For Rich or for Poor: When does Uncovered Interest Parity Hold?

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Abstract

We present a model that simultaneously explains why uncovered interest parity holds for some pairs of countries and not for others. The negative slope in the Fama regression arises when monetary volatility is low and the precautionary savings motive dominates the intertemporal substitution motive. When monetary volatility is high, the Fama slope is positive in line with uncovered interest parity. The model is simulated using the artificial economy methodology for 42 currencies. We conclude that, given the predominance of precautionary savings, the degree of monetary volatility explains whether or not uncovered interest parity holds.

Keywords: Monetary volatility; Uncovered interest parity; Forward bias puzzle; Habit persistence

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1. Introduction.

Eichenbaum (2010) investigates equally weighted carry trade portfolios for 20 portfolios over a thirty year period and finds consistent excess dollar returns. A ‘carry trade portfolio’ is a strategy of borrowing in the currency of the low interest currency and depositing the proceeds in the high interest currency, taking an open position to nominal exchange rate risk. He concludes that this vividly demonstrates the pervasiveness of the forward ‘bias’ puzzle. The contribution of this study is to specify and test a theory which provides for when the forward bias does and does not hold.

The literature that studies this puzzle or equivalently the failure of uncovered interest parity (UIP) is enormous. The classic survey is Engel (1996) and Sarno (2005) contains a good update. There is a multitude of relatively recent theoretical papers which explain the failure of uncovered interest parity. There are behavioural explanations (Burnside, Eichenbaum and Rebelo, 2009, Fisher, 2006 and Gourinchas and Tornell, 2004); rational inattention is offered by Bacchetta and Van Wincoop (2010); institutional features are emphasised by Carlson and Osler (2005); and Alvarez, Atkinson and Kehoe (2009) explain the forward bias by permitting a time varying degree of asset market participation.

Bansal and Shaliastovich (2009) and Backus, Gavazzoni, Telmer and Zin (2010) show how a model of the discount factor based on Epstein-Zin preferences can explain the uncovered interest parity puzzle. In a model related to the one explored in this paper, Verdelhan (2010) explains the forward bias puzzle using Campbell and Cochrane (1999) preferences in a non-monetary economy with trade costs. Engel (2011) confirms that there is a real analogue to the forward bias problem in the data. Moore and Roche (2010) show that the Verdelhan (2010) result can be obtained as a special case of our analysis, without the contrivance of trade costs.

A different line of empirical research has suggested that the forward bias problem does not always occur. Our main focus is on the claim that the forward bias problem is a feature which arises between developed countries and is much less likely to arise between emerging and developed countries. This was first suggested by Bansal and Dahlquist (2000). Frankel and Poonawala (2009), Chinn (2006) as well as Ito and Chinn (2007) find similar results. Flood and Rose (2002) buck the trend by finding no significant difference in the success of UIP between rich and poor countries. Most interestingly, they find that UIP works systematically better for 'crisis' countries. This is the starting point for this paper.

Moore and Roche (2002, 2008, 2010) have extended Campbell and Cochrane (1999) preferences to both a monetary and an international setting. In the standard model, when the domestic interest rate is relatively low, the exchange rate is expected to appreciate. The forward bias in our model arises because the preference specification provides two motives for savings. The first is the conventional desire to smooth consumption intertemporally. The less familiar motive is a precautionary savings effect, which is dominant in our calibration of the model. When times are relatively bad for the owner of the domestic endowment, the 'surplus consumption ratio' for the domestic good is low in relation to that for the foreign good, the time-varying risk aversion measured in the domestic good is relatively high and the own interest rate is relatively low. Consequently the holder of the domestic bond does not need to be compensated by as much of an expected appreciation and indeed may be content with an expected depreciation. The latter case, which arises easily in our model, is why the forward bias can occur.

The above argument is apparently so compelling that it begs the question as to why UIP would ever hold at all. The answer to this is that the argument in its extreme form applies to a world without nominal magnitudes. Moore and Roche (2010) show that when these are built into agents' optimising problems, the above result can be reversed. An increase in monetary volatility requires that nominal interest rates be raised to compensate the holder of nominal bonds. Note how precise this point is: we are not considering first moments such as money growth or inflation but a second moment: the variance of money growth. Whether or not UIP holds is a balance between the extent of monetary volatility and the dominance of the real precautionary savings motive.

It is easy to see how the failure or otherwise of UIP might follow developed/emerging economy lines. Emerging economies are far more likely to be characterised by volatile monetary regimes and therefore UIP is more likely to hold. The opposite is obviously true of developed countries.

The plan of the paper is as follows. In the next section, the model is outlined. Section 3 is the substantial contribution of the paper in which the model is calibrated and the results are reported. Section 4 makes some concluding remarks.

2. The Model

2.1 Basic Framework

The extension of Campbell and Cochrane (1999) to an international and monetary setting is fully discussed in Moore and Roche (2010). To facilitate the reader, the basic model is briefly summarised in this sub-section. A two-country, two-good, two-money representative agent model is developed similar to that of Lucas (1982). The markets for state-contingent money claims are made complete by having one period contingent nominal bonds as in Chari, Kehoe and McGrattan (2002).

