

Stavrakeva, Vania; Tang, Jenny

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Exchange Rates and the Yield Curve

Vania Stavrakeva and Jenny Tang

Abstract:

In this paper, we confront the data with the financial markets folk wisdom that an increase in a yield or forward rate of country i relative to j is associated with a contemporaneous appreciation of currency i . We find that while the folk wisdom prior to 2009:Q1 holds fairly well for all maturities and three major currency bases, the “coefficient curve” twisted during the zero-lower-bound period so that the relationship became stronger at the short end but weaker and even of the opposite sign at the long end of the curve. We attribute the structural breaks at the short end of the curve to a change in the relationship between expected excess currency returns and changes in relative yields/forwards. The breaks at the long end of the curve can be explained by changing relationships between yields/forwards and the part of exchange rate fluctuations due to changes in expectations over future short-term rates and long-run relative price levels. Alternatively, the twist of the coefficient curve can be attributed to the changing relationship between the exchange rate and the expectations hypothesis component of yields/forwards at the short end and the term premium component at the long end.

JEL Classifications: G15

Vania Stavrakeva is an assistant professor of economics at the London Business School. Jenny Tang is an economist in the research department of the Federal Reserve Bank of Boston. Their email addresses are vstavrakeva@london.edu and jenny.tang@bos.frb.org, respectively.

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1 Introduction

In financial markets, folk wisdom holds that a country's currency appreciates when its interest rate increases relative to that of other countries. The relationship is contemporaneous and yields are often perceived to be driven by monetary policy—both conventional and unconventional. Statements such as "Simon Derrick, chief currency strategist at BNY Mellon, notes a sharp correlation between the euro/dollar trade and changes in the yield differential between 10-year Treasury and Bund paper. 'The yield differential is the be-all and end-all,' he says,"¹ and "The dollar is almost universally expected to appreciate when US interest rates start rising..."² are becoming more prevalent in the financial press.

In this paper, we systematically dissect the empirical contemporaneous link between yields and exchange rate fluctuations at a quarterly frequency, relying on a parsimonious model based on expectations of short-run yields and inflation implied by a vector autoregression (VAR). We also rely on the assumption that the real exchange rate is stationary, which seems to be supported by the data. We consider 10 liquid currencies whose government debt is risk free and study currency pairs with respect to three base currencies—the U.S. dollar (USD), the British pound (GBP), and the euro (DEM/EUR).

We first analytically derive the theoretical relationship between contemporaneous changes in relative yields/forwards at all maturities and changes in the exchange rate assuming an AR(1) process for short-term interest rates. We show that while the theoretical predictions hold reasonably well for short-maturity yields—and correspondingly, short-horizon forward rates—they are rejected for longer-maturity yields and more-distant forward rates. Rejection of the theoretical relationship reflects a failure of our assumptions, which predominantly assert that changes in yields and forward rates do not correlate with changes in term premia

¹Roger Blitz and Claire Jones, "Merkel's 'Strong Euro' Comments Put Currency Investors on Alert," *Financial Times*, June 15, 2015. Available at <http://www.ft.com/intl/cms/s/0/8d3ec190-1121-11e5-9bf8-00144feabdc0.html#axzz3r2hkYvOy>.

²Anatole Kaletsky, "What a US interest rate rise really means for the dollar," *The Guardian*, November 17, 2015. Available at <http://www.theguardian.com/business/2015/nov/17/what-a-us-interest-rate-rise-really-means-for-the-dollar>.

and exchange rate components in a way that generates omitted variable bias. Our results indicate that while this omitted variable bias is weak at the short end of the curve, it becomes progressively stronger as the maturity increases. This implies that the theoretical relationship between changes in relative yields/forwards and exchange rate appreciation, which is also supported by the folk wisdom in financial markets, is less pronounced at the long end of the curve.

We also find that there is an important structural break in the relationship in 2009:Q1. Prior to the zero lower bound (ZLB) period, an increase of a forward rate of country i relative to country j led to an appreciation of currency i , as the theory would predict, for almost all maturities and currency pairs with the appreciation associated with longer-horizon forwards being much smaller than the appreciation associated with shorter-horizon forwards. During the ZLB period, the appreciation became even stronger at the short end of the forward curve, while at the long end, there was statistically significant depreciation. A similar twisting of this "coefficient curve" was present with yields as well, even though the coefficients estimated over the post-2009:Q1 subsample tended to maintain the same sign.

We take two different approaches in order to shed more light on the source of the structural breaks in the contemporaneous relationship between yields/forwards and exchange rates. First, using short-term rate expectations based on an estimated VAR, we decompose yields and forwards into expectations hypothesis and term premia terms. We find that the twisting of the coefficient curve can be attributed to a structural break in the relationship of exchange rates with the expectations hypothesis component of yields/forwards at short maturities and with the term premia components at longer maturities. In other words, during the ZLB period, an increase in the relative expectations hypothesis term was associated with an even larger exchange rate appreciation, while an increase in the relative term premia, especially at long maturities, led to an exchange rate depreciation. This result would imply that if Quantitative Easing (QE) worked entirely by lowering the term premium of long-term yields, it should have most likely led to currency appreciation rather than depreciation.

