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Exchange rates and long-term bonds*

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Abstract

Tentative evidence suggests that the empirical failure of uncovered interest parity (UIP) is confined to short-term interest rates. Tests of UIP for long-term interest rates are however hampered by various data problems. By focusing on short investments in long-term bonds, these data problems can be avoided. We study the relationship between the US dollar - Deutsch Mark exchange rate and German and American bond rates. The hypothesis that expected returns to investments in bonds denominated in the two currencies are equal cannot be rejected. This result is not simply due to low power as the β -coefficients are close to unity. For the corresponding short-term interest rates, the typical finding of a large and significantly negative β -coefficient is confirmed.

Keywords: Long-term interest rates, exchange rates, uncovered interest parity

JEL classifications: F31, F41

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1 Introduction

Uncovered interest parity (UIP) is typically soundly rejected in empirical tests. In contrast to the unity coefficient expected from UIP, numerous studies have reported β -coefficients that are significantly negative, and large. High interest rate currencies hence tend to appreciate rather than depreciate. This is the exchange rate risk premium puzzle or forward premium puzzle. A striking characteristic of the empirical literature on UIP is the exclusive focus on short-term interest rates. Before elevating the empirical failure of UIP to a stylized fact, *long-term* interest rates should also be investigated. Two recent papers that attempt to fill this void in the literature are Alexius (2001) and Meredith and Chinn (1998). They both test UIP for long investments in long-term bonds and conclude that the β -coefficients are typically significantly positive, but also significantly smaller than one. Similarly, Flood and Taylor (1996) run a UIP test for three-year bonds and obtain a positive coefficient.¹ Hence, the few existing results for long-term interest rates are much more favorable to UIP than the standard findings for short-term interest rates. The hypothesis of a *unity* β -coefficient is nonetheless rejected more often than not also for long-term interest rates.

The absence of UIP tests for long-term interest rates is probably due to data problems. This paper explores a possible route around these data problems, namely to focus on the relationship between nominal exchange rate changes and returns to short investments in long-term bonds. We study short (one to 32 weeks) investments in US and German ten-year benchmark

¹ Flood and Taylor (1996) is a survey of empirical exchange rate economics. Among other things, it includes regressions of three-year interest rates on corresponding exchange rate changes.

government bonds and corresponding changes in the US dollar - Deutsche Mark exchange rate. The hypothesis that expected returns are equal is not rejected and the β -coefficient is close to unity. We also match the data on exchange rate changes to corresponding short-term interest differentials. In this case, the standard finding of a negative and large β -coefficient is confirmed.

It is difficult to test UIP for long-term interest rates for two reasons. Large parts of the return to investments in long-term bonds stem from coupon payments that are made before maturity. This is not a technicality but a quantitatively important problem. For instance, 100 percent of the profit from holding a par bond to maturity takes the form of coupon payments. Sophisticated techniques for removing the effects of coupon payments are available, but they cannot be applied to the data used in Alexius (2001) or Meredith and Chinn (1998) because required information about e.g. the term structure of interest rates is not available for these long time series and/or wide range of countries.

A second problem is that information about the maturity of the long-term bonds is inexact. An observation designated as a ten-year bond could well be an eight-, nine- or eleven-year bond. This induces a timing mismatch between e.g. ten-year exchange rate changes and the bond investments. The size of the resulting measurement error depends on the slope of the yield curve at long maturities. If this segment of the yield curve is completely flat, nine-year interest rates and ten-year interest rates are equal and it does not matter whether a ten-year exchange rate change is matched to a nine-year interest rate or a ten-year interest rate. The measurement errors are large if

the yield curves are steep and have different slopes in different countries.

The combined effect of the two types of measurement errors in the data on long-term interest rates could bias the results from the UIP tests in either direction. Since coupon payments are typically positively high in times of high nominal interest rates, the measurement errors due to the presence of coupon payments work in favor of a positive β -coefficient. On the other hand, to the extent that the measurement errors are random, the β -coefficient is biased towards zero. UIP could hold even better for long-term interest rates than what is documented in Alexius (2001) and Meredith and Chinn (1998) – or the positive β -coefficients found in these two studies could merely be a consequence of systematic measurement errors. Alexius (2001) uses two crude methods to remove the effects of coupon payments from the data on long-term interest rates but is left with imprecise information about maturity. Her results indicate that UIP fares better the more carefully the data on returns to investments are constructed. Meredith and Chinn (1998) are able to obtain data on synthetic constant maturity bonds for some countries, hence avoiding the second problem, but they do not attempt to deal with the presence of coupon payments. Again, their results are more favorable to UIP when more accurate data on bond returns are used.