The household in country 1 is assumed to choose consumption of goods and services of country j, C_{1t}^{j}, to purchase state contingent nominal bonds $B^{1}(\tau_{t+1})$ (each of which pays off one unit of home currency in state τ_{t+1}) and $B^{2}(\tau_{t+1})$ (each of which pays off one unit of foreign currency in state τ_{t+1}) to maximize the following expected discounted lifetime utility function[1]:

$$E_t\left(\sum_{t=0}^{\infty}\beta^t\left\{\frac{(C_{1t}^1-H_{1t}^1)^{(1-\gamma)}}{1-\gamma}+\frac{(C_{1t}^2-H_{1t}^2)^{(1-\gamma)}}{1-\gamma}\right\}\right) \tag{1}$$

subject to the equation of motion for wealth in country 1, W_{t+1}^{1}:

$$W_{t+1}^1=\sum_{\tau_{t+1}\in \mathrm{T}}B^1(\tau_{t+1})+S_{t+1}\sum_{\tau_{t+1}\in \mathrm{T}}B^2(\tau_{t+1})+P_t^1Y_t^1 \tag{2}$$

and the wealth constraint:

$$W_t^1=P_t^1C_{1t}^1+S_tP_t^2C_{1t}^2+\sum_{\tau_{t+1}\in \mathrm{T}}Q^1(\tau_{t+1},\tau_t)B^1(\tau_{t+1})+S_t\sum_{\tau_{t+1}\in \mathrm{T}}Q^2(\tau_{t+1},\tau_t)B^2(\tau_{t+1}) \tag{3}$$

In (1) β is the discount factor, γ is a curvature parameter and H_{it}^{j} as the subsistence consumption (or habit) of goods and services of country *j* by the household of country *i*. Following Ravn, Schmitt-Grohe, and Uribe (2007) and Moore and Roche (2002, 2008. 2010) we assume that habits are goods rather than country specific. In (2) T is the set of all possible states, S_t is the level of the spot exchange rate, P_t^{j} is the price of good *j*, and Y_t^{j} is the

[1] The superscript denotes country of origin of the good. Uppercase letters denote variables in levels; lowercase letters denote variables in log levels, including growth and interest rates. Greek letters without time subscripts denote parameters. Bars over variables denote steady states. With the exception of the money growth process discussed below, we assume that the parameters in the model are identical for both countries. This is not essential to the model but we have not found it necessary to emphasize cross-country parameter differences. Because of this and for simplicity of exposition, we present the problem for the home country (country 1) representative agent only. Country 2 is the foreign country.

endowment in country *j*. In (3) $Q^j(\tau_{t+1}, \tau_t)$ is the nominal bond price in the currency of country *j* of state τ_{t+1} bonds.

We assume the usual cash in advance constraint:

$$\frac{M_t^j}{P_t^j} = C_t^j, \quad j = 1, 2 \tag{4}$$

We assume an external habit specification and define X_t^j as the surplus consumption ratio of good *j*:

$$X_t^j = \frac{\bar{C}_t^j - H_t^j}{\bar{C}_t^j}, \quad j = 1, 2 \tag{5}$$

where $\bar{C}_t^j$ is aggregate consumption per capita rather than individual consumption: it does not affect individual choice at the margin. Identical individuals choose the same level of consumption in equilibrium, so $C_t^j = \bar{C}_t^j$.

It is assumed that endowment growth in country *j* is an iid process:

$$\Delta y_{t+1}^j = \bar{\mu}^j + v_{t+1}^j, \quad v_{t+1}^j \sim N(0, \sigma_{v^j}^2), \quad j = 1, 2 \tag{6}$$

We follow Chari, Kehoe and McGrattan (2002) and assume that money growth in country *j* follows a simple AR(1) process:

$$\Delta m_{t+1}^j = (1 - \rho_{\pi^j})\bar{\pi}^j + \rho_{\pi^j}\Delta m_t^j + u_{t+1}^j, \quad u_{t+1}^j \sim \text{NIID}\left(0, \sigma_{u^j}^2\right), \quad j = 1, 2 \tag{7}$$

The unconditional means of endowment and money growth for country j are defined as $\bar{\mu}^j$ and $\bar{\pi}^j$ respectively. The variances of shocks to endowment and money growth are defined as $\sigma_{v^j}^2$ and $\sigma_{u^j}^2$ respectively. We also assume that the four shocks are uncorrelated (thus (6) and (7) are four univariate processes).

We follow Campbell and Cochrane (1999) and assume that the log of the surplus consumption ratio of good *j* evolves as follows:

$$x_{t+1}^{j} = (1-\phi)\overline{x}^{j} + \phi x_{t}^{j} + \lambda\left(x_{t}^{j}\right)\left(v_{t+1}^{j}\right), \quad j = 1,2 \tag{8}$$

where $\phi < 1$, is the habit persistence parameter and $\overline{x}$ is the steady state value for the logarithm of the surplus consumption ratio which we define below in equation (10). The function $\lambda(x_t^j)$ describes the sensitivity of the log surplus consumption ratio to endowment innovations. It depends non-linearly on the current log surplus consumption ratio:

$$\begin{aligned} \lambda\left(x_{t}^{j}\right) &= \frac{\sqrt{1-2\left(x_{t}^{j}-\overline{x}^{j}\right)}}{\overline{X}^{j}} - 1 \quad \text{for} \quad x_{t}^{j} \leq \overline{x}^{j} + \frac{1-\left(\overline{X}^{j}\right)^{2}}{2} \\ &= 0 \qquad\qquad\qquad\qquad \text{for} \quad x_{t}^{j} > \overline{x}^{j} + \frac{1-\left(\overline{X}^{j}\right)^{2}}{2} \qquad j = 1,2 \end{aligned} \tag{9}$$

where $\overline{X}^{j}$ is the steady state value of the surplus consumption ratio for good *j* and is defined as:

$$\overline{X}^{j} = \sigma_{v^{j}} \sqrt{\frac{\gamma}{1-\phi-\delta/\gamma}} \tag{10}$$

where $1-\phi > \delta/\gamma$. This is an alternative given in Campbell and Cochrane (1999), which allows for some variation in real interest rates. We can now proceed to use the model solution to derive the conditions under which uncovered interest parity holds.

