Does the forward premium puzzle disappear over the horizon?

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Abstract

The forward rate unbiasedness hypothesis has received little empirical support in most extant studies that have predominantly used short horizon data. This paper examines the puzzle using short, medium and long horizon data for five US dollar currency pairs 1980-2005. It develops a behavioral finance model that predicts the bias will abate as the horizon is extended. A heteroskedastic and autocorrelation consistent bootstrap approach is employed to deal with the data-overlap problem at medium and long horizons. The results from using this procedure replicate the puzzle at short horizons but indicate that it disappears at medium horizons and beyond.

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1. Introduction

The forward rate unbiasedness hypothesis (FRUH) implies that the forward rate is an unbiased predictor of the corresponding future spot rate under the assumptions of risk neutrality and rational expectations. Another version is that the forward premium is an unbiased predictor of the change in the future spot rate. Empirical tests of the latter relationship typically yield negative coefficients on the forward premium where positive values are expected. The implication is that high interest rate currencies are predicted to appreciate rather than depreciate in line with the FRUH. Despite widespread tests across different time frames and currencies, this result has remained stubbornly robust and has become known as the forward premium (FP) puzzle.

There is a voluminous literature on the FP puzzle (see, *inter alios*, Sarno, 2005 and Engle, 1996). Burnside, Eichenbaum, and Rebelo (2009) comment that much of the literature is characterized by two key features. On one hand, the foreign exchange market is modeled as a frictionless Walrasian market despite the fact that it is an OTC market in which market makers play a key role as Sarno and Taylor (2001) highlight. On the other hand, empirical and other studies emphasize risk-based explanations for the forward premium despite the fact that existing risk-based explanations seem to fail.[1] Some behavioral explanations of the puzzle have begun to emerge. Burnside et al. (2009) develop a microstructure model with frictions in which adverse selection and behavioral biases can explain the forward premium puzzle. Han et al. (2007) develop a behavioral model in which overconfidence explains the FP puzzle.

The first contribution is the development of a theoretical model that draws on several sources. One is the seminal Frankel and Froot (1991) distinction between rational and chartist investors. Another is the Meredith and Chinn (2006) paper that suggests that the

[1] One interesting new type of risk is rare disaster risk as in the Gabaix and Mehra (2009) complete markets model that can explain both the FP puzzle and the excess volatility of exchange rates.

puzzle disappears at long horizons. Finally, it draws inspiration from the recent behavioral[2] explanation for the FP puzzle by Han *et al.* (2007). Our model eschews the representative agent framework and instead is in the spirit of disagreement among agents in Hong and Stein (2009). Specifically, it builds on survey evidence that foreign exchange dealers base their trades exclusively on fundamentals at horizons of typically more than one year.[3] For short horizon trades, dealers employ heuristics such as technical trading and other 'non-fundamental' rules. For instance, Cheung and Chinn (2001) suggest US-based traders believe that short-run exchange rate dynamics are predominantly a function of non-fundamentals such as 'bandwagon effects' or 'overreactions to news'.

As a consequence, our model is set in a heterogeneous agent framework in which rational and *overconfident* traders operate. Following Han *et al*. (2007), the latter overestimate the precision with which they can assess the impact of new information and therefore overreact in the short term. Like Daniel *et al.* (2001), our model assume that all investors are risk averse and thus prices reflect a weighted average of the beliefs of different investors. This allows the explicit modeling of the ratio of overconfident to rational investors (c_k). This crucial role of this ratio is one the features that distinguishes our model from the Han et al. model.

The vast majority of empirical studies have employed short horizon[4] data of 1 to 3 months to examine the FP puzzle. Very few papers systematically investigate the puzzle at other horizons with the notable exception of Chinn and Meredith (2004).[5] The second contribution is that the model can explain the anomalous short horizon regressions results and provide a rationale for the disappearance of the puzzle as the horizon increases. Consider

[2] See Hirshleifer (2001) for an overview of recent investor overconfidence models and the areas of finance to which they have been applied.
[3] See Allen and Taylor (1990) and Cheung and Chinn (2001).
[4] The term 'horizon' represents the time to maturity of the relevant forward contract.
[5] For example, Chaboud and Wright (2005) use intraday data to estimate slope coefficients commensurate with the FRUH. However, this result is fragile as it is "destroyed" by adding as little as a few hours to the horizon.

that the effects of overconfidence are most keenly felt at short horizons since overconfident dominate rational investors over such time scales (i.e., relatively large c_k). Moreover, these behavioral effects are also felt more in the forward market relative to other markets, as investors find it cheaper to speculate using forwards or swaps to avoid the costs of marking-to-market incurred in futures trading.[6] The implication is that the forward exchange rate overreacts more to new information than the spot exchange rate over short horizons. In this case, and analogously to Han *et al.* (2007), it can be shown that the forward premium will be negatively correlated with the change in the future spot rate, during the period when the overshot exchange rates are correcting towards the new equilibrium. However, beyond short horizons, rational investors begin to hold sway (i.e., progressively smaller c_k) and the model predicts that forward and spot exchange rates become more closely aligned with fundamentals. In this manner, the model is consistent with the stylized survey findings and provides a theoretical basis for examining the term structure of the FP puzzle or, in other words, how swiftly it disappears? By contrast, the existing empirical literature has focused on the puzzle at either short or long horizons, ignoring medium horizons[7].

The third contribution is to address the problem of inference in the presence of data-overlapping arising from the use of medium and long horizon data. Consider that the variables in the typical regression to examine FRUH are non-contemporaneous. When the horizon used in the regression exceeds the sampling frequency, the data overlap problem induces serial correlation in the error term. This may not be pertinent at shorter horizons as the obvious solution is to use non-overlapping data. However this procedure is not appropriate for longer horizons since it would result in the loss of too much information. One solution is to use a heteroskedastic and autocorrelation consistent (HAC) estimator

[6] See Table 2 in Melvin and Taylor (2009) showing that FX forward swap spreads were considerably cheaper that spot spreads up to the crisis in 2007.

[7] In the literature, the prefix ‘short’ typically refers to horizons of less than a year. On the other hand, ‘long’ refers to horizons of five years or more. Consequently, we define intermediate (medium) horizons as between one and five years.

which corrects the standard errors and allows for appropriate inference. This type of correction falls short of fully dealing with this problem, as shown by Nelson and Kim (1993). They find that the distribution of the *t*-statistic is severely shifted even when using HAC corrected standard errors**.**

This paper addresses the data overlap problem through the application of a bootstrap procedure. This is similar to that found in the predictive regression literature in both foreign exchange and stock markets (Killian and Taylor, 2003; Rapach and Wohar, 2005). The bootstrap procedure generates HAC *t*-statistics from pseudo data forming an empirical distribution against which the original *t*-statistic can be compared. The associated *p*-values can be used to draw statistical inference even in the presence of large data overlaps. Our results indicate that the negativity of the FP slope coefficient reduces and tends to unity as we advance towards medium horizons. Employing monthly data on the five most traded currencies relative to the US dollar over the period 1980-2005, there is a clear pattern where the strong rejection of the unbiasedness null at the short horizon is overturned at medium horizons. In particular, the bootstrap results support the FRUH found at horizons ranging from 2 to 5 years in all cases and overturn some of the rejections based on HAC standard errors. These findings are consistent with an ‘overshooting’ behavioral finance model and with the oft-quoted maxim of Flood and Taylor (1997) that “fundamental things apply as time goes by.”