For the recent decades, the quality and availability of data on returns to investments in long-term bond is satisfactory. Long time series are however essential when testing UIP for long-term interest rates. Even the 40 years of data on ten-year bonds used in Alexius (2001) only contain three non-overlapping observations on realized exchange rate changes.

More and better studies of the relationship between long-term interest

rates and exchange rates are obviously needed. It is less obvious how to proceed given that better data on bond yields cannot be collected in retrospect. By focusing on returns to short investments in long-term bonds, a large number of observations can be constructed using only the recent period for which high quality data are readily available. The cost of this strategy is that while UIP is defined as a relationship between expected exchange rate changes and two deterministic interest rates, the returns to short investments in long-term bonds are stochastic. What can be tested is whether expected returns to investments in long-term bonds denominated in different currencies are equal given the exchange rate movements. This hypothesis is closely related to, but not equivalent with, UIP.

The finding that UIP holds for long-term interest rates but not for short-term interest rates could be a consequence of the long investment horizons used in Flood and Taylor (1996), Alexius (2001), and Meredith and Chinn (1998). Holding periods in studies of short-term interest rates are always short, three months in the case of three-months interest rates and so on. If it takes time before fundamental relationships affect exchange rates, long investment horizons may be needed to discover that the fundamental UIP hypothesis holds. Flood and Taylor (1996) interpret their results in this manner: “Fundamental things apply as time goes by.” There is however a possible alternative explanation. The relationship between short-term interest rates and *ex post* exchange rate changes could be special and hence different from the relationship between long-term interest rates and *ex post* exchange rate changes. Short-term interest rates differ from other financial assets in that they constitute the main monetary policy instrument in most

industrialized countries with flexible exchange rates. Several authors have tried to explain the negative coefficients in UIP tests in terms of the endogenous response of monetary policy to shocks. For instance, McCallum (1994), Meredith and Chinn (1998), Kugler (2000), and Alexius (2000) construct models where the co-movements of short interest rates and exchange rates in response to shocks generates negative coefficients in standard UIP tests. The finding that UIP holds for long interest rates but not for short interest rates would then be due to the maturity of the instrument *per se* rather than to the length of the investment horizon.

If the length of the investment horizon determined the relationship between exchange rates and interest rates, the results from UIP tests using long-term investments in e.g. T-bills would coincide with those using long-term investments in long-term bonds. In this paper, we study identical (short) investment horizons for long-term bonds and corresponding short-term interest rates. Thereby, we are able to distinguish between the two explanations. We find β -coefficients of +1 for short investments in long-term bonds and β -coefficients of -3 for corresponding short-term interest rates. This implies that it is the maturity of the instrument rather than the length of the investment horizon that matters for the results.

2 Theoretical framework

In contrast to the short-term interest rates typically used in UIP tests, short investments in long-term bonds are risky because the investment horizon does not coincide with the maturity of the bonds. Hence, we will consider

investments in bonds, whose returns are risky in their respective currencies of denomination. Assuming no arbitrage, a stochastic discount factor that prices the bonds can be derived. Employing the (nominal) stochastic discount factors of domestic and foreign assets, M_t and M_t^f , conditional on the information set Ω_t we have

$$1 = E \left[M_{t+\tau} R_{t+\tau}^d \mid \Omega_t \right] \quad \text{and} \quad 1 = E \left[M_{t+\tau}^f R_{t+\tau}^f \mid \Omega_t \right], \quad (1)$$

where $R_{t+\tau}^d$ is the nominal gross return on the domestic T -period bond over a τ -period horizon and $R_{t+\tau}^f$ is the corresponding nominal gross return on the foreign bond. For instance, letting $M_{t+\tau}$ denote the intertemporal marginal rate of substitution between periods t and τ , (1) is the first order conditions of the consumption capital asset pricing model.