2.2 Calculating the Slope Coefficient from the Forward Regression

Fama (1994) shows that the slope of the regression of the expected spot return on the forward discount is:

$$b_{1} = \frac{Var\left[E_{t}\left(s_{t+1}-s_{t}\right)\right] + Cov\left[\left(f_{t}-E_{t}\left(s_{t+1}\right),\left(E_{t}\left(s_{t+1}-s_{t}\right)\right)\right)\right]}{Var\left[f_{t}-s_{t}\right]} \tag{11}$$

Fama also shows that two conditions are necessary for a negative slope. The covariance in the numerator must be negative and the variance of expected spot returns must not be too high. The covariance condition arises quite easily in this model (see Appendix 1). When calibrated to western levels of monetary volatility, the second Fama condition is also readily met and the model correctly identifies the Fama regression slope as negative. The point of the analysis that follows is that where there is a sufficiently high level of monetary volatility, the second Fama condition may not be met and the (still) negative covariance above is overwhelmed thereby restoring a positive sign for the regression slope.

Assuming that the nominal stochastic discount factors are log-normal, Backus, Foresi and Telmer (2002) demonstrate that home and foreign nominal interest rates are given by

$$i_t^j = -\ln E_t\left(NMRS_{t+1}^j\right) = -E_t\left(nmrs_{t+1}^j\right) - \frac{1}{2}Var_t\left(nmrs_{t+1}^j\right), \qquad j = 1,2 \tag{12}$$

where *NMRS* is the nominal marginal rate of substitution (lower case *NMRS* is the logarithm). Backus, Foresi and Telmer (2001) show that the expected forward profit is

$$f_t - E_t\left(s_{t+1}\right) = \frac{1}{2}Var_t\left(nmrs_{t+1}^2\right) - \frac{1}{2}Var_t\left(nmrs_{t+1}^1\right) \tag{13}$$

and the home and foreign countries nominal interest rate differential (or forward premium) is

$$\begin{aligned} i_t^1 - i_t^2 = f_t - s_t &= \ln E_t\left(NMRS_{t+1}^1\right) - \ln E_t\left(NMRS_{t+1}^2\right) \\ &= \left(-E_t\left(nmrs_{t+1}^1\right) - \frac{1}{2}Var_t\left(nmrs_{t+1}^1\right)\right) - \left(-E_t\left(nmrs_{t+1}^2\right) - \frac{1}{2}Var_t\left(nmrs_{t+1}^2\right)\right) \\ &= \left(E_t\left(nmrs_{t+1}^2\right) - E_t\left(nmrs_{t+1}^1\right)\right) + \frac{1}{2}\left(Var_t\left(nmrs_{t+1}^2\right) - Var_t\left(nmrs_{t+1}^1\right)\right) \end{aligned} \tag{14}$$

Thus subtracting (13) from (14) the expected change in the spot exchange rate is

$$E_t\Delta s_{t+1} = E_t\left(nmrs_{t+1}^2\right) - E_t\left(nmrs_{t+1}^1\right) \tag{15}$$

In the model described in Section 2.1 the home and foreign $NMRS_{t+1}^{j}$ are

$$NMRS_{t+1}^{j} = \beta\left[\left(\frac{Y_{t+1}^{j}}{Y_{t}^{j}}\right)^{1-\gamma}\left(\frac{X_{t+1}^{j}}{X_{t}^{j}}\right)^{-\gamma}\left(\frac{M_{t+1}^{j}}{M_{t}^{j}}\right)^{-1}\right], \qquad j=1,2. \tag{16}$$

Assuming that the parameters γ, δ and ϕ are the same in both countries, it can be shown[2] that the expected currency excess return is

$$f_t - E_t\left(s_{t+1}\right) = \xi_1 - \left(\gamma\left(1-\phi\right)+\theta_2\right)x_t^2 + \left(\gamma\left(1-\phi\right)+\theta_1\right)x_t^1 \tag{17}$$

where

$$\xi_1 \equiv \frac{1}{2}\left(\sigma_{u^2}^2 - \sigma_{u^1}^2\right) + \frac{1}{2}\left(\left(1-\frac{\gamma}{\overline{X}^1}\right)\sigma_{v^1}^2 - \left(1-\frac{\gamma}{\overline{X}^2}\right)\sigma_{v^2}^2\right) + \left(\gamma\left(1-\phi\right)+\theta_1\right)\overline{x}^1 - \left(\gamma\left(1-\phi\right)+\theta_2\right)\overline{x}^2$$

and

$$\theta_j \equiv -\delta - \sigma_{v^j}\sqrt{\gamma(1-\phi)-\delta} \qquad j=1,2$$

It can also be shown that the forward discount is:

$$f_t - s_t = \xi_2 + \rho_{\pi^1}\Delta m_t^1 - \rho_{\pi^2}\Delta m_t^2 + \theta_1 x_t^1 - \theta_2 x_t^2 \tag{18}$$

where

$$\xi_2 = \xi_1 + \left((1-\rho_{\pi^1})\overline{\pi}^1 - (1-\rho_{\pi^2})\overline{\pi}^2\right)$$

Subtracting (17) from (18) the expected change in the spot exchange rate is

$$E_t\left(s_{t+1}-s_t\right) = -\gamma\left(1-\phi\right)\left(x_t^1 - x_t^2\right) + \left((1-\rho_{\pi^1})\overline{\pi}^1 - (1-\rho_{\pi^2})\overline{\pi}^2\right) + \left(\rho_{\pi^1}\Delta m_t^1 - \rho_{\pi^2}\Delta m_t^2\right) \tag{19}$$

In Appendix 1 we derive the conditional moments in the Fama (1994) regression (11). Thus the theoretical slope coefficient is:

[2] See Appendices A1 to A4 in Moore and Roche (2010)

$$b_1 = \frac{\dfrac{\rho_{\pi^1}^2 \sigma_{u^1}^2}{1-\rho_{\pi^1}^2} + \dfrac{\rho_{\pi^2}^2 \sigma_{u^2}^2}{1-\rho_{\pi^2}^2} - \gamma\left(1-\phi\right)\left(\theta_1 \sigma_{x^1}^2 + \theta_2 \sigma_{x^2}^2\right)}{\dfrac{\rho_{\pi^1}^2 \sigma_{u^1}^2}{1-\rho_{\pi^1}^2} + \dfrac{\rho_{\pi^2}^2 \sigma_{u^2}^2}{1-\rho_{\pi^2}^2} + \left(\theta_1{}^2 \sigma_{x^1}^2 + \theta_2{}^2 \sigma_{x^2}^2\right)} \tag{20}$$