The remainder of this paper is organized as follows. Section 2 provides a discussion of the FP puzzle and the related long horizon literature. Section 3 presents a simple behavioral model of exchange rates with investor overreaction. Section 4 outlines the econometric methodology and presents the empirical results. Finally, Section 5 summarizes the conclusions from the paper.

2. The forward premium puzzle

2.1. Short horizon spot return regression

The forward premium puzzle arises in the context of testing the FRUH or forward market efficiency. Researchers typically estimate the following spot return regression:

$$\Delta s_{t+k} = a_k + b_k (f_{t,t+k} - s_t) + u_{t+k} \tag{1}$$

where k is the horizon, s_t and f_t are the logarithms of the spot and forward exchange rates at time t respectively, and u_{t+k} is a zero-mean error term. Under FRUH one would expect $a_k = 0$, $b_k = 1$ and u_{t+k} to be serially uncorrelated, so that an investor cannot earn excess returns. However, for small values of k, the slope coefficient on the forward premium is typically negative and this constitutes the FP puzzle.[8]

The extant literature employs two broad approaches to explain the deviations of b_k from its theoretical value of unity. The first approach seeks a theoretical justification and encompasses explanations such as an omitted risk premium (Fama, 1984), irrational expectations (Frankel and Froot, 1987; Frankel, 1989) and the Peso problem (Lewis, 1989; Evans and Lewis, 1995). The second approach is statistical, examining the time-series properties of the inputs to equation (1). For example, there is overwhelming statistical evidence that spot exchange rates are I(1) and hence spot returns are I(0). However, less clear is the order of integration for the forward premium, although it typically described as highly persistent which leads to inference problems[9] in (1). Some studies (Crowder, 1994; Kellard *et al.*, 2000) go further and suggest the forward premium presents a unit root. Notably, Evans and Lewis (1995) argue this is unrealistic under Fama's model of a risk

[8] Froot and Thaler (1990) find that the average reported slope to be -0.88 over 75 published articles.
[9] Newbold *et al*. (1998) also note that the forward premium is relatively small in magnitude and thus its time series properties can be empirically swamped by those of the larger spot return.

premium, implying that given an I(0) spot return and error term, the risk premium itself would have to be I(1) and cointegrated with the forward premium.

The recent literature addresses the controversy over integration order by seeking to model the forward premium as a more complex process. For example, Baillie and Bollerslev (2000) and Maynard and Phillips (2001) examine whether fractionally integrated long memory in the forward premium can explain the negative estimates of b_k. Specifically, Baillie and Bollerslev (2000) simulate a model of the foreign exchange market where the assumed long memory behaviour of the conditional variance is inherited by the forward premium. Interestingly, the simulated results are broadly consistent with the empirical features of the forward premium puzzle, b_k converging slowly to its true value of unity. In support of this approach, Maynard and Phillips (2001) develop the relevant asymptotic theory.

Sakoulis and Zivot (2001) comment that long memory modeling requires a relatively large number of observations to provide reliable estimates of an integration order between 0 and 1. Furthermore, the long memory properties of the data can be produced by the stochastic break model of Bai (1997). Therefore, Sakoulis and Zivot advocate modeling the forward premium as a process with infrequent structural breaks and uncover evidence of reduced bias in estimates of b_k. Drawing together two strands of the literature, Choi and Zivot (2007) present empirical results that the long memory properties of the forward premium are jointly due to structural breaks and fractionally integrated behavior[10].

The statistical strand of the literature has yet to definitely resolve the puzzle. Indeed, Maynard (2006) questions whether the statistical properties of the differenced spot rate and forward premium are the sole contributor to the bias. Employing three estimation methods that aim to circumvent the relevant econometric problems: a covariance based test (Maynard

[10] As the authors note, there is no formal statistical theory to justify their approach.

and Shimotsu, 2009), the exact finite sample signs tests (Campbell and Dufour, 1997) and an optimal conditional test (Jansson and Moreira, 2006), Maynard (2006) shows that whilst the bias is reduced, b_k is still far short of its theoretical value of unity. Following this mixed evidence on the source of the bias, he concludes that statistical methods alone are unlikely to resolve the puzzle. Finally, Kellard and Sarantis (2008) provide evidence that the economic and statistical explanations of the puzzle may be combined. Employing a rational expectations framework, it is shown that a consumption CAPM implies that long memory in the true risk premium and the conditional variance of the spot rate can explain analogous time series behaviour in the forward premium; the induced long memory in the forward premium then resulting in the long tailed distributions that help produce the large number of negative b_k coefficients.

2.2. Long horizon regressions

There is little research examining the FP puzzle at long horizons[11]. However, a recent survey by Chinn (2006), concludes that by incorporating longer horizons, the lack of supportive evidence for FRUH from the current float is not as great as commonly thought. However, it is still the case that short-term interest rate differentials and hence forward premia, remain a biased predictor of *ex post* exchange rate changes.[12]

The early study of Flood and Taylor (1997) highlights how the magnitude of the bias is a function of horizon. Employing 3-year government bond rates and 3-year exchange rate changes, they find a significant coefficient on the interest rate differential (forward premium) of 0.596 but the null of unity is rejected. In a similar vein, Alexius (2001) estimates positive b_k coefficients using 14 long-term government bond rates from 1957-1997.

[11] The focus of much of this literature is on the very long horizon. Our paper is distinct from the extant literature as we evaluate more horizons, with particular focus on the medium term.

[12] Bekaert *et al.* (2007) examine the puzzle over the term structure and suggest it is the choice of currency pair that best serves at reducing the bias rather than horizon.

Chinn and Meredith (2004) make an important contribution to the literature by testing the largest number of horizons to date: 3 months, 6 months, 12 months, 5 years and 10 years.[13] At the short horizon and employing US dollar bilateral rates for 6 countries, their results are in agreement with the extant literature; the majority of b_k coefficients being significantly negative. For example, at the 3-month horizon, negative estimates range from -2.887 for the Japanese Yen to -0.477 for the Canadian dollar.[14] Interestingly, at the 12 month horizon, 5 out of 6 b_k estimates are more (less) positive (negative) than at either the 3 or 6-month horizons but are mainly insignificant.[15] On the other hand, the evidence at the long horizon is more striking. The 10-year benchmark government bond dataset yields slope b_k coefficients from 0.197 to 1.120; all are positive and many close to unity. Moreover, estimates range from 0.603 to 0.918 for the 10-year constant maturity series and 0.373, 0.455 and 0.870 for the 5-year constant maturity series. Clearly it is noteworthy that b_k estimates at the 5-year horizon are all closer to unity than the corresponding country result at the short horizon, and further, that the 10-year results (using both datasets) are closer still. Thus, Chinn and Meredith (2004) establish that b_k estimates typically tend towards their hypothesized unity value as the horizon increases, even if there are still a few rejections of FRUH at the long horizon.

3. Models of the forward premium puzzle

3.1. Extant models

[13] Chinn and Meredith focus on the G7 countries. See Madarassy and Chinn (2002) for a similar methodology focusing on non-G7 developed countries where support for the FRUH is weaker than for the G7 countries. Interestingly, Bansal and Dahlquist (2000) show it is easier to find support for developing countries at shorter horizons.