Second, we decompose exchange rate changes into five terms with useful economic interpretations: (i) relative lagged interest rates (the uncovered interest rate parity term), (ii) lagged expected excess returns, (iii) changes in expectations over current and future excess returns, (iv) changes in expectations over current and future short-term interest rates (which largely capture the expected path of monetary policy), and (v) changes in expectations over the long-run nominal exchange rate.³ Given our assumption of stationary real exchange rates, component (v) can be also interpreted as changes in expectations over the path of relative inflation rates. Through the prism of this decomposition, we offer a different breakdown of the drivers behind the twisting of the coefficient curve. The following interesting patterns emerge. The structural break at the short end of the curve can be attributed to the fact that prior to the ZLB period, a cross-currency bond position that enjoyed an increase in yields was associated with higher contemporaneous and future expected excess currency returns, which lower the current value of the currency. The opposite is true during the ZLB period. The twisting at the long end of the curve can be attributed to the fact that, during the ZLB period, an increase in relative yields/forwards was associated with higher expected relative long-run price levels and smaller increases in expected future relative short-term rates.

Most of the existing economics research that studies exchange rates and yields has focused on the relationship between lagged short-term or long-term relative yields and exchange rate changes (see, for example, Bekaert, Wei, and Xing (2007) and Engel (2014) for a literature review, along with Engel (2016) and Valchev (2016) for more recent contributions). In the finance literature, our work is closely related to Ang and Chen (2010) and Chen and Tsang (2013). Both of these papers estimate forecasting equations for either exchange rate changes or excess currency returns, using yield curve variables. The key distinction is that these papers focus on the ability of *lagged* yield curve variables to predict future exchange rate

³Other papers that consider a similar decomposition of the exchange rate change include Froot and Ramadorai (2005), Engel and West (2005; 2006), Engel, Mark, and West (2008), Mark (2009), Engel, West, and Zhu (2010), and Engel (2014; 2016).

changes or returns, while we examine whether *contemporaneous changes* in yields comove with exchange rate changes in a manner consistent with economic theory. These literatures have ignored the contemporaneous relationship between yields and exchange rates.

A different literature has studied the contemporaneous relationship between yields and exchange rates, but for the purpose of finding common factors that drive both exchange rates and yields. These papers typically link exchange rates and bond yields by structurally estimating two-country term structure models. Examples include Bansal (1997), Backus, Foresi, and Telmer (2001), Ahn (2004), Brennan and Xia (2006), Wu (2007), Sarno, Schneider, and Wagner (2012), and Yung (2014). Our paper approaches this relationship from the other direction. Rather than estimating a model using structural assumptions, we estimate reduced-form relationships and compare those to theoretical predictions. Furthermore, our theoretical relationships are derived using fewer restrictions—in particular, we do not assume complete markets or a particular functional form for the stochastic discount factor, as much of this literature does. Thus, our empirical findings may be less subject to misspecification. Furthermore, we decompose exchange rate changes in a way that allows us to understand these relationships in the context of macroeconomic concepts such as monetary policy surprises and expectations of future interest rate changes.

This paper is also related to the literature examining the effect of monetary policy shocks—both conventional and unconventional—on exchange rates and yields, using both VAR and high-frequency identification techniques (see Eichenbaum and Evans (1995), Rogers (1999), Faust and Rogers (2003), and Scholl and Uhlig (2008), Rogers, Scotti, and Wright (2015)). One main difference from this literature is that we study the unconditional relationship between exchange rates and yields/forwards rather than the relationship conditional on a particular shock.

2 Empirical Framework

To build our empirical framework, we start with definitions of expected excess returns on bonds of different maturities in different currencies. Both in this section and in our data, one time period is a quarter. $S_t^{i,j}$ denotes the exchange rate of currency i per one unit of currency j . This exchange rate is defined such that an increase of $S_t^{i,j}$ corresponds to a depreciation of currency i against currency j . Below, we introduce some notation and definitions. The one-period return in units of currency i on an n -period bond denominated in currency i is

$$R_{t+1}^{n,i} = \frac{P_{t+1}^{n-1,i}}{P_t^{n,i}},$$

where $P_t^{0,i} = 1$. The log yield-to-maturity of this same bond is defined as

$$y_t^{n,i} = -\frac{1}{n} \ln (P_t^{n,i})$$

and the forward rate from periods $t + n$ to $t + n + m$ at time t is given by

$$f_t^{n,n+m,i} \equiv (n + m) y_t^{n+m,i} - n y_t^{n,i}. \quad (1)$$

The yield on a one-period (“short-term”) bond is given by $i_t^i \equiv y_t^{1,i} = -\ln (P_t^{1,i})$ and its main determinant is the monetary policy rate targeted by the central bank of country i .⁴

2.1 One-Period, Risk-Free Bonds

First, we define the expected excess return from taking a long position in one-period, risk-free bonds of currency j and a simultaneous short position in one-period, risk-free bonds of currency i . Denoting logs of variables with lowercase letters, the expected excess return from this trade is

$$\sigma_t^{i,j} \equiv E_t \Delta s_{t+1}^{i,j} - (i_t^i - i_t^j) \quad (2)$$

⁴The correlation between the average quarterly monetary policy target rate and the end-of-period three-month yield is very strong prior to the zero lower bound period.