The expression $\dfrac{\rho_{\pi^j}^2 \sigma_{u^j}^2}{1-\rho_{\pi^j}^2}$ is the conditional variance of monetary growth in the foreign country and we interpret this as an index of monetary volatility. It is increasing in the absolute value of the persistence of money growth, ρ_{π^j}, as well as the unconditional variance of money growth, $\sigma_{u^j}^2$. In the extreme case, where monetary volatility is zero, the slope in equation (20) simplifies to:

$$b_1 = \frac{-\gamma\left(1-\phi\right)\left(\theta_1 \sigma_{x^1}^2 + \theta_2 \sigma_{x^2}^2\right)}{\left(\theta_1{}^2 \sigma_{x^1}^2 + \theta_2{}^2 \sigma_{x^2}^2\right)} \tag{21}$$

In this case, at least one of the $\theta_j > 0$ is necessary for a negative slope. A sufficient condition is $\theta_1 = \theta_2 > 0$. It is shown in Moore and Roche (2010) as well as Verdelhan (2010) that the latter corresponds to the case where the precautionary demand for savings in both countries is more important than the standard intertemporal substitution motive. In such circumstances, UIP never holds and the slope in the Fama regression is reassuringly negative.

Now consider the case where $\theta_1 = \theta_2 = 0$. In this case, the two real savings motives are of equal importance, only monetary volatility matters and equation (20) simplifies to unity. Consequently, the model explains the diverse values for the Fama slope by the relative importance of monetary volatility in relation to the balance of real savings motives. At this stage, we evaluate the importance of the model by calibrating the model to actual data.[3]

[3] In the early literature, it was speculated by, for example, Domowitz and Hakkio (1985) that time-varying volatility in the conditional volatility of money growth might explain the 'risk premium'. The irony of equation (20) is that variations in the conditional volatility of money growth helps us to understand UIP rather deviations from it.

3. Simulation

In this section we calibrate the model and generate its predictions for the slope coefficient for a large number of developed (13) and less-developed countries (29).

3.1 Calibration

The parameters governing the money growth processes are allowed to be different for each country and for each identified monetary regime. The baseline parameterization is presented in Table I. We let the U.S. be country 1. We use the growth rate in M1 to proxy for the cash-in-advance money growth rate.[4] An AR(1) process for U.S. money growth is estimated over the period 1983:11 to 2010:10 as one-month spot and forward rates are available for many developed countries for this time period. The unconditional mean of U.S. money growth per capita $\bar{\pi}_1$ is estimated to be 0.38% per month. The AR(1) coefficient of money growth ρ_{π^1} is estimated to be 0.17. The standard deviation of shocks to money growth σ_{u^1} is estimated to be 0.79% per month.

For all developed countries and many less developed countries we can also construct the growth rates of money using M1 from the IMF international financial statistics databank (we have to use M0 for Bolivia, Chile, Georgia, Poland and Sweden). This data is available from Datastream Advance[5]. For most countries the data ends in 2010:10. The sample starts at different time periods for many countries. This is due to either data availability or that in prior periods the exchange rate was pegged to the U.S. dollar. The precise sample periods for all 42 countries are presented in Appendix 2: Data.

[4] The data is available from the Federal Reserve Bank of St. Louis website http://research.stlouisfed.org/fred2/.
[5] http://www.thomson.com/content/financial/brand_overviews/Datastream_Advance

We also collected monthly data on interest rates, inflation rates and spot and forward exchange rates from Datastream Advance. One-month spot and forward rates are available for most developed countries from 1983:11-2010:10.[6] In order to construct forward premia for the less developed countries we assume that covered interest parity holds and use interest rates. For most of these countries we can use money market interest rates. For others we use either the treasury bill rate (Hungary India, Israel, Nigeria, Zimbabwe) or the prime interest rate (Bolivia , Russia, Slovakia, Thailand). This is due to data availability. We use the equivalent one-month interest rate for the domestic (i.e. the U.S.) interest rate and calculate the forward premium for all bilateral pairs via the U.S. dollar.

Consider the numerator in equation (20). If

$$\frac{\rho_{\pi^2}^2 \sigma_{u^2}^2}{1-\rho_{\pi^2}^2} > \frac{\rho_{\pi^1}^2 \sigma_{u^1}^2}{1-\rho_{\pi^1}^2} - \gamma\left(1-\phi\right)\left(\theta_1 \sigma_{x^1}^2 + \theta_2 \sigma_{x^2}^2\right) \tag{22}$$

then $b_1 > 0$. Thus we define a volatile monetary regime when the expected variance of foreign money growth meets this criterion. Since the main thesis in this paper is that it is the monetary regime that is the main driver of whether uncovered interest rate parity holds we take considerable care in identifying changes in the regime by observing money supply, interest rates and inflation for each country in time periods where the exchange rate was market driven.

If there are sufficient numbers of observations on foreign money growth per capita we estimate AR(1) models for both stable and volatile monetary regime periods and perform simple Chow and Hansen tests for parameter stability. The time periods for possible breaks for the following thirteen countries are New Zealand (1997:12), Singapore (1997:12), Bolivia (2005:5), Brazil (1991:2), Mexico (1995:12), Uruguay (2003:7), Venezuela (2003:3), Bulgaria(1997:7), Estonia

[6] A description of the data and time periods is given in Appendix 2 and more detail is contained in a not for publication Data Appendix.

(2004:6), Georgia (2004:12), Slovenia (1996:4), Philippines (1997:12) and Cyprus (2000:12). The null hypotheses in these tests are rejected[7].