We then price the foreign bond in terms of the domestic currency:

$$1 = E \left[M_{t+\tau} R_{t+\tau}^f \frac{S_{t+\tau}}{S_t} \mid \Omega_t \right], \quad (2)$$

where S_t is the price of foreign currency in terms of domestic currency. We let lowercase letters denote logarithms and assume that all variables are log-normally distributed. Using equation (1) and taking logs of (2) we obtain the expression

$$E \left[r_{t+\tau} \mid \Omega_t \right] = -\frac{1}{2} Var \left(r_{t+\tau} \mid \Omega_t \right) - Cov \left(m_{t+\tau}, r_{t+\tau} \mid \Omega_t \right), \quad (3)$$

where $r_{t+\tau} = r_{t+\tau}^d - r_{t+\tau}^f + s_{t+\tau} - s_t$.

The first term on the right hand side in (3) is the Jensen inequality term (JIT). A number of papers have studied the size of the JIT and found it to be negligible (Engel (1984), Cumby (1988), Hodrick (1989), and Backus, Gregory and Telmer (1993)). In a consumption based model where agents

exhibit constant relative risk aversion, the last term on the right hand side could (ignoring inflation) be rewritten as $-\gamma Cov(c_{t+\tau}, r_{t+\tau} \mid \Omega_t)$, where $c_{t+\tau}$ is the rate of growth of consumption for the period $(t, t + \tau)$ and γ is the coefficient of relative risk aversion. A risky asset with a positive risk premium has low returns in times of high marginal utility (low consumption). With risk-neutral agents ($\gamma = 0$) the covariance term would disappear from the right hand side of (3). Thus, this covariance term is the risk premium, comprising the net of the term premia for the two bonds as well as a foreign exchange premia.

Assuming rational expectations with respect to the information set $Z_t \subset \Omega_t$ and disregarding the JIT, we can derive the following expression from equation (3):

$$s_{t+\tau} - s_t = r_{t+\tau}^d - r_{t+\tau}^f + Cov(-m_{t+\tau}, r_{t+\tau} \mid Z_t) + u_{t+\tau}, \quad (4)$$

where $E[u_{t+\tau} \mid Z_t] = 0$. Assuming that there exists a stochastic discount factor such that the risk premium (the covariance term in (4)) is equal to a constant p for all t , we arrive at

$$s_{t+\tau} - s_t = p + r_{t+\tau}^d - r_{t+\tau}^f + u_{t+\tau}. \quad (5)$$

Equation (5) implies that expected returns to investments in bonds denominated in the two currencies are equal, possibly allowing for a constant risk premium. This hypothesis is closely related to UIP.

3 Data and empirical results

Data on returns to weekly investments in ten-year US and German benchmark government bonds have been constructed by Dahlquist, Hördahl and Sellin (2000). The interest rates (yields to maturity) are collected approximately at closing time of the European markets on Tuesdays (Wednesdays if Tuesdays are holidays). The presence of coupon payments is handled using the Nelson and Siegel (1987) approach. Returns at horizons above one week are obtained by summation. Matching data on the USD/DEM exchange rate are collected from the BIS database, as are corresponding data on risky investments in short-term interest rates consists of rolling overnight interest rates. The sample period is October 1993 to November 1998.

We regress *ex post* exchange rate changes on relative bond returns as in (6) and investigate whether $[\alpha, \beta]$ equals $[0, 1]$ for different choices of investment horizon τ . Alternatively, a constant risk premium is allowed and only the hypothesis that β equals one is tested.

$$s_{t+\tau} - s_t = \alpha + \beta \left(r_{t+\tau}^d - r_{t+\tau}^f \right) + u_{t+\tau}. \quad (6)$$

Since the error term $u_{t+\tau}$ and the realized return $r_{t+\tau}^d - r_{t+\tau}^f$ are simultaneously determined, (6) is estimated using GMM. It is important that the instruments are conditioned only on information available at t , i.e. that they are unaffected by events occurring between t and $t+\tau$. The weekly frequency of the data is a second impediment. We use three instruments: Zero coupon interest rate differentials (observed in t), lagged bond returns $r_{t-\tau}^d - r_{t-\tau}^f$, and short-term interest differentials (also observed in t). The model does not display obvious signs of misspecification. The LM tests indicate first or-

der autocorrelation in the residuals at the weekly, non-overlapping horizon. The Ljung-Box tests for higher order autocorrelation (16 lags) is however insignificant (18.04), as are the LM tests for second and higher order autocorrelation. The Engle (1982) test also indicates some heteroscedasticity. For holding periods above one week, we have overlapping data, which induces $MA(\tau - 1)$ autocorrelation. The weighting matrix in the GMM estimation therefore allows for heteroscedasticity and autocorrelation along the lines of Newey and West (1987). The lag length is set to one in the weekly data and to $(\tau - 1)$ when the investment horizon τ exceeds one week.