Section 6.1 in the Appendix shows that in a simple model based on a standard no-arbitrage condition, $\sigma_t^{i,j}$ is equal to the currency risk premium plus Jensen's inequality terms. However, in reality, there are numerous frictions in addition to currency risk premia that might enter the expected excess return. These include the inability of traders to borrow at the risk-free government bond rate, counterparty risk, as well as binding net worth or value-at-risk constraints. If $\sigma_t^{i,j}$ is assumed to be zero, then equation (2) gives the strong form of the uncovered interest rate parity (UIRP) condition, where the currency of country i is expected to depreciate when the interest rate of country i exceeds that of country j .

Using equation (2), the actual change in the exchange rate can be written as

$$\Delta s_{t+1}^{i,j} = i_t^i - i_t^j + \sigma_t^{i,j} + \Delta s_{t+1}^{i,j} - E_t \Delta s_{t+1}^{i,j}. \quad (3)$$

A vast literature has focused on studying the link between the exchange rate change, the UIRP term, $i_t^i - i_t^j$, and the currency risk premium in period t , $\sigma_t^{i,j}$. In this paper, we argue that understanding the expectational error term $\Delta s_{t+1}^{i,j} - E_t \Delta s_{t+1}^{i,j}$ is equally important for understanding the dynamics of the exchange rate change. The expectational error is assumed to be mean zero and uncorrelated with variables in the information set used to form exchange rate expectations in period t . This term then contains only period $t + 1$ information, which cannot be used for forecasting. In this paper, we show that there is a key link between contemporaneous (period $t + 1$) movements of yields, which are largely driven by monetary policy actions—both conventional and unconventional—and exchange rate changes, a link that appears through this expectational error term.

To further delve into this expectational error, we iterate equation (2) forward to obtain

$$s_t^{i,j} = -E_t \sum_{k=0}^{\infty} [i_{t+k}^i - i_{t+k}^j + \sigma_{t+k}^{i,j}] + E_t \lim_{k \rightarrow \infty} s_{t+k}^{i,j}. \quad (4)$$

First-differencing equation (4) and combining the resulting expression with equation (2)

implies that the expectational error can be expressed as

$$\begin{aligned} \Delta s_{t+1}^{i,j} - E_t \Delta s_{t+1}^{i,j} &= - \underbrace{\sum_{k=0}^{\infty} (E_{t+1} \tilde{i}_{t+k+1}^{i,j} - E_t \tilde{i}_{t+k+1}^{i,j})}_{\varphi_{t+1}^{EH}} \\ &\quad - \underbrace{\sum_{k=0}^{\infty} (E_{t+1} \sigma_{t+k+1}^{i,j} - E_t \sigma_{t+k+1}^{i,j})}_{\sigma_{t+1}^F} + \underbrace{E_{t+1} \lim_{k \rightarrow \infty} s_{t+k}^{i,j} - E_t \lim_{k \rightarrow \infty} s_{t+k}^{i,j}}_{s_{t+1,\infty}^{\Delta E}}, \end{aligned} \quad (5)$$

$$\text{where } \tilde{i}_t^{i,j} \equiv i_t^i - i_t^j.$$

Equation (5) allows us to express the realized exchange rate changes as forward-looking variables which, in addition to the period t interest rate differential and expected excess return, also reflect changes in expectations in contemporaneous ($t + 1$) and future relative short-term rates, φ_{t+1}^{EH} , expected excess returns, σ_{t+1}^F , and long-run nominal exchange rate levels, $s_{t+1,\infty}^{\Delta E}$. We show in Section 3 that if the real exchange rate is stationary, $s_{t+1,\infty}^{\Delta E}$ will reflect changes in expectations over long-run relative price levels or the infinite sum of future relative inflation. To simplify notation, we suppress the “ i, j ” superscript from this point forward. Combining equations (2) and (5) implies that

$$\Delta s_{t+1} = \tilde{i}_t - \varphi_{t+1}^{EH} + \sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}. \quad (6)$$

One of the most important components that enters the expectational error in equation (6) is the expectations hypothesis term, φ_{t+1}^{EH} . It can be decomposed into a contemporaneous surprise $i_{t+1}^i - E_t i_{t+1}^i$ and changes in expectations of future short-term policy rates (period $t + 2$ and onwards), $\sum_{k=1}^{\infty} (E_{t+1} \tilde{i}_{t+k+1}^{i,j} - E_t \tilde{i}_{t+k+1}^{i,j})$. To the extent that the short-term interest rate, when not constrained by the ZLB, captures monetary policy, φ_{t+1}^{EH} captures changes in expectations over $t + 1$ and future policy rates.

These expectational changes can arise from a surprise change in the policy rate in $t + 1$ or from information acquired within the quarter that is useful for forecasting the future state of the economy and future monetary policy responses to those developments. Traders may

also believe that there is inherent gradualism in the rate-setting process, so that current monetary policy surprises foreshadow future policy rate changes. For details on monetary policy smoothing, where policymakers are assumed to smooth either the level or change of the policy rate, see Coibion and Gorodnichenko (2012).

Finally, even in the absence of contemporaneous macroeconomic news or monetary policy surprises, expectations regarding future policy may change due to a central bank engaging in unconventional monetary policy actions such as forward guidance or QE. With forward guidance, policymakers explicitly communicate their plans for future policy and, if credible, these statements will inform the market's forecast of future monetary policy. QE can affect the market's expectations of future policy rates by acting as a signal that the central bank will be committed to expansionary policy for some time. For examples of empirical and theoretical work on the signaling channel of QE, see Bauer and Rudebusch (2014) and Bhattarai, Eggertsson, and Gafarov (2015).