[14] GMM estimation is used throughout to deal with the problem of data-overlapping. The estimate for the remaining country, Italy, is the only positive coefficient at 0.518.

[15] The unity null is rejected at all short horizons for 4 out of the 6 countries tested.

Existing models build on macroeconomic or behavioral finance principles. In the former category, the McCallum (1994) model is based upon a two equation system consisting of a monetary policy rule and an uncovered interest parity (UIP) relationship. The policy rule models the response of short-term interest rates to monetary shocks and, in particular, assumes that monetary authorities tend to smooth such rates and prevent them from departing too far from their value in the recent past. Although McCallum shows for conventional model parameters the negative bias for the estimates of b_k can be reproduced, Mark and Wu (1996) question the significance of policy rule coefficients for a number of developed countries. Offering an alternative policy rule, their model does not re-produce the negative bias for a reasonable change in that policy rule.

Chinn and Meredith (2004) generalize the McCallum model to allow for fuller macroeconomic interaction. Introducing output and inflation equations, they allow the monetary reaction function to adjust in response to movements in both variables. Simulated spot return regressions yield increasingly positive (less negative) estimates of b_k as the horizon increases. However, one drawback is that the model does not provide underlying market shocks that are sufficient to generate the observed volatility in exchange rates.

Turning to a behavioral finance framework, Han *et al.* (2007) highlight the role *overconfidence* might play in the FP puzzle.[16] Specifically, when some investors receive a noisy signal for next period's inflation differential between domestic and foreign markets, they overestimate its accuracy and therefore both spot and forward exchange rates overshoot (in the same direction) their equilibrium values. However, the signal is assumed to have a larger impact on the forward rate compared to the spot rate. Han *et al*. allow this important asymmetry in the model by noting that whilst the spot rate is partly a function of the transactions demand for money, the forward rate is more driven by speculation. Given this, a

[16] See Hirshleifer (2001) for a review of models on investor overconfidence.

positive (negative) contemporaneous forward premium implies a negative (positive) future spot return as the overreaction of the spot rate is corrected.

Therefore, in the Han *et al*. (2007) model, the sign of b_k depends on the strength of the overreaction. On the one hand, in a theoretical one period 'short horizon' model, overreaction is shown to generate the required negative b_k coefficient given plausible parameter values. On the other hand, examining a two period 'longer horizon' model, they show that whilst the two period coefficient is still negative, it is less so than the coefficient in the one period regression. Specifically, Han *et al*. suggest that at longer horizons, the overreaction effect is dominated by the conventional UIP effect.

3.2. A simple model of investor overreaction

Given the work by Han *et al.* (2007), a behavioral explanation for the FP puzzle is promising. Additionally, since the binary 'short' and 'long' horizon dichotomy is too restrictive, we posit an overreacting mechanism suggesting a dynamic time path for the b_k estimate. Therefore, we present a model of foreign exchange market overreaction which allows greater focus on the 'term structure of the FP puzzle' or, in other words, how swiftly it disappears.

Firstly, assume relative PPP holds on average:

$$\Delta s_t = \overline{\pi}_t + \varepsilon_t \tag{2}$$

where s_t is the logarithm of the spot exchange rate at time t and ε_t is a zero-mean error term. Additionally, the realized inflation differential is defined as $\overline{\pi}_t \equiv \pi_t - \pi_t^* = \Delta p_t - \Delta p_t^*$, with p_t representing the logarithm of the home country price level and asterisks indicating any foreign country analog.

Secondly, and similarly to Han *et al*., allow covered interest rate parity to hold and use the Fisher equation[17] so:

$$d_t = f_{t,t+1} - s_t = i_t - i_t^* = E_t(\bar{\pi}_{t+1}) \qquad (3)$$

where $f_{t,t+1}$ is the logarithm of the one-month forward exchange rate at time t, i_t is the nominal rate of return on a domestic one-month bond and E is an expectations operator. Next, let the realized inflation differential follow the error correction type process:

EMBED Equation.3 ~~(4~~)

where λ>0 to correct for ~~situations~~ the weighted average ~~***expectation is inflated by overconfidence and***~~ thus exceeds the rational expectation. U***nlike Han*** et al., we assume that all investors are risk averse and thus prices reflect a weighted average of the beliefs of different investors. Therefore:

EMBED Equation.3 (5)

where EMBED Equation.3 is the proportion of overconfident (EMBED Equation.3) investors, relative to rational investors (EMBED Equation.3), in the spot exchange market. In other words, the realized inflation differential is equal to the market expectation of the future inflation differential plus a correction to the previous period's error from overconfident investors. Hereafter, the equations in our model diverge from those in the Han et al. (2007) model due the crucial role the proportion of overconfident investors plays as the horizon of the forward contracts is extended.

Assume all investors receive a signal about next period's inflation differential at time t. Overconfident investors overreact to this signal compared to rational investors:

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[17] The Fisher equation states $i_t = r_t + E_t(\pi_{t+1})$, where r_t is the real rate of return on a domestic one-month bond. For equation (3) to hold, it is necessary to assume $r_t = r_t^*$.

[18] Since our approach does not employ a risk-based explanation, we assume a constant but negligible risk premium that we do not model explicitly. Including this would not qualitatively alter our results.

[19] For simplicity, we assume that rational investors assess without error, any signal for next period's inflation differential.

[20] Han *et al*. (2007) require only overconfident, risk neutral investors to receive a signal.

$$E_t^o(\bar{\pi}_{t+1}) = \alpha E_t^r(\bar{\pi}_{t+1}) \tag{6}$$

where, in the presence of a signal, $\alpha > 1$; otherwise $\alpha = 1$. The reaction in the forward exchange market to a signal can be assessed by denoting c' as the proportion of overconfident to rational investors. We assume that typically $c' > c$. This is because the forward market contains relatively more speculative and therefore, overconfident investors as it is cheaper to trade forward. Substituting (6) into a weighted average version of (3) gives at time t:

$$\begin{aligned} d_t &= f_{t,t+1} - s_t = E_t(\bar{\pi}_{t+1}) = c' E_t^o(\bar{\pi}_{t+1}) + (1-c')E_t^r(\bar{\pi}_{t+1}) \\ &= c'\alpha E_t^r(\bar{\pi}_{t+1}) + (1-c')E_t^r(\bar{\pi}_{t+1}) \\ &= [c'\alpha + (1-c')]E_t^r(\bar{\pi}_{t+1}) \end{aligned} \tag{7}$$

Given $0 \geq c' \geq 1$ and $\alpha > 1$, then (7) indicates that the contemporaneous forward premium is always positive (negative) in response to a positive (negative) signal about the future inflation differential. Contrastingly, in the spot market at time $t+1$, from (2) and (4) we have:

$$\Delta s_{t+1} = \bar{\pi}_{t+1} + \varepsilon_{t+1} = \omega_{t+1} + \lambda[E_t^r(\bar{\pi}_{t+1}) - \omega_t] + \varepsilon_{t+1} \tag{8}$$

where from (5) and (6):

$$\omega_{t+1} = c\alpha E_{t+1}^r(\bar{\pi}_{t+2}) + (1-c)E_{t+1}^r(\bar{\pi}_{t+2}) \tag{9}$$