Table 1 shows the results from applying GMM to (6) for the returns to investments in long-term government bonds as the investment horizon is extended from one to 30 weeks. Partial R^2 are around 0.4 and the Sargan test indicates that the instruments are uncorrelated with the error terms at all except two investment horizons. Rather than re-optimizing the choice of instruments for each investment horizon, we present the results from using alternative sets of instruments as well. At least one combination of instruments pass the Sargan test for every investment horizon.

As shown in the second column of Table 1, all the intercepts are small and insignificantly different from zero. Hence, neither currency has carried a constant risk premium over the sample period. The third column contains the slope coefficients. For weekly investments in long-term bonds, the point estimate of β is 0.11 and insignificant. As the holding period is extended, the β -coefficient rises up to a maximum of 1.423 at investment horizons of 11 weeks. The strict version of our hypothesis, $[\alpha, \beta] = [0, 1]$, cannot be rejected for any investment horizon. For horizons of 2 weeks or longer, the point

estimates of β are above 0.6 and insignificantly different from unity. These findings contrast blatantly with the typical finding of a significantly negative relationship between exchange rate changes and interest differentials. Here, the hypothesis that the returns are equal cannot be rejected. Furthermore, the results are not simply due to low power to reject the null hypothesis as the point estimates of the β -coefficient are close to unity. The β -coefficients and 95 percent confidence intervals for the baseline set of instruments appear in Figure 1.

The Hausman test (not reported) indicates that it is appropriate to use instruments. We nevertheless present the results from the no-instrument case in order to demonstrate that the findings are robust. Figure 2 shows the point-estimates of β using two alternative sets of instruments (the two bond market variables and zero coupon interest rates only) as well as the original bond returns $r_{t+\tau}^d - r_{t+\tau}^f$. The point estimates of $[\alpha, \beta]$ typically differ significantly between the instrument sets, but the qualitative results from an economist's perspective remain the same across the choices of instruments. Almost all point estimates of β are significantly positive and the hypothesis of a unity coefficient can rarely be rejected. Given this main finding, several qualifying observations can be made in light of Figures 1 and 2. The point estimates of β are lowest for weekly investments and tend to peak around the quarterly horizon. They are higher in the benchmark model and the no-instrument case than for the two alternative sets of instruments. When the two long-term bond market variables are used, β falls towards 0.3 as the investment horizon is extended to six months. The two most stable curves stem from the no-instrument case and from using only zero coupon

Table 1: Results for long-term bond yields.

horizon	α	β	$H_0 : \beta = 1$	$H_0 : [\alpha, \beta] = [0, 1]$	Sargan
1 week	-0.000 [-0.304]	0.121 [0.258]	4.357 (0.037)	5.033 (0.081)	1.400 (0.497)
2 weeks	-0.001 [-0.370]	0.619 [1.404]	0.747 (0.388)	0.903 (0.637)	9.560 (0.008)
4 weeks	-0.001 [-0.239]	0.683 [1.821]	0.714 (0.398)	0.787 (0.675)	4.144 (0.126)
8 weeks	0.001 [0.243]	1.121 [3.003]	0.104 (0.747)	0.157 (0.925)	4.535 (0.104)
12 weeks	-0.003 [-0.418]	1.423 [2.762]	0.460 (0.498)	0.496 (0.780)	3.971 (0.137)
20 weeks	0.001 [0.096]	1.189 [2.565]	0.242 (0.886)	0.166 (0.683)	3.218 (0.200)
30 weeks	-0.004 [-0.306]	0.805 [4.962]	1.441 (0.230)	2.230 (0.317)	4.086 (0.130)

Tested equation: $s_{t+\tau} - s_t = \alpha + \beta \left(r_{t+\tau}^d - r_{t+\tau}^f \right) + u_{t+\tau}$.

t -statistics within brackets, p -values within parentheses. The final column contains the Sargan test of the overidentifying restrictions that the instruments are uncorrelated with the error terms. The test statistics has a $\chi^2(2)$ distribution.

yields as instrument. Adding two instruments whose correlation with the independent variable is relatively low hence introduces some volatility. The standard errors of the β -coefficients are much larger when instruments are used than with original bond returns $r_{t+\tau}^d - r_{t+\tau}^f$.