Since long-term yields and forwards capture both contemporaneous and future short-term rates, they are tightly linked to the φ_{t+1}^{EH} term, which is what we show in the next two subsections.

2.2 Long-Term Yields and Forward Rates

Given that long-term yields directly capture current short-term rates and expectations of future short-term rates, in addition to term premia, it should be interesting to explore the theoretical relationship between yields of different maturities or forwards at different horizons and exchange rate changes. In our empirical work, we decompose the yield on an n -period bond denominated in currency i into terms that we refer to as the expectations hypothesis

component and term premium

$$y_t^{n,i} = EH_t^{n,i} + TP_t^{n,i} \quad (7)$$

$$\text{where } EH_t^{n,i} \equiv \frac{1}{n} \sum_{k=0}^{n-1} E_t \tilde{l}_{t+k}^i. \quad (8)$$

More generally, this term premium is just the expected excess return from taking a long position in this n -period bond and short positions in one-period bonds for each of the next n periods. Substituting equations (7) and (8) into equations (3) and (5) reveals that the quarterly exchange rate change can be expressed as a function of the change in relative n -period bond yields:

$$\Delta s_{t+1} = -n \Delta \tilde{y}_{t+1}^n - \sum_{k=n}^{\infty} (E_{t+1} \tilde{l}_{t+k+1} - E_t \tilde{l}_{t+k}) + n \Delta \widetilde{TP}_{t+1}^n + \sigma_t - \sigma_{t+1}^F + s_{t+1, \infty}^{\Delta E}, \quad (9)$$

where $\widetilde{TP}_t^n \equiv TP_t^{n,i} - TP_t^{n,j}$, $\widetilde{EH}_t^n \equiv EH_t^{n,i} - EH_t^{n,j}$ and $\tilde{y}_t^n \equiv y_t^{n,i} - y_t^{n,j}$. Equation (9) is suggestive as to why we might expect a strong relationship between exchange rate changes and changes in yields at different horizons.

From the definition of forward rates given in equation (1), we can analogously define the expectations hypothesis component and term premia associated with forward rates as follows:

$$\begin{aligned} f_t^{n,n+m,i} &= TP_t^{n,n+m,i} + EH_t^{n,n+m,i} \\ \text{where } EH_t^{n,n+m,i} &\equiv (n+m) EH_t^{n+m,i} - n EH_t^{n,i} = \sum_{k=n}^{n+m-1} E_t \tilde{l}_{t+k}^i \\ \text{and } TP_t^{n,n+m,i} &\equiv f_t^{n,n+m,i} - EH_t^{n,n+m,i} = (n+m) TP_t^{n+m,i} - n TP_t^{n,i}. \end{aligned}$$

Thus, an alternate expression for exchange rate changes involving m -period forward rates

and forward term premia is

$$\begin{aligned} \Delta s_{t+1} = & - \sum_{k=0,m,2m,\dots,(N-1)m} \Delta \tilde{f}_{t+1}^{k,k+m} - \underbrace{\left[-E_t \tilde{v}_{t+Nm} + \sum_{k=Nm}^{\infty} (E_{t+1} \tilde{v}_{t+k+1} - E_t \tilde{v}_{t+k+1}) \right]}_{\Delta \widetilde{EH}_{t+1}^{Nm,\infty}} \\ & + \sum_{k=0,m,2m,\dots,(N-1)m} \Delta \widetilde{TP}_{t+1}^{k,k+m} + \sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}, \end{aligned}$$

where $\tilde{f}_{t+1}^{k,k+m} \equiv f_{t+1}^{k,k+m,i} - f_{t+1}^{k,k+m,j}$ and $\Delta \widetilde{EH}_{t+1}^{Nm,\infty} = \lim_{T \rightarrow \infty} (\Delta EH_{t+1}^{Nm,T,i} - \Delta EH_{t+1}^{Nm,T,j})$.

It will be interesting to explore the link between exchange rate changes and forward rates separately from yields, since forward rates capture information about a particular horizon in the future. Likewise, we can also disentangle movements of the yield curve due to short-term rate expectations or term premia at different horizons. Focusing in on particular horizons will be useful, since many unconventional policy actions such as forward guidance or quantitative easing have historically had their largest effects on particular regions of the yield curve. Swanson (2015) finds that the effect of forward guidance typically peaks at a maturity of about one year, while Greenwood, Hanson, and Vayanos (2015) finds the effects of QE announcements to be the largest for 1-year forward rates five and seven years ahead.

2.3 Persistence of the Policy Rate

As discussed above, reasons such as monetary policy smoothing can give rise to persistence in short-term interest rates. To illustrate the implications of this persistence for the link between contemporaneous exchange rate dynamics and changes in short- and long-term interest rates or forward rates, this section considers a parsimonious example where the short-term rate follows an AR(1) process,

$$i_{t+1}^j = \bar{v}_j + \rho_j i_t^j + \varepsilon_{t+1}^j, \quad (10)$$

where we assume that ε_{t+1}^i has a mean of zero and is a serially uncorrelated random variable and that $|\rho| < 1$. In Appendix Table 8, we present the results from the simple AR(1)

regression for a sample starting in 1990:Q1 and the following countries: Australia, Canada, Germany/euro area, Japan, New Zealand, Norway, Sweden, Switzerland, the United Kingdom, and the United States.⁵ These countries and data sample were selected based on the availability of data on zero-coupon bond yields in addition to having low default probabilities and highly liquid sovereign debt. See Section 6.3 in the Appendix for a description of the data. As the AR(1) regressions of the policy rates in Table 8 show, short-term interest rates are highly persistent, with estimated ρ 's above 0.87 for all 10 countries. This simple AR(1) process is also a good descriptor of the level of the short-term interest rate, as evidenced by adjusted R^2 's of above 0.9 for all countries.

