0$, the last expression reveals directly that $\beta^{FW} > 0$ for small enough υ. In addition, equations (23), (D.3) and (D.4) imply that $Cov\,(q_t, E_t\pi_{t+1}) > 0$ if and only if $\frac{\mu}{1-b\mu} - \frac{1-\mu}{\kappa\sigma} > 0$. Since the left-hand side of the latter inequality increases strictly in μ, there is a unique threshold μ^l such that $d\left(\mu^l\right)/d\,(\kappa\sigma) < 0$ and $Cov\,(q_t, E_t\pi_{t+1}) > 0$ for $\mu \in \left(\mu^l, 1\right)$.

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Table 1: Fama regression[20]						
	before inflation targeting			*inflation targeting (or whole sample)*		
	$\widehat{\alpha}$	$\widehat{\beta}$	$\overline{R}^2$	$\widehat{\alpha}$	$\widehat{\beta}$	$\overline{R}^2$
Australia (AUD)				0.005	-0.67^{***}	0.01
				(0.007)	(0.58)	
				1984:Q1-2008:Q3		
Canada (CAD)	0.005	-0.49^{*}	0.01	-0.006	-2.36^{***}	0.04
	(0.004)	(0.75)		(0.004)	(1.22)	
	1977:Q4-1994:Q2			*1994:Q3-2008:Q3*		
Switzerland (CHF)				-0.014	-0.99^{***}	0.01
				(0.007)	(0.59)	
				1977:Q4-2008:Q3		
United Kingdom (GBP)	0.015	-3.22^{***}	0.14	-0.001	2.87	0.07
	(0.008)	(0.93)		(0.008)	(2.35)	
	1977:Q4-1992:Q2			*1992:Q3-2008:Q3*		
Sweden (SEK)	0.002	0.86	0.002	-0.004	-2.98^{***}	0.10
	(0.012)	(1.63)		(0.006)	(1.18)	
	1977:Q4-1993:Q4			*1994:Q1-2008:Q3*		
New Zealand (NZD)				0.003	-0.98	0.01
				(0.015)	(2.08)	
				1992:Q3-2008:Q3		
Euro area (euro)				-0.012	-3.21^{**}	0.05
				(0.007)	(1.87)	
				1999:Q2-2008:Q3		

[20] Estimates of equation (2), based on quarterly spot and three-month forward exchange rates with the US dollar (sources: Bloomberg and BIS). Newey-West robust standard errors are in parentheses. Sample periods are in italics. For the first five currencies, a structural break is tested for in an eight-quarter window around the officially announced switch to explicit inflation targeting by the corresponding central bank (see footnote 7). If a Chow test identifies such a break (at the 95% confidence level), equation (2) is estimated over two non-overlapping time periods. Because of shorter data series, no break is allowed for in the case of the last two currencies. Superscripts *, ** and *** denote a slope coefficient that is different from 1 at the 90%, 95% and 99% confidence level, respectively.

Table 1.1: Dependent variable: CPI inflation[21]						
	before inflation targeting			*inflation targeting (or whole sample)*		
	$\widehat{\alpha}_1$	$\widehat{\beta}_1$	$\overline{R}^2$	$\widehat{\alpha}_1$	$\widehat{\beta}_1$	$\overline{R}^2$
Australia (AUD)				−0.003	0.73	0.01
				(0.005)	(0.50)	
					1984:Q1-2008:Q3	
Canada (CAD)	−0.002	0.77	0.01	−0.002	−0.13	0.00
	(0.003)	(0.54)		(0.003)	(0.90)	
		1977:Q4-1994:Q2			*1994:Q3-2008:Q3*	
Switzerland (CHF)				0.000	0.57	0.00
				(0.005)	(0.47)	
					1977:Q4-2008:Q3	
United Kingdom (GBP)	0.002	0.31	0.00	−0.005	1.14	0.00
	(0.007)	(0.72)		(0.005)	(1.17)	
		1977:Q4-1992:Q2			*1992:Q3-2008:Q3*	
Sweden (SEK)	−0.004	1.22	0.06	−0.003	0.35	0.00
	(0.006)	(0.50)		(0.005)	(0.93)	
		1977:Q4-1993:Q4			*1994:Q1-2008:Q3*	
New Zealand (NZD)				−0.007	0.89	0.00
				(0.007)	(0.84)	
					1992:Q3-2008:Q3	
Euro area (euro)				−0.001	0.88	0.00
				(0.006)	(1.54)	
					1999:Q2-2008:Q3	

[21] Estimates of equation (7), based on quarterly CPI inflation and three-month forward exchange rates with the US dollar (sources: Bloomberg and BIS). CPI inflation is estimated with nominal exchange rates and the corresponding real exchange rates, as implied by equation (5). Newey-West robust standard errors are in parentheses. Sample periods, in italics, are the same as in Table 1.