However, because there is no signal at time $t+1$, $\alpha = 1$ and (9) reduces to $\omega_{t+1} = E_{t+1}^r(\bar{\pi}_{t+2})$. Additionally, because inflation differentials are persistent (and there is no new information):

$$\omega_{t+1} = E_{t+1}^r(\bar{\pi}_{t+2}) = E_t^r(\bar{\pi}_{t+1}) \tag{10}$$

Substituting (10) into (8) and simplifying:

$$\begin{aligned} \Delta s_{t+1} &= E_t^r(\bar{\pi}_{t+1}) + \lambda[E_t^r(\bar{\pi}_{t+1}) - \omega_t] + \varepsilon_{t+1} \\ &= E_t^r(\bar{\pi}_{t+1}) + \lambda[E_t^r(\bar{\pi}_{t+1}) - (cE_t^o(\bar{\pi}_{t+1}) - (1-c)E_t^r(\bar{\pi}_{t+1}))] + \varepsilon_{t+1} \\ &= [1 + \lambda c(1-\alpha)]E_t^r(\bar{\pi}_{t+1}) + \varepsilon_{t+1} \end{aligned} \tag{11}$$

Equation (11) shows that, if $1+\lambda c(1-\alpha)<0$, then the one-period ahead spot return will be negative (positive) in response to a positive (negative) signal at time *t*. Note that, while this requires a significant degree of overreaction, it is consistent with the absolutely large negative coefficients found in the literature. In other words, given appropriate levels of overconfidence in the spot market, the exchange rate will overshoot at time *t* and subsequently correct itself by $t+1$. To formalize the relationship between the one-period ahead spot return and the contemporaneous forward premium, in the presence of a future inflation differential signal, rearrange (7):

$$E_t^r(\bar{\pi}_{t+1}) = \frac{1}{c'\alpha + (1-c')} d_t \tag{12}$$

Substituting (12) into (11) gives:

$$\Delta s_{t+1} = \frac{1+\lambda c(1-\alpha)}{c'\alpha + (1-c')} d_t + \varepsilon_{t+1} \tag{13}$$

where the slope coefficient in the spot return regression is:

$$b_1 = \frac{1+\lambda c(1-\alpha)}{c'\alpha + (1-c')} \tag{14}$$

The expression on the right hand side of (14) shows that if $1+\lambda c(1-\alpha)<0$, then the slope coefficient will be negative (b_1<0), the typically observed empirical result for short horizon spot return regressions. As expected, the slope coefficient is decreasing in both the overconfidence parameter (α) and the proportion of overconfident investors (c).

$\partial\beta/\partial\alpha = -c\lambda[c'(\alpha\text{-}1)+1] - c'[1\text{-}(\alpha\text{-}1)c\lambda\ ] /[c'(\alpha\text{-}1)+1]^2$

$\partial\beta/\partial\alpha < 0$ for $c' > c(-\lambda)$

The negative response to overconfidence parameter (α) depends on both the proportion of overconfident investors (*c*) and on λ which is the error correction term in equation (4).

Similarly

$\partial\beta/\partial c = (1\text{-}\alpha)\lambda/[c'(\alpha\text{-}1)+1]^2 < 0$ for $\alpha>1$

The negative response to the proportion of overconfident investors (c) depends on α which is the overreaction parameter in equation (6) and satisfies the restriction α >1 in the presence of a signal.

Of course, we are not only interested in the short horizon but how the FP puzzle evolves as the time horizon extends. The result in (14) can be generalised to a k - period ahead model by allowing appropriate k - period analogues of equations (3) to (6):

$$d_{t,t+k} = f_{t,t+k} - s_t = i_{t,t+k} - i^*_{t,t+k} = E_t(\bar{\pi}_{t+k}) \tag{15}$$

$$\bar{\pi}_t = \omega_t + \lambda[E^r_{t-k}(\bar{\pi}_t) - \omega_{t-k}] \tag{16}$$

$$\omega_t = c_k E^o_t(\bar{\pi}_{t+k}) + (1-c_k)E^r_t(\bar{\pi}_{t+k}) \tag{17}$$

$$E^o_t(\bar{\pi}_{t+k}) = \alpha_k E^r_t(\bar{\pi}_{t+k}) \tag{18}$$

Now at time t, assume all investors receive a signal about the k period's inflation differential. In that case, the previous algebra holds and the slope coefficient is:

$$b_k = \frac{1 + c_k \lambda_k (1-\alpha_k)}{c'_k \alpha_k + (1-c'_k)} \tag{19}$$

We assume that, as the horizon increases, both the proportion of overconfident investors (c_k) and the overreaction parameter (α) in (19) will increase. As the parameter α_k approaches unity or overshooting decreases, then the slope coefficient turns positive. The restriction for the slope coefficient in (19) to be unity is that that α_k=1 or that overreaction has died out. Equation (19) suggests a 'term structure' for the forward premium puzzle in which the expectations hypothesis may not hold at short horizons but is valid at medium and longer horizons. The horizon at the slope coefficient becomes indistinguishable from unity is an empirical question and the subject of the next section.

4. Data, methodology and results

4.1. Data

Whilst forward contracts are common for maturities of up to one year, this is not true for longer horizons. Therefore, for any maturity greater than one year, forward premia have to be calculated via the covered interest parity relationship. In fact, this relationship is used to provide *all* forward premia as Maynard and Phillips (2001**)** show that interest rate differentials provide a cleaner measure of the premia than those calculated with actual forward contracts.

For the short-term rates, we employ Eurocurrency yields calculated using the number of days as described by Bauer (2001) for the following horizons: 1 month, 3 months, 6 months and 12 months. This involves calculating the exact number of days between each monthly observation when calculating monthly rates from annual returns.[21] Benchmark government bond yields are subsequently used at the 2-year, 3-year, 5-year, 7-year and 10-year horizons for the medium and long-term rates.

The dataset comprises of bilateral US dollar exchange rates for the following countries: Canada, Germany, Japan, Switzerland and the United Kingdom. Sampled at a monthly frequency, these represent the majority of the most heavily traded currencies over the sample period.[22] The *actual* period covered varies according to data availability by country, spanning 1980-2005; see Appendix A. Additionally, for robustness we use the quarterly data of Chinn and Meredith (2004).[23] These data, covering 1973-2001, include 10-year benchmark government bond yields and 5 and 10-year constant maturity government bond yields. Again, the number of countries under each yield type varies due to data availability; see Appendix B.

[21] As Bauer notes, the Eurocurrency compounding convention is the actual number of days divided by 365 for the British pound and 360 for all other currencies.
[22] For Switzerland there are no results for the 2-year, 7-year and 10-year horizons as these government bond series are not available. See BIS (1999) for representative currency trading volumes.
[23] We are grateful to Menzie Chinn and Guy Meredith for making these data available.

4.2. Results

In the first and second sub-section below, results are presented for our monthly frequency data and the quarterly data of Chinn and Meredith (2004), respectively. Typically, the extant literature estimates the forward premium regression in equation (1) and tests the null hypothesis of a unit slope coefficient. As noted earlier, our equivalent null hypothesis is a zero slope coefficient in regression (21) and therefore all such coefficients estimates are reported after adding unity for comparability purposes.

4.2.1. Monthly frequency results using HAC estimation

Preliminary estimates of (21) using HAC estimation *without bootstrapping* are shown in Table 1.