There are a number of other parameters to choose in the model. The baseline parameterization is presented in Table I. Campbell and Cochrane (1999) choose the AR(1) coefficient of the log of the surplus consumption ratio, ϕ, to mimic the first order serial correlation coefficient of the log price-dividend ratio in the United States. The data on monthly forward and spot exchange rates that we present below covers the period 1983:11 to 2010:10. Using monthly price-dividend ratio for the U.S. over this period we estimate ϕ to be 0.994.[8] The power parameter in the utility function, γ, is set equal[9] to 2 as in Campbell and Cochrane (1999). The parameter δ has to be negative for the precautionary savings motive to dominate the intertemporal substitution savings motive. If the absolute value of δ is too large, interest rates and the forward premia will be too volatile. We follow Moore and Roche (2010) and set $\delta = -0.005$. We follow Campbell and Cochrane (1999) who use seasonally adjusted real consumption expenditure on non-durables and services per capita to proxy for endowments in the United States. The following parameters for monthly endowment growth rates are based on Table 1 in Campbell and Cochrane (1999). The unconditional mean $\bar{\mu}_1$ is set equal to 0.1575% per month. The standard deviation of shocks to endowment growth σ_{v^1} is set to 0.433% per month. The parameters ϕ, γ, δ and σ_{v^1} imply that the steady state surplus consumption ratio is approximately 6% as in Campbell and Cochrane (1999).

[7] Test results available on request.

[8] The data can be downloaded from Robert Shiller's website http://aida.econ.yale.edu/~shiller/data.htm.

[9] Moore and Roche (2010) used $\gamma = 0.5$ as their benchmark value. Figure 1 is not as neat in this case because very low levels of the power parameter makes the volatility of surplus consumption so high that it takes a relatively higher level monetary volatility to achieve a positive Fama regression slope. However, the qualitative message of this paper remains unchanged: only high monetary volatility can achieve a positive Fama regression slope once the precautionary savings motive is dominant.

An expression for the variance of the log of surplus consumption, σ_x^2 cannot be derived as a closed form solution. Therefore we simulate equations (8)-(9) for a sample size of 1,000,000 and assume that the resulting estimate of σ_x^2 is the population variance. We estimate $\sigma_x = 0.5359$. This estimate is used along with the expected money growth parameters to generate the model's prediction for the slope coefficient.

Since we are interested in examining the effects of changing monetary regimes on the slope coefficient we initially assume that the real parameters for the foreign countries are the same as those for the domestic country. Another reason for this assumption is that volatile monetary regimes are of relatively short durations and it would be impossible to get precise estimates of the endowment process using annual consumption expenditure data for less developed countries. However it might be reasonable to assume that the standard deviation of shocks to endowment growth in LDCs might be larger than that in the U.S. In sensitivity analysis we set the standard deviation of shocks to endowment growth in LDCs to double that we use for the U.S. Thus σ_{v^2} is set to 0.866% per month for LDCs. This affects the variance of the log of surplus consumption and the parameter θ_j in (20), in opposite directions. In our baseline experiment we estimate $\sigma_x = 0.5359$ and $\theta_j = 0.0044$. When we double the size of the real shock in LDCs we estimate $\sigma_x = 0.5024$ and $\theta_j = 0.005$. One might have expected that variance of the log of surplus consumption would get larger as the shock gets larger but the sensitivity factor in (8) decreases on average as steady state surplus consumption increases. As we will see in the next section the net effect is too small to alter our main results.

3.2 Results

In Table II we present the time period, the standard deviation of the expected monetary growth per month, the average annual inflation rate and the slope coefficient estimated using actual and simulated data under the stable monetary regime. The slope coefficient estimated using simulated data is in the column headed Model A for the baseline parameterization and in the column headed Model B for the simulation where we set the standard deviation of shocks to endowment growth in LDCs to double that we use for the U.S. In Table III we present the same statistics for countries that experienced a volatile monetary regime for some time period. If countries appear (in the first column) in both tables we attach 1 to the country name in the stable regime and 2 to the country name in the volatile regime. Fourteen out of forty-two countries have both types of regime.

Each Table has two panels. In the upper panel, we show the results for developed countries and in the lower panel for less developed countries. In the course of our sample, a number of countries, such as South Korea, have graduated from less developed to developed status. It is also true that some countries such as South Africa and perhaps even New Zealand have moved in the other direction. Our classification of countries as developed is based on how they would have been regarded at the very beginning of the sample.

In stable monetary regimes the estimated slope coefficients in the data are always negative for developed countries. New Zealand and Singapore are the two developed countries that display both stable and volatile monetary periods. The stable periods for both countries occur after the Asian financial crisis when the standard deviation of expected money growth fell and the estimated slope coefficients in the data and in the model are negative for both countries (labelled New Zealand 1 and Singapore 1 in Table II).

We find that during periods where the monetary regime is stable, when the standard deviation of expected money growth is low, the estimated slope coefficients in the data and in the model are negative for less developed countries in 14 out of 16 cases. For Georgia, the slope coefficient is close to zero in the model and in the data. For South Korea the model produces a negative slope coefficient while it is positive in the data.

Table III shows the results for volatile monetary regimes. Most of the countries that are represented here are in the less developed category but there are three exceptions. The annual growth rates in M1 for New Zealand and Singapore were very erratic up to 1997:12 and the slope coefficients were 0.78 for both countries (labelled New Zealand 2 and Singapore 2 in Table III). The model successfully reflects these positive slopes. The model is less successful with Hong Kong where the Fama coefficient is negative while the model simulates the slope as close to plus unity. The reason is not hard to find: the annual growth rates in M1 for Hong Kong were very erratic throughout the sample period. It may well be that M1 is not the most appropriate measure of the money stock in Hong Kong. For less developed countries, we find that during periods where the monetary regime is relatively unstable, when the standard deviation of expected money growth is high, the estimated slope coefficients in the data and in the model are positive in all 25 cases that we report in Table III.