Table 1.2: Dependent variable: real exchange rate depreciation[22]						
	before inflation targeting			*inflation targeting (or whole sample)*		
	$\widehat{\alpha}_2$	$\widehat{\beta}_2$	$\overline{R}^2$	$\widehat{\alpha}_2$	$\widehat{\beta}_2$	$\overline{R}^2$
Australia (AUD)				0.008	-1.40^{***}	0.05
				(0.006)	(0.52)	
				1984:Q1-2008:Q3		
Canada (CAD)	0.008	-1.26^{***}	0.06	-0.004	-2.23^{***}	0.04
	(0.003)	(0.42)		(0.004)	(1.18)	
	1977:Q4-1994:Q2			*1994:Q3-2008:Q3*		
Switzerland (CHF)				-0.014	-1.56^{***}	0.05
				(0.007)	(0.59)	
				1977:Q4-2008:Q3		
United Kingdom (GBP)	0.014	-3.53^{***}	0.21	-0.004	1.74	0.01
	(0.007)	(0.90)		(0.008)	(2.31)	
	1977:Q4-1992:Q2			*1992:Q3-2008:Q3*		
Sweden (SEK)	0.006	-0.26	0.00	-0.002	-3.33^{***}	0.17
	(0.009)	(1.33)		(0.005)	(1.10)	
	1977:Q4-1993:Q4			*1994:Q1-2008:Q3*		
New Zealand (NZD)				0.01	-1.36^{**}	0.01
				(0.009)	(1.00)	
				1992:Q3-2008:Q3		
Euro area (euro)				-0.01	-4.09^{***}	0.14
				(0.006)	(1.53)	
				1999:Q2-2008:Q3		

[22] Estimates of equation (8), based on quarterly real exchange rates and three-month forward exchange rates with the US dollar (sources: Bloomberg and BIS). Newey-West robust standard errors are in parentheses. Sample periods, in italics, are the same as in Table 1. Superscripts ** and *** denote a slope coefficient that is different from 1 at the 95% and 99% confidence level, respectively.

Table 2: Country characteristics[23]						
	$\widehat{a}$	$\widehat{\mu}$	$\sqrt{\widehat{V}(\eta)}$	$\widehat{\rho}$	$\sqrt{\widehat{V}(\varepsilon)}$	$\widehat{\upsilon} \equiv \frac{\widehat{V}(\eta)}{\widehat{V}(\varepsilon)}$
Australia	0.22	0.96	0.043	0.69	0.029	2.20
Canada	0.36	0.98	0.024	0.67	0.028	0.68
Switzerland	0.41	0.94	0.050	0.76	0.012	17.4
Great Britain	0.29	0.92	0.044	0.15	0.028	2.47
Sweden	0.39	0.96	0.047	0.42	0.032	2.16
New Zealand	0.21	0.96	0.051	0.15	0.047	1.18
Euro area	0.31	0.95	0.030	0.81	0.012	6.25
Average	*0.31*	*0.95*	*0.041*	*0.52*	*0.027*	*2.34*

Table 3: Benchmark calibration			
	parameter	*value*	*based on*
coefficient of $E_t(\pi_{t+1})$	b	0.968	Woodford (2003)
sensitivity of π to x	κ	0.0157	Woodford (2003)
sensitivity of x to i	σ	0.5	Woodford (2003)
openness	a	0.31	Table 2
q persistence	μ	0.95	Table 2
r^n persistence	ρ	0.52	Table 2
relative volatility	υ	2.34	Table 2
policy parameters (*TR* and *FW* only)	$\phi_\pi = \varphi_\pi$ $\phi_x = \varphi_x$	1.5 0.025	

[23] Sources: Datastream, IMF, OECD, BIS. The estimate of openness to trade ($\hat{a}$) is equal to the average quarterly imports-to-GDP ratio from 1980:Q1 to 2008:Q3, in the particular country or area. The estimates of the parameters of the AR(1) processes followed by the real exchange rates with the US dollar ($\widehat{\mu}$ and $\widehat{V}(\eta)$) and the natural real interest rates ($\widehat{\rho}$ and $\widehat{V}(\varepsilon)$) are based on equations (9) and (10). In each case, the underlying samples comprise quarterly data from 1984:Q1 to 2008:Q3. The natural real interest rate is estimated with three-month treasury rates and the three-month ex-post inflation rates implied by GDP deflators.