Under equation (10), φ_{t+1}^{EH} , defined in (5), can be expressed as

$$\varphi_{t+1}^{EH} = \frac{1}{1 - \rho_i} \varepsilon_{t+1}^i - \frac{1}{1 - \rho_j} \varepsilon_{t+1}^j.$$

For the remainder of this section, we consider the special case $\rho = \rho_i = \rho_j$, which does not appear to be an unrealistic assumption for our sample of countries based on the results in Table 8. In this case, we can combine the above expression with equations (6) and (10) to express the exchange rate change as

$$\Delta s_{t+1}^{i,j} = -\frac{1}{1 - \rho} \Delta \tilde{v}_{t+1}^{i,j} + \sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}. \quad (11)$$

If the correlation between the interest rate changes and the remaining terms, $\sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}$, is small and if the AR(1) model is an approximately accurate representation of how traders form expectations, our ρ estimates imply that the β_1 coefficient in the regression,

$$\Delta s_{t+1} = \alpha_1 + \beta_1 \Delta \tilde{v}_{t+1} + error_{t+1}, \quad (12)$$

should be between -23.3 and -7.9 . Table 1 reports the results of the regression specified in (12).

⁵Throughout the paper, short-term interest rates and macroeconomic data for Germany are replaced with euro-area data starting in 1999:Q1. For long-term yields, German zero-coupon yields are used throughout the entire sample period, since our analysis applies to risk-free sovereign debt.

The β_1 coefficients are negative and significantly different from zero (as suggested by equation (11)) in 11 out of 24 cases. In 19 out of 24 cases, we cannot reject that β_1 is within the theoretical range suggested by a simple AR(1). This result is quite remarkable, considering that this regression may have omitted variable bias due to the missing $\sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}$ term. Hence, Table 1 provides suggestive evidence that for many currency pairs, the historical relationship between short-term interest rate changes and exchange rate changes is close to the one implied by a simple AR(1) process for short-term interest rates with no correlation between $\sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}$ and $\Delta \tilde{y}_{t+1}^{i,j}$.

If we consider long-term yields, equation (10) implies that equation (9) can be re-written as

$$\Delta s_{t+1}^{i,j} = -\frac{1}{1-\rho^n} \left(n\Delta \tilde{y}_{t+1}^n - n\Delta \widetilde{TP}_{t+1}^n \right) + \sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}. \quad (13)$$

If we now assume that the correlation between $\frac{1}{1-\rho^n} n\Delta \widetilde{TP}_{t+1}^n + \sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}$ and $n\Delta \tilde{y}_{t+1}^n$ is zero, then we would expect a coefficient of $\beta_{y,n} = -\frac{1}{1-\rho^n}$ in the following regression, which generalizes the regression in (12) to longer maturities:

$$\Delta s_{t+1} = \alpha_{y,n} + \beta_{y,n} n\Delta \tilde{y}_{t+1}^n + error_{t+1}. \quad (14)$$

Note that the theory-implied $\beta_{y,n}$ approaches -1 as the maturity of the bond grows. The results from this regression for $n = 20$ are presented in Table 2. In the absence of any omitted variable bias and given the estimates of the AR(1) short-term rate process, a reasonable range for $\beta_{y,20}$ is between -1.7 and -1.1 . The estimates of $\beta_{y,20}$ are indeed negative and significantly different from zero in 13 out of 24 cases. In 12 out of 24 cases, we cannot reject the null hypothesis that the estimate is inside the range predicted by the simple AR(1) model under the assumptions above.

Table 3 presents the results of regression (14) for 10-year yields ($n = 40$). Compared to the above results with 5-year yields, the results here are quite puzzling. In the absence of any omitted variable bias and given the estimates of the AR(1) short-term rate process, a

reasonable range for $\beta_{y,40}$ is between -1.2 and -1.005 . The estimates of $\beta_{y,40}$ are now only negative and significantly different from zero in 7 out of 24 cases while two currency pairs actually yield positive and significant coefficients. Furthermore, for all 24 currency pairs, we can quite strongly reject the null that $\beta_{y,40}$ is inside the range implied by the simple AR(1) model under the assumptions above.⁷

The results in Tables 1, 2, and 3 suggest that the relationship between exchange rates and yields is approximated well by the simple AR(1) model with little or no omitted variable bias at short horizons. In other words, since yields cumulate short-term rate expectations and term premia at multiple horizons, the results in Table 1 might be in large part credited for the results in Table 2. To take a closer look at this possibility, we now examine changes in relative forward rates for which equations (1) and (13) imply:

$$\Delta s_{t+1}^{i,j} = -\frac{1}{\rho^n(1-\rho^m)} \left(\Delta \tilde{f}_{t+1}^{n,n+m} - \Delta \widetilde{TP}_{t+1}^{n,n+m} \right) + \sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}. \quad (15)$$