[Table 1 around here]

Table 1 indicates that all five currency pairs have significantly negative slope coefficients at the 1% critical value except for the DM which is significant at the 5% critical value. The average slope coefficient is -1.894 which is typical of the estimates reported elsewhere in the FP puzzle literature.

Some three of the five currency pairs exhibit slope coefficient estimates that are monotonically increasing up to 12 months. All of the coefficients are negative and one can reject the FRUH null of a unit slope at the 1% significance level in most cases.[24] Further examination of Table 1 indicates that the trend of increasing coefficients continues into the medium and long horizon. However, whilst the majority of the 2-year horizon coefficients remain negative, this is reversed at the 3-year horizon where the coefficients are predominantly positive. Moreover, the FRUH is supported for three countries at the 3-year horizon and never rejected at the 5% level or better for the 5-year horizon. Generally, once

[24] In the case of the DM, the results hold only at the 5% and 10% significance levels for some horizons.

support is found for the null at these intermediate horizons, it continues into the long horizon. The noticeable exception is the UK where support for FRUH is found only at the 3 and 5-year horizons.

Considering the results over the full spectrum of available horizons, there appears *prima facie* support for the effects of investor overreaction. More specifically, the initial negative point estimates reflect overshooting in the short run. Thereafter, the correction of the overreaction is reflected in the estimates turning positive and reverting towards their desired value of unity at medium horizons for all sample countries. This can be clearly seen from Figures 1 to 3:

[Figures 1-3 around here]

Inspection of these figures illustrates the nature of the mean reversion to the theoretically consistent slope coefficient of unity. The change from negative to positive coefficient values is clearly visible for all countries, though the trajectories differ markedly between currencies. This resonates with the Bekaert et al. (2004) argument that the FP puzzle varies by currency.

4.2.2 Results addressing the data overlap issue

Instead of contrasting short and long horizon results, we address a different issue. We seek to establish how soon support for the FRUH can be garnered. In other words, the question becomes at which horizon support for the FRUH first appears. However, given monthly data, the use of variables with horizons of longer than one month involves data overlap. In the presence of overlapping observations, *t*-statistics can be severely biased as asymptotic standard errors are too small in finite samples. Nelson and Kim (1993) show the effect of overlapping can dramatically shift the distribution of a *t*-statistic even when using heteroskedastic and autocorrelation consistent (HAC) estimation.

To address the overlapping issue, we deploy a bootstrap procedure similar to that of Killian and Taylor (2003) and Rapach and Wohar (2005).[25] In particular, we recast the forward premium anomaly using an analogous predictability specification to facilitate the application of the bootstrap procedure. First recall that if FRUH holds, the current *k*-period forward rate provides an unbiased forecast of the *k*-period ahead spot exchange rate:

$$E[s_{t+k} - f_{t,t+k}] = 0 \qquad (20)$$

From (20) we derive the forecast error regression:

$$(s_{t+k} - f_{t,t+k}) = \alpha_k + \beta_k(f_{t,t+k} - s_t) + \varepsilon_{t+k}$$

$$\text{or } sf_{t+k} = \alpha_k + \beta_k(fp_{t+k}) + \varepsilon_{t+k} \qquad (21)$$

for horizon *k*, where $s_{t+k} - f_{t,t+k} = sf_{t+k}$, and $f_{t,t+k} - s_t = fp_{t+k}$ and ε_{t+k} is the error term. The null hypothesis $\beta_k = 0$ is analogous to $\beta_k = b_k - 1$ in spot return regression (1).

In equation (21) the predictive ability of the forward premium, fp_{t+k}, with respect to the forward rate forecast error, sf_{t+k}, is assessed by the *t*-statistic of the OLS estimate of β_k. However, for $k > 1$, the observations of the endogenous variable overlap, inducing serial correlation in the error term, ε_{t+k}. In a first stage, this is addressed using HAC estimation, applying the Newey and West (1987) adjustment using the Bartlett kernel. In a second stage, the bootstrap procedure can be defined on the basis of the HAC estimated regression. Specifically, for a given horizon, the data generating processes (DGPs) are defined for sf_{t+k} and fp_{t+k} under the null of predictability, H_0: $\beta_k = 0$. In our case then, sf_{t+k} is generated by a random walk with drift and fp_{t+k} by an AR model, with the lag order selected by the AIC up to a maximum of order 5:

$$sf_{t+k} = \alpha_k + \varepsilon_{t+k} \qquad (22)$$

[25] Note that Killian and Taylor (2003) and Rapach and Wohar (2005) examine predictability in foreign exchange markets and stock markets respectively. For other literature, see Mark (1995), Kothari and Shanken (1997), Killian (1999) and Faust et al. (2003).

$$fp_{t+k} = AR(AICMax[5]) \tag{23}$$

DGPs (22) and (23) are fitted by OLS and residual series and parameter estimates obtained. Using these parameter estimates, a pseudo sample is built by sampling in tandem from the OLS residuals, thus preserving the structure of the contemporaneous correlation in the original data. This sampling is carried out with replacement and the initial 100 start-up observations are discarded. Using the pseudo data, we estimate equation (21) via our HAC estimator and store the *t*-statistic. The formation of the pseudo data, estimation and storage of the *t*-statistic is repeated 2000 times to construct the empirical distribution of *t*-statistics. Finally, to generate the *p*-values under the null, we calculate the absolute proportion of the *t*-statistics that are less than the *t*-statistic observed using the original data.

Table 2 presents the bootstrapped *p*-values based on a pseudo-distribution of HAC *t*-statistics:

[Table 2 around here]

Again, these results clearly show a rejection of the null hypothesis at the shorter horizons and support for the null at medium and long horizons. Of particular interest is the location of the interim horizon at which start to find support for the FRUH begins to appear. Considering the 10% significance level and the bootstrapped *p*-values, support for an unbiased forward rate emerges at horizons of 3 years or greater for Canada. In the case of Germany, the same support appears at horizons of 2 years or greater. For Switzerland support is found at the 5 year horizon. In all these cases, the HAC *t*-statistic and bootstrapped *p*-value results are qualitatively the same.

Interestingly, there are clear differences between the bootstrapped and non-bootstrapped results for the case of Japan and the UK. Based solely on the HAC *t*-statistic, support for an unbiased forward rate can only be found at the 7 and 10-year horizons for Japan, and the 3 and 5-year horizons for the UK. By contrast, using the bootstrapped *p*-

values, support emerges at horizons of 5 years or greater for Japan and 3 years or greater for the UK. The bootstrap *p*-values overturn the anomalous UK result found under non-bootstrapped HAC *t*-statistics by suggesting that the British pound supports the FRUH at both medium and long horizons rather than the former alone.

Overall, results indicate that support for FRUH emerges at the 3 to 5-year horizon. Specifically, support is found as early as the 2-year horizon for the DM, for 3 out of 5 currencies at the 3-year horizon, and for all currencies at longer horizons. This result is quite striking. As predicted by the behavioral finance model presented in the previous section, the FP puzzle disappears at medium horizons and beyond.