The results presented in Tables II and III are graphed in Figure I. The South-West quadrant depicts the case where the slope is negative in the model (read along the vertical axis) and in the data. The North-East quadrant depicts the case where the slope is positive in the model and in the data. The simple correlation coefficient between the slope coefficient simulated in the model and estimated in the data is 0.78. Almost all points lie in these quadrants and close to the 45^0 line.

Our results do not depend on the level of inflation as Bansal and Dalquist (2000) suggest. For example the annual average inflation rate was high for Zimbabwe (at 830%) and Brazil 2 (at 806%) and was low for Thailand (at 2.94%) and Chile (at 4.8%) during the periods where the monetary regime is unstable and all slope coefficients were positive in both the model and the data (see Table III). While the annual average inflation rate was high for Venezuela 1 (at 21%) and was low for Cyprus 1 (at 2.85%) during the periods where the monetary regime is stable and all slope coefficients were negative in both the model and the data (see Table II). Venezuela had relatively high inflation rates during both its monetary regimes but its estimated slope coefficients were positive in the unstable regime and negative in the stable monetary regime. With the exception of Estonia and Georgia, the stable monetary regime time period is generally from the late 1990s to the end of sample.

4. Conclusion

This paper has proposed a modelling strategy that makes substantial progress towards explaining why the forward bias puzzle only arises between some pairs of countries and not for others. A model that combines Campbell and Cochrane (1999) habit persistence defined over individual goods in a monetary framework identifies two different forces at work. Where monetary policy is stable, interest rates are primarily determined by real behaviour. In those circumstances, the importance of the precautionary savings motive ensures that the forward bias typically arises. This is why uncovered interest parity is not usually observed between developed countries. In contrast, in countries where monetary volatility dominates, something closer to interest parity is observed because nominal bond holders have to be compensated for the nominal volatility.

Monetary volatility is defined quite precisely here: it means a high conditional variance for money growth. We calibrated this model to 13 developed and 29 emerging economies. We are successfully able to explain when UIP holds and does not hold without referring explicitly to the income, inflation rate nor level of development of the countries concerned. As far as we are aware, we are the first to develop a theoretical model which achieves this.

Obviously, this model is highly stylised. We are assuming complete markets with perfect international risk sharing. This is particularly difficult when dealing with some of the very underdeveloped economies that we examine here. The evidence of Figure 1 is all the more compelling because of this.

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Table I
Baseline parameterization

	Endowment growth	Money growth
Unconditional mean	0.1575%	0.38%
AR(1) coefficient	0.00	0.17
Standard deviation of shock	0.433%	0.79%
Curvature of the utility function γ		2.000
AR1 coefficient of log surplus consumption ϕ		0.994
Parameter in steady state surplus consumption δ		-0.005

Notes to the table: The endowment growth parameters are taken from Campbell and Cochrane (1999) but transformed to monthly frequency. The money growth parameters are estimated from an AR(1) model for money growth using U.S. data over the period 1983:11 to 2010:10. The curvature of the utility function is taken from Campbell and Cochrane (1999). Following Campbell and Cochrane (1999) the AR1 coefficient of log surplus consumption is estimated using monthly price-dividend ratio for the U.S. over the period 1983:11 to 2010:10.

Table II
Stable Monetary Regimes

Country	Time-Period	% Standard Deviation Expected Money Growth	% Mean Inflation Rate	Slope Coefficient		
				Data	Model A	Model B
Developed countries						
Australia	1985:1-2010:10	0.07	3.98	-1.357	-2.029	-2.048
Canada	1985:1-2010:10	0.14	2.76	-0.329	-1.722	-1.697
Denmark	1985:1-2010:10	0.43	2.54	-0.056	-0.333	-0.234
Euro	1999:1-2010:10	0.02	2.06	-2.438	-2.142	-2.181
Japan	1983:11-2010:10	0.25	0.64	-1.494	-1.165	-1.086
New Zealand 1	1998:1-2010:10	0.37	2.30	-1.129	-0.555	-0.454
Norway	1993:1-2010:10	0.33	2.10	-0.652	-0.735	-0.637
Singapore 1	1998:1-2010:10	0.18	1.29	-1.129	-1.550	-1.504
South Africa	1983:11-2010:10	0.30	0.29	-0.424	-0.881	-0.787
Sweden	1985:1-2010:10	0.32	2.97	-0.005	-0.772	-0.674
Switzerland	1985:1-2010:10	0.49	1.71	-0.156	-0.122	-0.028
UK	1986:9-2010:10	0.03	3.78	-0.522	-2.132	-2.168
Less developed countries						
Bolivia 1	2005:6-2008:11	0.13	7.92	-0.910	-1.782	-1.764
Brazil 1	1999:2-2010:10	0.21	6.86	-0.994	-1.995	-2.008
Bulgaria 1	1997:8-2010:10	0.19	13.86	-0.814	-1.495	-1.443
Cyprus 1	2001:1-2007:12	0.29	2.85	-0.411	-0.920	-0.828
Estonia 1	1997:1-2004:6	0.13	4.93	-0.251	-1.777	-1.759
Georgia 1	1995:12:2004:12	0.51	11.13	0.067	-0.055	0.037
Hungary	1998:2-2010:10	0.11	7.52	-0.266	-1.867	-1.861
India	1995:10-2006:7	0.32	5.98	-0.920	-0.773	-0.675
Latvia	1993:9-2006:11	0.23	10.35	-0.257	-1.265	-1.192
Mexico 1	1996:1-2010:10	0.10	9.73	-0.100	-1.938	-1.943
Philippines 1	1998:1-2010:10	0.04	5.51	-0.773	-2.110	-2.143
Slovenia 1	1996:4-2006:3	0.15	6.71	-0.570	-1.689	-1.659
South Korea	1991:1-2010:10	0.33	4.16	0.304	-0.728	-0.630
Turkey 1	2004:8-2010:10	0.40	8.71	-0.172	-0.436	-0.335
Uruguay 1	2003:7-2006:8	0.39	7.78	-0.940	-0.474	-0.373
Venezuela 1	2003:3-2006:6	0.18	20.75	-1.360	-1.516	-1.467