If we assume that $\frac{1}{\rho^n(1-\rho^m)} \Delta \widetilde{TP}_{t+1}^{n,n+m} + \sigma_t - \sigma_{t+1}^F + s_{t+1,\infty}^{\Delta E}$ and $\Delta \tilde{f}_{t+1}^{n,n+m}$ are uncorrelated, then we would expect a coefficient of $\beta_{f,n,m} = -\frac{1}{\rho^n(1-\rho^m)}$ in the following regression

$$\Delta s_{t+1} = \alpha_{f,n} + \beta_{f,n,m} \Delta \tilde{f}_{t+1}^{n,n+m} + error_{t+1}. \quad (16)$$

The results from this regression for 7-year forward rates three years ahead ($n = 12$ and $m = 28$) are presented in Table 4. The estimates of $\beta_{f,12,28}$ are negative and significant in only one out of 24 cases and are actually positive and significant in five out of 24 cases. All of the estimates are quite far from the theory-implied coefficient range of -5.2 to -2.4 , and the null hypothesis that the true coefficients fall within this range are resoundingly rejected for all currency pairs.

To summarize, reduced-form bilateral regressions tell us that the theory-implied empirical relationships between exchange rate dynamics and changes in yields with short maturities are not completely overpowered by omitted variable bias. However, this is not the case when

⁷We can reject $\beta_{y,40} \leq -1.005$ at a 1 percent level for 22 out of the 24 currency pairs.

one considers forward rates and yields at longer maturities. To examine this issue further we consider panel regressions.

2.3.1 Panel Regressions

For the remainder of the paper, we pool data by currency base and estimate panel regressions, allowing for currency-pair fixed effects, and present the same regressions over yearly horizons up to 10 years, $n \in \{0, 4, \dots, 40\}$, where, with respect to forwards, we use 1-year rates, that is, $m = 4$. We allow a break in the coefficients starting in 2009:Q1, which corresponds to the quarter in which the United States first hit the ZLB.⁹

Figures 1 and 2 present the panel regression results of regressions (14) and (16) for the following base currencies: USD, GBP, DEM/EUR. The bars denote 80 percent confidence intervals. Starting with regression (14), we observe the following patterns. For the USD and DEM/EUR bases and to a lesser degree for the GBP base, the short end of the curve exhibits less negative coefficients prior to the ZLB period than during the ZLB period. The caveat is that the point estimates are less precisely estimated during this latter period. For maturities longer than 20 quarters, this relationship flips for all bases, implying that the point estimates at the long end of the yield curve are less negative during the ZLB period than earlier. In many cases, they even become significantly positive.

To delve further into this interesting pattern, we can examine the panel regressions results from (16). By focusing on a particular horizon using forward rates, this pattern becomes even more striking and is apparent for all currency bases. The slope coefficients associated with changes in 1-year forward rates that are more than four or five years ahead are positive and statistically significant in the ZLB period, rather than negative. This echoes the finding in Table 4 implying that the omitted variable bias is strong enough at longer horizons to

⁹Our yields data for Japan end in 2009:Q2, so there is a difference in the relative weight that JPY pairs receive in the two subsamples. The results are qualitatively the same if we drop all JPY pairs from our analyses.

produce estimated coefficients that are opposite in sign to those implied by equation (15).

Although the reduced-form regressions reveal interesting patterns, many questions remain, such as: what is the source of the omitted variable bias? In other words, is it due to a departure of interest rate expectations from those implied by AR(1) dynamics or to non-zero covariances between changes in yields/forwards and expected excess currency return terms, σ_t and σ_{t+1}^F , or to changes in expectations over the long-run nominal exchange rate, $s_{t+1,\infty}^{\Delta E}$? Also, what is the relationship between the expectations hypothesis and the term premium components of yields/forwards and the exchange rate change? Last but not least, which one of the exchange rate components in decomposition (6) is the main driver of the structural breaks observed in the panel regressions?

In order to be able to answer these questions, the next section describes how we can calculate the various terms in equation (6) as well as the expectations hypothesis and term premia components of yields/forwards using a VAR approach.

3 VAR Approach

To compute the terms in equation (6), we assume that exchange rates and short-term interest rates can be described by the following reduced-form VAR(p) process:

$$F_{t+1} = \bar{\mathbf{F}} + \gamma(L) F_t + \varepsilon_{F,t+1} \quad (17)$$

$$\text{where } \gamma(L) \equiv \gamma_1 + \gamma_2 L + \dots + \gamma_p L^{p-1}$$

$$\text{and } F_{t+1} \equiv [q_{t+1}^{i,j}, x_{t+1}^i, z_{t+1}^i, x_{t+1}^j, z_{t+1}^j, x_{t+1}^{US}, z_{t+1}^{US}]'. \quad (18)$$

Here, $q_{t+1}^{i,j}$ is the level of the real exchange rate between currencies i and j . By including the real exchange rate in levels, we are estimating a specification where a stable estimate of the VAR implies that long-run purchasing power parity holds and VAR-based expectations of the long-run real exchange rate are constant. The vector x_{t+1} is a set of financial variables

that includes the three-month bill rate as well as the empirical term structure slope and curvature factors defined as follows:

$$\begin{aligned} sl_{t+1}^i &= y_t^{40,i} - i_t^i \\ c_{t+1}^i &= 2y_t^{8,i} - (y_t^{40,i} + i_t^i). \end{aligned}$$

The vector z_{t+1} represents other variables that may be useful for forecasting either short-term interest rates or changes in the exchange rate. Importantly, we always include a quarterly inflation rate (measured using the GDP deflator) in z_{t+1} . This allows us to compute VAR-based expectations of nominal exchange rate changes from our estimates of the real exchange rate and inflation equations. The other variables in z_{t+1} include the GDP gap and the current-account-to-GDP ratio.