4.2.3 Chinn and Meredith (2004) data

As a comparison, the methodology used in the previous section is applied to the quarterly dataset of Chinn and Meredith (2004). Table 3 presents the estimation results of equation (21) for the 10-year benchmark and constant maturity 5 and 10-year government bond yields.[26]

[Table 3 around here]

Using 10-year benchmark data (Panel 3a) it is not possible to reject the null of an unbiased forward rate for Canada, France and Germany based on the non-bootstrapped HAC *t*-test. However, FRUH is roundly rejected for Italy, Japan and the UK. These results are qualitatively the same as Chinn and Meredith (2004), rejecting the null hypothesis for analogous countries. For the 5 and 10-year constant maturity government bonds, more non-bootstrapped HAC regressions (Panels 3b-3c) yield estimated coefficients where the null hypothesis is not rejected in any case, with the exception of Japan. Again, this is similar to Chinn and Meredith (2004), where findings cannot support unbiasedness for Japan or the

[26] Once again, the estimated slope coefficients of (21) are presented after adding unity.

UK. Turning to the bootstrap p-values in Panels 3a-3c indicate that support for the unity null increases and there are no rejections at the 5% level or better. Applying the bootstrap procedure clearly provides even more support for an unbiased forward rate at long horizons.

5. Conclusions

This paper investigates the nature of the forward premium anomaly across a spectrum of horizons from the short to the long run. A behavioral model of exchange rates is developed that draws on the Frankel and Froot (1991) distinction between rational and chartist investors and recent theoretical work by Han *et al*., (2007) that examines the role of overconfident investors. The model assumes a relatively high proportion of overconfident to rational investors at short horizons that generates the negative slope coefficients typically found in the extant literature. As the horizon increases, the proportion of overconfident investors decreases and overreaction tends towards zero. This causes the slope coefficient to approach its long-run theoretical value of unity. In essence, the model suggests a 'term structure' for the forward premium where the expectations hypothesis fails in the short run due to overshooting but holds at medium and long horizons. It highlights the empirical question of how quickly unbiasedness is achieved.

Employing a monthly sampled dataset of bilateral US dollar exchange rates for the most heavily traded currencies (Canada, Germany, Japan, Switzerland and the United Kingdom; spanning 1980-2005), the dynamic time path of the forward premium puzzle is examined after accounting for the data-overlap problems inherent in intermediate and long horizon studies. Specifically, a HAC-bootstrap approach is used to correct for the resultant shift in the relevant t-distribution. Results reveal estimated slope coefficients congruent with the behavioral model. In particular, the HAC-bootstrap procedure provides clear evidence that forward rate unbiasedness commonly emerges at around the 3-year horizon!

Appendix A

Table A1 shows the start and end dates for the monthly data.

Table A1: Dataset coverage

	Start	End	No. observations
Canada	February 1985	December 2004	239
Germany	January 1980	December 1998	228
Japan	January 1984	December 2004	252
Switzerland	January 1981	December 2004	288
UK	January 1980	December 2004	300

Source: Thomson DataStream

Appendix B

Table B1 shows the data availability at the country level for the 10-year benchmark, 5-year constant maturity, and 10-year constant maturity government bonds. These data are quarterly, from 1973:Q1-2004Q4, with the exception of Italy (1977:Q1-2000Q4).

Table B1: Dataset coverage

10-year benchmark government bonds	5-year constant maturity	10-year constant maturity
Canada	Canada	Canada
France	Germany	Germany
Germany	UK	Japan
Italy		UK
Japan		
UK		

Source: Chinn & Meredith (2004)

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Table 1: Tests for an unbiased forward rate: non-bootstrapping approach

	1 month	3 months	6 months	12 months	2 years	3 years	5 years	7 years	10 years
Canada	-1.3220 (0.5399***)	-1.1591 (0.4523***)	-0.9790 (0.4013***)	-0.9723 (0.4265***)	-0.5675 (0.4850***)	0.2664 (0.4711)	0.5964 (0.4741)	1.2749 (0.2623)	0.7450 (0.3339)
Germany	-0.9484 (0.9763**)	-0.6856 (0.8654*)	-0.8454 (0.7283**)	-0.5627 (0.5696***)	0.1302 (0.5321)	0.3942 (0.5459)	1.0309 (0.2808)	1.3209 (0.2160)	1.4184 (0.3512)
Japan	-2.9044 (1.1570***)	-3.1562 (1.047***)	-3.1789 (0.8532***)	-3.1552 (0.6062***)	-2.5651 (0.6002***)	-1.6844 (0.4909***)	-0.0159 (0.5200*)	0.9742 (0.3235)	1.4954 (0.4678)
Switzerland	-1.6309 (0.8586***)	-1.4431 (0.8949***)	-1.3120 (0.7591***)	-1.1378 (0.5761***)	-	-0.2135 (0.4293***)	0.9456 (0.1953)	-	-
UK	-2.6641 (1.0897***)	-2.4370 (1.1285***)	-2.1302 (1.0481***)	-1.4934 (0.8596***)	-0.9091 (0.7266***)	0.1670 (0.5730)	0.5680 (0.4314)	0.5065 (0.2386**)	0.5136 (0.1850***)

Notes: Point estimates are from regression (21). HAC standard errors are in parenthesis. *, **, *** indicates different from null of one at the 10%, 5% and 1% marginal significance level respectively.

Table 2: Tests for an unbiased forward rate: bootstrap *p*-values approach

Canada	Horizon	1 month	3 months	6 months	12 months	2 years	3 years	5 years	7 years	10 years
	Slope	-1.3220	-1.1591	-0.9790	-0.9723	-0.5675	0.2664	0.5964	1.2749	0.7450
	BS p-val	0.0005	0.0005	0.0005	0.0010	0.0170	0.2460	0.5805	0.5005	0.6265
	HAC Significant	***	***	***	***	***				
	p-val Significant	†	†	†	†	†				

Germany	Horizon	1 month	3 months	6 months	12 months	2 years	3 years	5 years	7 years	10 years
	Slope	-0.9484	-0.6856	-0.8454	0.5627	0.1302	0.3942	1.0309	1.3209	1.4184
	BS p-val	0.0980	0.0790	0.0310	0.0180	0.1710	0.3420	0.9240	0.2095	0.4165
	HAC Significant	**	*	**	***					
	p-val Significant	†	†	†	†					

Japan	Horizon	1 month	3 months	6 months	12 months	2 years	3 years	5 years	7 years	10 years
	Slope	-2.9044	-3.1562	-3.1789	-3.1552	-2.5651	-1.6844	-0.0159	0.9742	1.4954
	BS p-val	0.0025	0.0005	0.0000	0.0000	0.0000	0.0000	0.1790	0.9580	0.4215
	HAC Significant	***	***	***	***	***	***	*		
	p-val Significant	†	†	†	†	†	†			

Switzerland	Horizon	1 month	3 months	6 months	12 month	2 years	3 years	5 years	7 years	10 years
	Slope	-1.6309	-1.4431	-1.3120	-1.1378	-	-0.2135	0.9456	-	-
	BS p-val	0.0265	0.0225	0.0105	0.0025	-	0.0210	0.797	-	-
	HAC Significant	***	***	***	***	-	***		-	-
	p-val Significant	†	†	†	†	-	†		-	-

UK	Horizon	1 month	3 months	6 months	12 months	2 years	3 years	5 years	7 years	10 years
	Slope	-2.6641	-2.437	-2.1302	-1.4934	-0.9091	0.1670	0.5680	0.5065	0.5136
	BS p-val	0.0030	0.0475	0.0650	0.0445	0.0735	0.4360	0.7385	0.2250	0.2680
	HAC Significant	***	***	***	***	***			**	***
	p-val Significant	†	†	†	†	†				

Notes: Point estimates are from regression (21). *, **, *** indicates different from null of one at the 10%, 5%, and 1% marginal significance level respectively. † indicates a bootstrapped (BS) *p*-value less than 10%.