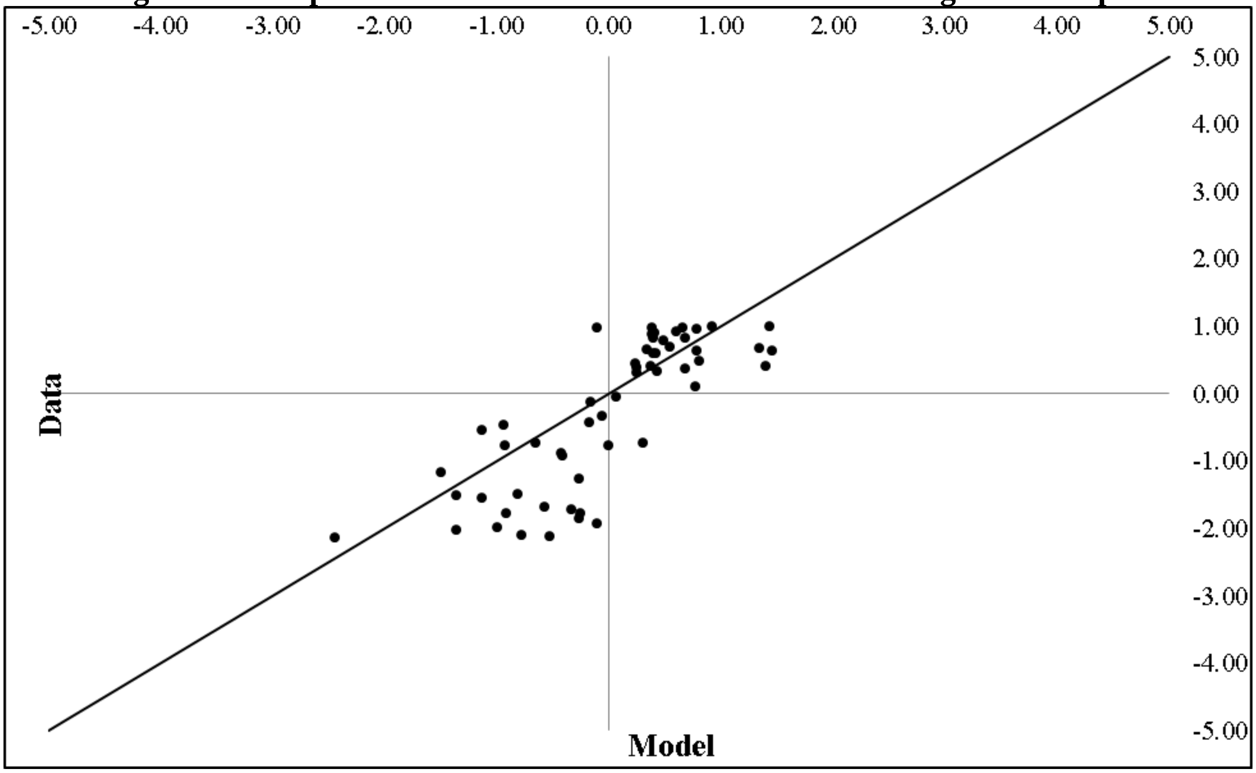
Notes to the table: The standard deviation for expected money growth is estimated from parameters of the AR (1) model for money growth. The slope coefficient is estimated from a regression of the spot change on the forward premium. All exchange rates are via the U.S. dollar. The column headed Model A refers to a simulation where the real side of the economy is the same for the U.S. and the other country. The column headed Model B refers to a simulation where the standard deviation of the endowment shocks for LDCs are double that of the U.S. economy.

Table III
Volatile Monetary Regimes

Country	Time-Period	% Standard Deviation Expected Money Growth	% Mean Inflation Rate	Slope Coefficient		
				Data	Model A	Model B
Developed countries						
Hong Kong	1983:11-2010:10	4.03	4.18	-0.107	0.974	0.978
New Zealand 2	1985:1-1997:12	1.00	5.89	0.783	0.632	0.675
Singapore 2	1985:1-1997:12	3.04	1.77	0.783	0.955	0.961
Less developed countries						
Argentina	2002:02-2010:10	0.76	11.20	1.401	0.409	0.471
Bolivia 2	1991:1-2005:5	3.69	6.79	0.663	0.970	0.973
Brazil 2	1991:1-1999:1	43.53	806.20	0.927	1.000	1.000
Bulgaria 2	1995:12-1997:7	4.83	143.80	0.388	0.982	0.984
Chile	1994:1-2006:8	0.82	4.82	0.810	0.477	0.534
Columbia	1995:3-2010:10	0.57	9.04	0.772	0.092	0.177
Cyprus 2	1996:1-2000:12	2.32	2.82	0.608	0.924	0.933
Estonia 2	2004:7-2010:10	0.76	4.52	0.374	0.407	0.470
Georgia 2	2005:1-2009:9	1.07	7.81	1.349	0.672	0.710
Indonesia	1988:1-2010:10	0.74	12.73	0.248	0.385	5.731
Israel	1984:6-2010:10	1.80	37.77	0.388	0.876	0.891
Lithuania	2002:2-2010:10	0.70	3.13	0.429	0.333	0.402
Mexico 2	1985:12-1995:12	1.45	47.69	0.681	0.813	0.836
Nigeria	2000:1-2010:10	0.95	13.80	0.428	0.599	0.644
Philippines 2	1982:1-1997:12	2.08	12.31	0.412	0.906	0.918
Poland	1993:6-2010:10	1.32	19.26	0.495	0.776	0.803
Romania	1995:8-2006:11	0.94	41.25	0.400	0.586	0.633
Russia	1995:6-2010:10	0.79	29.00	0.238	0.448	0.507
Slovakia	1993:1-2008:11	7.14	7.07	1.437	0.992	0.993
Slovenia 2	1991:12-1996:3	0.68	21.70	0.250	0.304	0.375
Thailand	1997:7:2010:10	1.08	2.94	0.543	0.680	0.718
Turkey 2	1998:1-2004:7	1.04	8.37	0.348	0.655	0.695
Uruguay 2	1993:12-2003:06	0.72	20.22	0.690	0.362	0.429
Venezuela 2	1996:1-2003:2	1.01	37.49	1.460	0.635	0.677
Zimbabwe	2004:1-2006:10	1.45	830.33	0.400	0.812	0.835