This reduced-form VAR(p) can be transformed into the VAR(1) companion form, as follows:

$$\underbrace{\begin{bmatrix} F_{t+1} \\ \vdots \\ F_{t-p+2} \end{bmatrix}}_{\mathbf{x}_{t+1}} = \underbrace{\begin{bmatrix} \bar{\mathbf{F}} \\ 0 \\ 0 \end{bmatrix}}_{\bar{\mathbf{x}}} + \underbrace{\begin{bmatrix} \gamma_1 & \gamma_2 & \cdots & \gamma_p \\ & \mathbf{I} & & \mathbf{0} \end{bmatrix}}_{\mathbf{\Gamma}} \underbrace{\begin{bmatrix} F_t \\ \vdots \\ F_{t-p+1} \end{bmatrix}}_{\mathbf{x}_t} + \underbrace{\begin{bmatrix} \varepsilon_{F,t+1} \\ 0 \\ \vdots \end{bmatrix}}_{\boldsymbol{\varepsilon}_{t+1}}.$$

To ameliorate the problem of overparameterization in unrestricted VARs, we follow Cushman and Zha (1997) in restricting both the contemporaneous and the lagged relationships between the variables in the VAR. More specifically, we consider a specification where each country's financial variables follow a smaller three-variable VAR.¹⁰ This specification is widely used in the asset pricing literature. One advantage of this specification versus one that models the short-term interest rate as a function of macroeconomic variables (such as a Taylor rule) is that it uses information from long-term yields in a parsimonious way. This allows the estimates to better capture the effects of forward guidance, among other things, and is

¹⁰One caveat is that we do not impose a zero lower bound in the VAR. However, we observe very few, small negative short-term rate forecasts. Also short-term rates entered negative territory during the ZLB.

therefore more appropriate for a sample that includes zero lower bound episodes. We present evidence below that our VAR-implied, short-term interest rate expectations accord well with those implied by futures market data.¹¹

Our next set of restrictions concerns the macroeconomic variables. We assume that changing economic conditions in the United States affect expectations over macro variables in other countries through spillovers from the United States into the macroeconomy of these other countries. Commodity-producing countries seem to be the most sensitive to the lagged U.S. output gap. This result is not surprising given the high sensitivity of growth in these economies to world growth. See Miranda-Agrippino and Rey (2015) for VAR-based evidence of such spillovers. At the same time, we restrict U.S. macroeconomic variables to depend only on lags of themselves and U.S. financial variables. Lastly, we allow the real exchange rate to enter as a lag only in its own equation. To summarize, if we partition each matrix $\{\gamma_1, \dots, \gamma_p\}$ into seven blocks corresponding to the partitioning of F_{t+1} given in (18), then the above restrictions imply the following zero restrictions on $\{\gamma_1, \dots, \gamma_p\}$:

$$\gamma_l = \begin{bmatrix} \bullet & \bullet & \bullet & \bullet & \bullet & \bullet & \bullet \\ 0 & \bullet & 0 & 0 & 0 & 0 & 0 \\ 0 & \bullet & \bullet & 0 & 0 & \bullet & \bullet \\ 0 & 0 & 0 & \bullet & 0 & 0 & 0 \\ 0 & 0 & 0 & \bullet & \bullet & \bullet & \bullet \\ 0 & 0 & 0 & 0 & 0 & \bullet & 0 \\ 0 & 0 & 0 & 0 & 0 & \bullet & \bullet \end{bmatrix} \quad \text{for } l = 1, \dots, p.$$

We estimate this VAR for each i, j pair by OLS with a lag length of two quarters. Since we use the VAR only for the purpose of calculating expectations of the short-term interest rates, inflation, and exchange rates, we do not attempt to identify the structural shocks. Notice that, for most of the results that follow, the exact real exchange rate specification is

¹¹The use of futures market data for this comparison becomes more problematic for longer-horizon forecasts due to term premia in futures prices.

irrelevant, besides the fact that purchasing power parity is assumed to hold in the long run, which is confirmed in the data. We clarify this later on.

As a check for external validity, we can compare our model-implied interest rate expectations with market-based measures of short-term interest rate surprises computed using futures prices by adapting the method used in Bernanke and Kuttner (2005) to a quarterly frequency. We find that the one-period-ahead expected U.S. short-term interest rate change, $\hat{E}_t [i_{t+1}^{US}] - i_t^{US}$, has a correlation of 75 percent with the expected component of the U.S. three-month eurodollar interest rate change over the full sample, and 83 percent over the pre-ZLB period. The model-implied quarterly U.S. short-term interest rate surprise, $i_{t+1}^{US} - \hat{E}_t [i_{t+1}^{US}]$, has a correlation of 77 percent with the market-based federal funds rate surprise measure over the full sample and 79 percent over the pre-ZLB period. Tables 5 and 6 show these correlations over the full sample for a number of additional countries. The high correlations for a large majority of the countries we consider are evidence that our VAR-based short-term interest rate expectations accord reasonably well with expectations held by financial market participants.¹²