The results in Table 1 indicate that the relationship between long-term bonds and exchange rate changes is roughly consistent with standard asset pricing theory also for short investment horizons. Even the longest, semi-annual, investment horizons studied here would be classified as short within the empirical literature on UIP, and in particular relative to the five- and ten year horizons used in Alexius (2001) and Meredith and Chinn (1998). Be-

Table 2: Results for short-term interest rates

horizon	α	β	$H_0 : \beta = 1$	$H_0 : [\alpha, \beta] = [0, 1]$
1 week	0.000 [0.013]	-0.467 [-0.311]	0.955 (0.323)	1.280 (0.527)
2 weeks	-0.002 [-0.979]	-1.939 [-0.987]	2.239 (0.135)	3.967 (0.312)
4 weeks	-0.005 [-1.729]	-3.153 [-1.641]	4.671 (0.031)	5.464 (0.065)
8 weeks	-0.011 [-1.805]	-3.923 [-2.006]	6.336 (0.012)	7.015 (0.030)
12 weeks	-0.021 [-2.107]	-4.793 [-2.515]	9.241 (0.002)	9.838 (0.007)
20 weeks	-0.038 [-2.723]	-5.156 [-3.231]	14.877 (0.000)	15.835 (0.000)
30 weeks	-0.069 [-6.712]	-3.896 [-3.863]	23.568 (0.000)	103.195 (0.000)

Tested equation: $s_{t+\tau} - s_t = \alpha + \beta \left(r_{t+\tau}^d - r_{t+\tau}^f \right) + u_{t+\tau}$.
 t -statistics within brackets, p -values within parentheses.

cause the β -coefficients are positive and close to unity for short investments in long-term bonds, it appears to be the maturity of the instrument and not the length of the investment horizons that matters for the results. Additional light can be shed on this issue by repeating the exercise using corresponding data on short-term interest rates, i.e. by matching the data on exchange rate changes to short-term interest differentials instead of relative returns to bond investments. Because bond investments are risky, we want corresponding risky investments in a short instrument over the same horizons. Rolling investments in overnight interest rates is a good alternative as they are risky for these horizons, readily available and easily computed for different holding periods.

Table 2 shows the results from estimating (6) for short-term interest rates using lagged short-term interest differentials as instrument. This appears to be the most appropriate choice here since adding more instruments with a weak relationship to the independent variable mainly increases the standard errors and introduces considerable volatility in the estimates of β . The qualitative results are however robust to the choice of instruments in the sense that the β -coefficient remains negative and large but significant only for investment horizons above one or two months.

The residuals from the weekly regressions using non-overlapping data display significant first (but not higher) order autocorrelation. They are also slightly heteroscedastic. Again, overlapping data induce $MA(\tau - 1)$ autocorrelation for investment horizons above one week. Hence, the GMM standard errors are corrected for heteroscedasticity, $MA(\tau - 1)$ autocorrelation for τ larger than one and $MA(1)$ autocorrelation in the weekly regressions using non-overlapping data. As shown in the third column of Table 2, the standard finding of a negative β -coefficient is confirmed for the short-term interest rates. The β -coefficient approaches -4 as the investment horizon increases, which is consistent with the typical finding. It is not significantly different from zero for weekly or fortnightly investments but becomes significant at the four week horizon.

The β -coefficients and the 95 percent confidence intervals from the UIP tests for short-term interest rates as the investment horizon is extended from one to 30 weeks are shown in Figure 3. The difference between these findings and the results for long-term interest rates in Figures 1 and 2 is striking, especially for investments above one month.

Traditional explanations for the well-documented negative relationship between exchange rate changes and interest differentials focus either on time varying risk premia, expectational errors and/or “peso problems” (broadly interpreted to include switches between appreciating and depreciating regimes for nominal exchange rates). The standard approaches do not appear likely to provide an explanation for the result that expected returns to investments in long-term bonds denominated in different currencies are equal while the β -coefficient is negative and large for the corresponding short-term interest rates. An important characteristic of short-term interest rates is that they are used as the principal instrument of monetary policy in most industrialized countries with flexible exchange rates. The approach with the greatest potential to explain the present findings appears to be models where monetary policy is conducted in a manner that creates negative co-movements of short-term interest rates and exchange rates in response to shocks. McCallum (1994), Meredith and Chinn (1998) and Alexius (2000) provide examples of how such models can be constructed.