Table 4: Model-implied slope coefficients			
	Strict CPI	*TR*	*FW*
$Cov\left(E_t \pi_{t+1}, i_t\right)$	0	0.09 bp	0.02 bp
$+Cov\left(E_t \Delta q_{t+1}, i_t\right)$	-51 bp	-0.70 bp	-0.12 bp
$= Cov\left(E_t \Delta e_{t+1}, i_t\right)$	-51 bp	-0.61 bp	-0.10 bp
$V\left(i_t\right)$	61881 bp	8.56 bp	0.05 bp
$\widehat{\beta} = \frac{Cov(E_t \Delta e_{t+1}, i_t)}{V(i_t)}$	-0.00	-0.07	-2.09

Figure 1A: Sensitivity Analysis

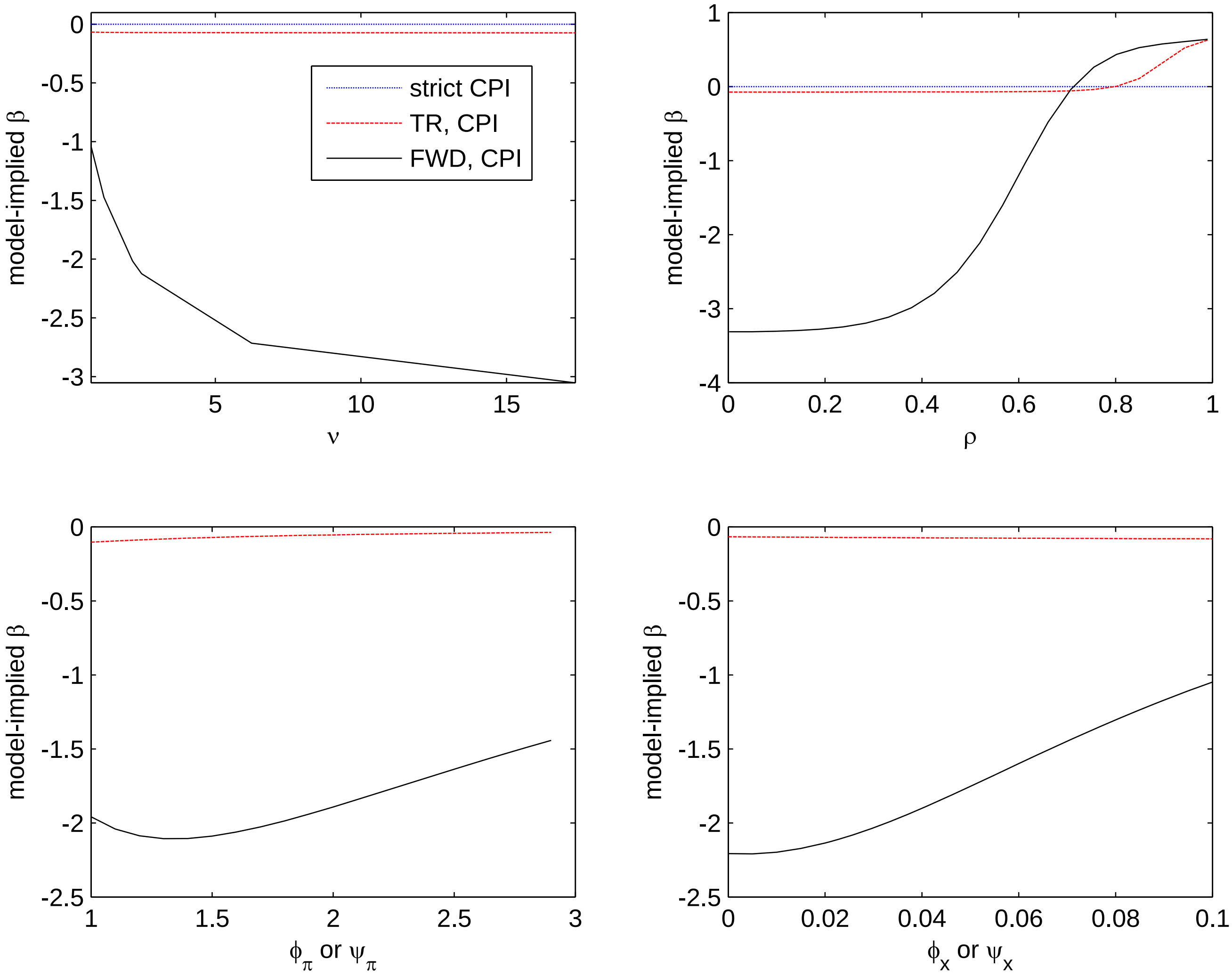

Figure 1B: Sensitivity Analysis