Table 3: Long horizon tests for an unbiased forward rate: Chinn and Meredith (2004) quarterly data

Panel 3a. Benchmark Government Bond Yields, 10-Year Maturity

	$\widehat{\beta}_k$		BS *p*-value	N
Canada	1.1269	(0.2845)	0.6980	72
France	0.8258	(0.6693)	0.8380	72
Germany	0.9105	(0.1967)	0.7515	72
Italy	0.1870	(0.3002***)	0.1735	56
Japan	0.3808	(0.2216***)	0.0525	72
UK	0.5647	(0.1425***)	0.0785	72

Notes: Point estimates are from the regression in equation (21). HAC standard errors are in parenthesis. *, **, *** indicates different from null of one at the 10%, 5%, and 1% marginal significance level respectively. Bootstrapped (BS) *p*-values in the fourth column. Data period 1973:Q1-2000:Q4 except Italy (1977:Q1-2000Q4).

Panel 3b. Constant Maturity Government Bond Yields, 5-Year Maturity

	$\widehat{\beta}_k$		BS *p*-value	N
Canada	0.1631	(0.5393)	0.5020	92
Germany	0.2670	(0.5108)	0.2880	92
UK	0.4635	(0.4158)	0.6605	92

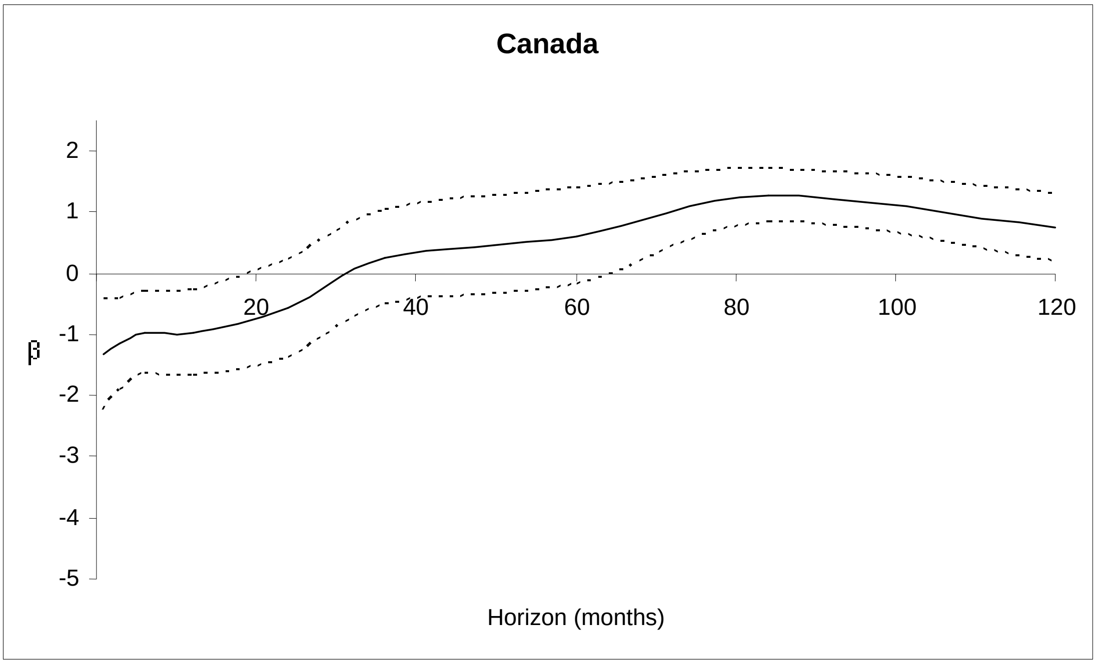
Notes: Point estimates are from the regression in equation (21). HAC standard errors are in parenthesis. *, **, *** indicates different from null of one at the 10%, 5%, and 1% marginal significance level respectively. Bootstrapped (BS) *p*-values in the fourth column. Data period 1973:Q1-2000:Q4.

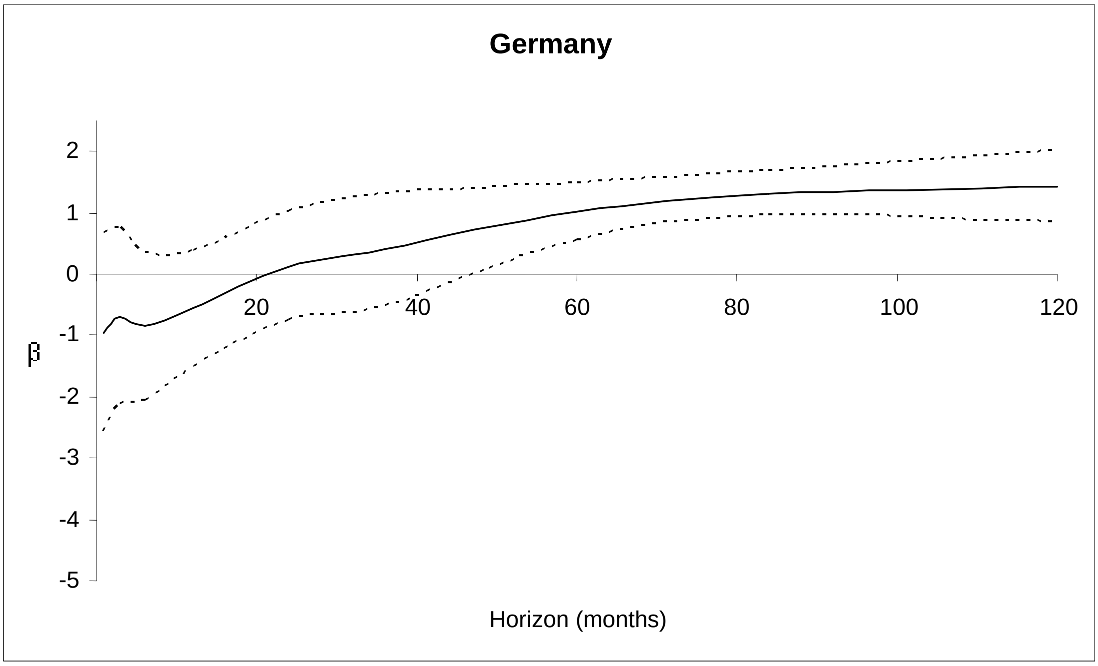
Panel 3c. Constant Maturity Government Bond Yields, 10-Year Maturity

	$\widehat{\beta}_k$		BS *p*-value	N
Canada	0.6048	(0.3293)	0.5735	72
Germany	0.9052	(0.1863)	0.7285	72
Japan	0.4117	(0.2672**)	0.1395	72
UK	0.7110	(0.1762)	0.6520	72

Notes: Point estimates are from the regression in equation (1). HAC standard errors are in parenthesis. *, **, *** indicates different from null of one at the 10%, 5%, and 1% marginal significance level respectively. Bootstrapped (BS) *p*-values in the fourth column. Data period 1973:Q1-2000:Q4.