4 Concluding remarks

The stylized fact that UIP fails in empirical tests has actually been established for short-term interest rates exclusively. The few existing studies of data on long-term interest rates lend considerable support to the hypothesis. It is however difficult to test UIP for long-term bonds because long time series of high quality data on long-term interest rates are unavailable. We focus on a short time series of carefully constructed returns to *short* investments

in long-term bonds. Thereby, a large amount of independent observations can be obtained using only the recent period, when high quality data can be found. The cost of this strategy is that while UIP is traditionally defined as a relationship between expected exchange rate changes and known interest rates, returns to short investments in long-term bonds are stochastic.

We test the hypothesis that expected exchange rate changes equal the expected difference between returns to investments in long-term bonds. The resulting β -coefficients are above 0.6 and insignificantly different from unity for investment horizons above two weeks. The point estimate rises up to a maximum of 1.423 for quarterly investments in long-term bonds, which is much higher than in previous studies. Furthermore, the joint hypothesis of a zero (constant) risk premium and a unity slope coefficient cannot be rejected for any investment horizon. In stark contrast to the well documented finding of a negative relationship between exchange rate changes and short-term interest differentials, the behavior of US dollar - Deutsche Mark exchange rate and returns to short investments in German and American long-term bonds appears to be consistent with standard asset pricing theory.

In previous studies of UIP for long-term interest rates, the positive β -coefficients are explained in terms of the long investment horizons needed to unravel relationships between exchange rates and fundamental variables such as interest rates. This explanation implies that returns to short investments in long-term bonds should behave like short-term interest rates. However, we obtain even larger β -coefficients for short investments in long-term bonds than what Alexius and Meredith and Chinn document for long investment horizons. We also match the data on exchange rate changes to short-term

interest rates rather than bond returns. In this case, the standard finding of a negative and large β -coefficient is confirmed. It hence appears to be the maturity of the instrument *per se* rather than the length of the investment horizon that matters for the results.

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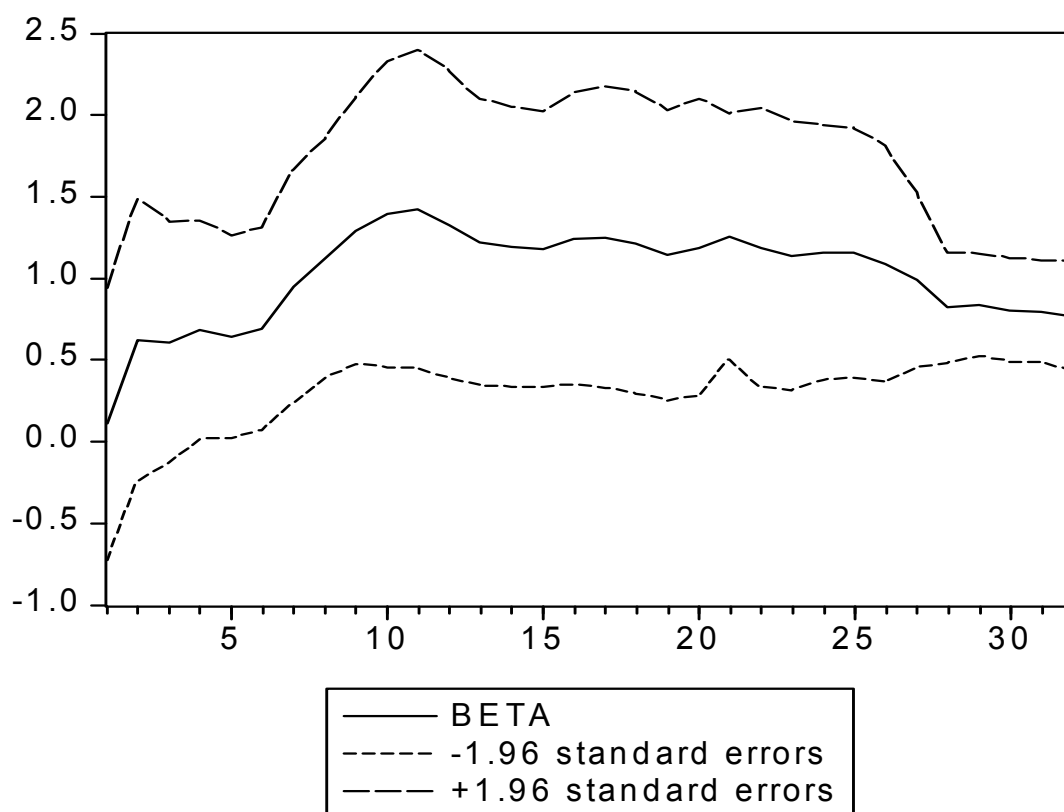
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Figure 1: Point estimates of β for long-term interest rates as the investment horizons is extended from one to 32 weeks, and 95 percent confidence intervals.*