Notes to the table: The standard deviation for expected money growth is estimated from parameters of the AR (1) model for money growth. The slope coefficient is estimated from a regression of the spot change on the forward premium. All exchange rates are via the U.S. dollar. The column headed Model A refers to a simulation where the real side of the economy is the same for the U.S. and the other country. The column headed Model B refers to a simulation where the standard deviation of the endowment shocks for LDCs are double that of the U.S. economy.

Figure 1. Comparison of Theoretical and Estimated Fama regression slopes.



Notes to the figure: The slope is obtained from a linear regression of the expected spot return on the forward discount or nominal interest rate differential. The estimated slope using actual data is plotted along the horizontal axis. The estimated slope from the theoretical model is plotted along the vertical axis.

Appendix 1: Calculating unconditional second moments

In this appendix we derive the conditional moments in the Fama (1994) regression (11). Using (19) the unconditional variance of the expected spot change is:

$$Var\left(E_t\left(s_{t+1}-s_t\right)\right)=\left(\frac{\rho_{1\pi}^2\sigma_{1u}^2}{1-\rho_{1\pi}^2}+\frac{\rho_{2\pi}^2\sigma_{2u}^2}{1-\rho_{2\pi}^2}+(1-\phi)^2\gamma^2\left(\sigma_{x^1}^2+\sigma_{x^2}^2\right)\right) \tag{23}$$

The covariance between expected forward profits (17) and the expected spot return (19) is:

$$\begin{aligned}Cov\left[\left(f_t-E_t\left(s_{t+1}\right),\left(E_t\left(s_{t+1}-s_t\right)\right)\right)\right]&=-\gamma\left(1-\phi\right)\left(\gamma\left(1-\phi\right)+\theta_1\right)\sigma_{x^1}^2\\&\quad-\gamma\left(1-\phi\right)\left(\gamma\left(1-\phi\right)+\theta_2\right)\sigma_{x^2}^2\end{aligned} \tag{24}$$

Given our baseline parameter values discussed in Section 3 the covariance (24) is negative, which is one of Fama's necessary conditions for a negative slope. Using (18) the unconditional variance of the forward discount is:

$$Var\left(f_t-s_t\right)=\frac{\rho_{\pi^1}^2\sigma_{u^1}^2}{1-\rho_{\pi^1}^2}+\frac{\rho_{\pi^2}^2\sigma_{u^2}^2}{1-\rho_{\pi^2}^2}+\left(\theta_1^2\sigma_{x^1}^2+\theta_2^2\sigma_{x^2}^2\right) \tag{25}$$

Appendix 2: Data

Country	Sample period	Money	Interest Rate
Developed countries			
Australia	1985:1-2010:10	M1	Forward Premium
Canada	1985:1-2010:10	M1	Forward Premium
Denmark	1985:1-2010:10	M1	Forward Premium
Euro	1999:1-2010:10	M1	Forward Premium
Hong Kong	1983:11-2010:10	M1	Forward Premium
Japan	1983:11-2010:10	M1	Forward Premium
New Zealand	1985:1-2010:10	M1	Forward Premium
Norway	1993:1-2010:10	M1	Forward Premium
Singapore	1985:1-2010:10	M1	Forward Premium
South Africa	1983:11-2010:10	M1	Forward Premium
Sweden	1985:1-2010:10	M0	Forward Premium
Switzerland	1985:1-2010:10	M1	Forward Premium
UK	1986:9-2010:10	M1	Forward Premium
Less developed countries			
Argentina	2002:02-2010:10	M1	Interbank
Bolivia	1991:1-2008:11	M0	Prime
Brazil	1991:1-2010:10	M1	Interbank
Bulgaria	1995:12-2010:10	M1	Interbank
Chile	1994:1-2006:8	M0	Interbank
Columbia	1995:3-2010:10	M1	Interbank
Cyprus	1996:1-2007:12	M1	Interbank
Estonia	1997:1-2010:10	M1	Interbank
Georgia	1995:12-2009:9	M0	Prime
Hungary	1998:2-2010:10	M1	TBill
India	1995:10-2006:7	M1	TBill
Indonesia	1988:1-2010:10	M1	Interbank
Israel	1984:6-2010:10	M1	TBill
Latvia	1993:9-2006:11	M1	Interbank
Lithuania	2002:2-2010:10	M1	Interbank
Mexico	1985:12-2010:10	M1	Interbank
Nigeria	2000:1-2010:10	M1	TBill
Philippines	1982:1-2010:10	M1	Interbank
Poland	1993:6-2010:10	M0	Interbank
Romania	1995:8-2006:11	M1	Interbank
Russia	1995:6-2010:10	M1	Prime
Slovakia	1993:1-2008:11	M1	Prime
Slovenia	1991:12-2006:3	M1	Interbank
South Korea	1991:1-2010:10	M1	Interbank
Thailand	1997:7:2010:10	M1	Prime
Turkey	1998:1-2010:10	M1	Interbank
Uruguay	1993:12-2006:8	M1	Interbank
Venezuela	1996:1-2006:6	M1	Interbank
Zimbabwe	2004:1-2006:10	M1	TBill