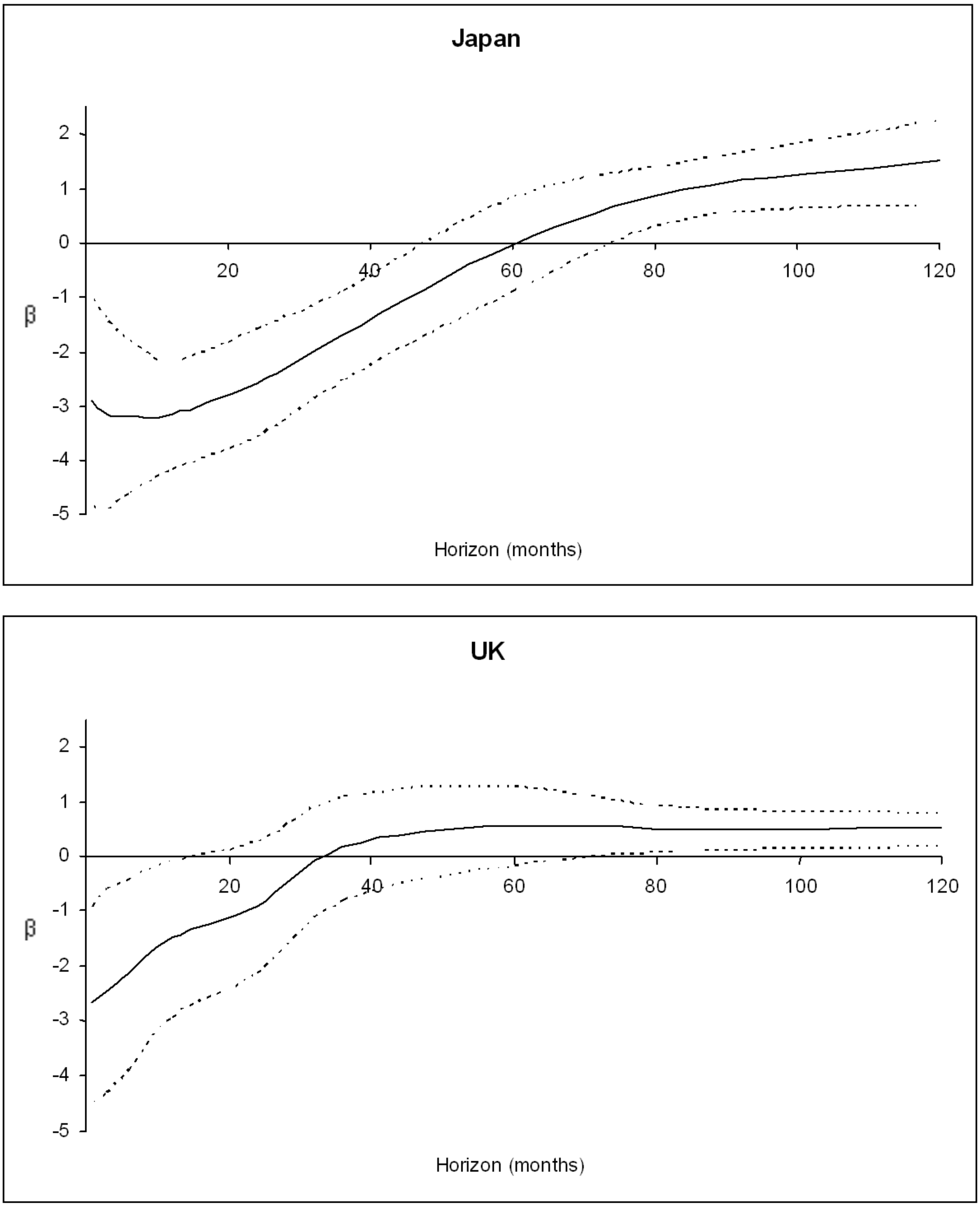
Figure 1: Equation (21) slope coefficient estimates with 10% rejection bands – Canada and Germany





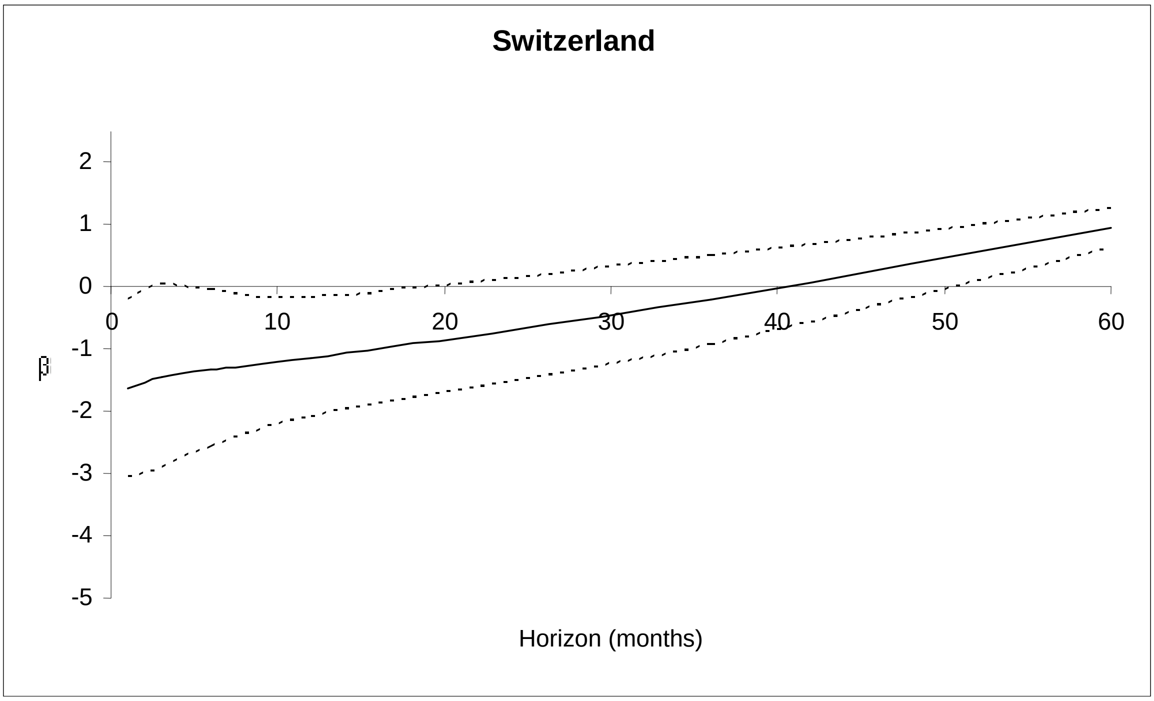
Notes: The solid line represents the point estimates from Tables 1 and 2. The dashed lines represent rejection bands at the 10% significance level.

Figure 2: Equation (21) slope coefficient estimates with 10% rejection bands – Japan and UK



Notes: The solid line represents the point estimates from Tables 1 and 2. The dashed lines represent rejection bands at the 10% significance level.

Figure 3: Equation (21) slope coefficient estimates with 10% rejection bands over all horizons – Switzerland



Notes: The solid line represents the point estimates from Tables 1 and 2. The dashed lines represent rejection bands at the 10% significance level.