* Baseline set of instruments: Zero coupon interest rate differentials, lagged bond returns and short-term interest rate differentials.

Figure 2: Point estimates of β using alternative sets of instruments: The original bond returns, lagged bond returns and zero coupon interest rates, and zero coupon interest rates only.

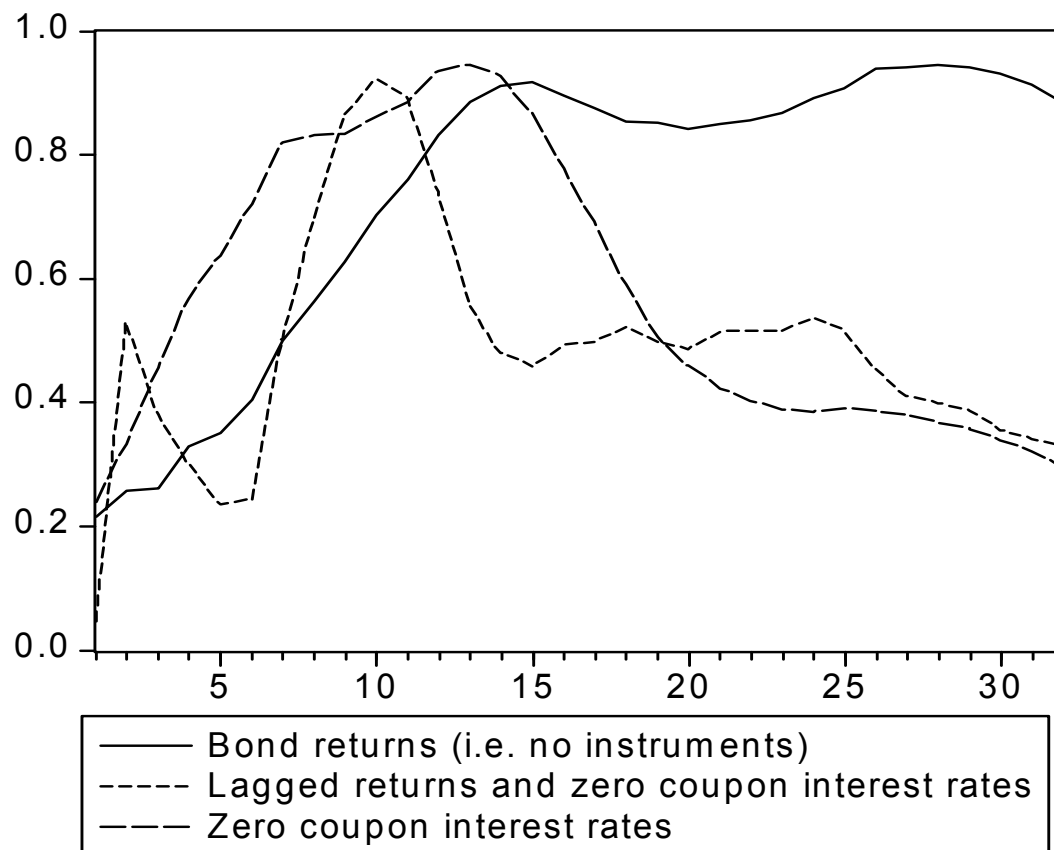


Figure 3: Point estimates of β for short-term interest rates as the investment horizon is extended from one to 32 weeks, and 95 percent confidence intervals.