With the estimated VARs, we can now decompose exchange rates into the five terms listed in equation (6). First, using the notation of the VAR, the exchange rate change and relative policy rates can be expressed as

$$\begin{aligned} \Delta s_{t+1}^{i,j} &\equiv \Delta q_{t+1}^{i,j} + \pi_{t+1}^i - \pi_{t+1}^j = (e_q + e_{\pi,i} - e_{\pi,j}) \mathbf{X}_{t+1} - e_q \mathbf{X}_t \\ \hat{i}_{t+1}^{i,j} &= (e_i^i - e_i^j) \mathbf{X}_{t+1}, \end{aligned}$$

where e_q is a row vector that selects $q_{t+1}^{i,j}$ from \mathbf{X}_{t+1} . That is, it has the same number of elements as \mathbf{X}_{t+1} with an entry of 1 corresponding to the position of $q_{t+1}^{i,j}$ in \mathbf{X}_{t+1} and zeros

¹²Note that the futures contracts we use are typically written on interbank interest rates, while our VAR produces expectations of three-month T-bill rates. By basing our comparisons on expected interest rate changes and surprises, we are able to abstract from differences in the rates that do not vary at a quarterly frequency. Nonetheless, the differences in the financial instruments might make it harder to detect a high correlation between our VAR-implied expectations and the ones implied by futures prices, even if our VAR accords well with the market's expectations formation process.

elsewhere. Likewise, e_i^i and e_i^j are selection vectors corresponding to the short-term interest rates of countries i and j , respectively.

VAR-based expectations of nominal exchange rates and interest rates can be obtained from those for future \mathbf{X}_{t+k+1} , as follows:¹³

$$\begin{aligned}\hat{E}_t \mathbf{X}_{t+k+1} &= (\mathbf{I} - \mathbf{\Gamma})^{-1} (\mathbf{I} - \mathbf{\Gamma}^{k+1}) \bar{\mathbf{X}} + \mathbf{\Gamma}^{k+1} \mathbf{X}_t \\ \hat{E}_t \Delta s_{t+k+1}^{i,j} &= (e_q + e_{\pi,i} - e_{\pi,j}) [(\mathbf{I} - \mathbf{\Gamma})^{-1} (\mathbf{I} - \mathbf{\Gamma}^{k+1}) \bar{\mathbf{X}} + \mathbf{\Gamma}^{k+1} \mathbf{X}_t] \\ &\quad - e_q [(\mathbf{I} - \mathbf{\Gamma})^{-1} (\mathbf{I} - \mathbf{\Gamma}^k) \bar{\mathbf{X}} + \mathbf{\Gamma}^k \mathbf{X}_t] \\ \hat{E}_t \tilde{\imath}_{t+k+1} &= (e_i^i - e_i^j) [(\mathbf{I} - \mathbf{\Gamma})^{-1} (\mathbf{I} - \mathbf{\Gamma}^{k+1}) \bar{\mathbf{X}} + \mathbf{\Gamma}^{k+1} \mathbf{X}_t].\end{aligned}$$

Note that the change in expectations over future \mathbf{X}_{t+k+1} can be written simply as a linear combination of the time $t + 1$ reduced-form residuals:

$$\hat{E}_{t+1} \mathbf{X}_{t+k+1} - \hat{E}_t \mathbf{X}_{t+k+1} = \mathbf{\Gamma}^k \mathbf{\Xi}_{t+1}.$$

As long as all eigenvalues of $\mathbf{\Gamma}$ are strictly inside the unit circle in the complex plane, the exchange rate change components can be constructed using equations (2) and (5) as follows:

$$\begin{aligned}\tilde{\imath}_t^{i,j} &= (e_i^i - e_i^j) \mathbf{X}_t \\ \sigma_t &= (e_q + e_{\pi}^i - e_{\pi}^j) (\bar{\mathbf{X}} + \mathbf{\Gamma} \mathbf{X}_t) - (e_q + e_i^i - e_i^j) \mathbf{X}_t \\ \varphi_{t+1}^{EH} &= (e_i^i - e_i^j) \mathbf{\Xi}_{t+1} + (e_i^i - e_i^j) \mathbf{\Gamma} (\mathbf{I} - \mathbf{\Gamma})^{-1} \mathbf{\Xi}_{t+1} \\ \sigma_{t+1}^F &= [(e_q + e_{\pi}^i - e_{\pi}^j) \mathbf{\Gamma} - (e_q + e_i^i - e_i^j)] (\mathbf{I} - \mathbf{\Gamma})^{-1} \mathbf{\Xi}_{t+1} \\ s_{t+1,\infty}^{\Delta E} &= (e_{\pi}^i - e_{\pi}^j) (\mathbf{I} - \mathbf{\Gamma})^{-1} \mathbf{\Xi}_{t+1}.\end{aligned}\tag{19}$$

In the decomposition of yields and forward rates, $EH_t^{n,i}$ can be calculated as

$$EH_t^{n,i} = e_i^i (\mathbf{I} - \mathbf{\Gamma})^{-1} \left\{ \bar{\mathbf{X}} + \frac{1}{n} (\mathbf{I} - \mathbf{\Gamma}^n) (\mathbf{X}_t - (\mathbf{I} - \mathbf{\Gamma}_q)^{-1} \bar{\mathbf{X}}) \right\}$$

¹³The \hat{E}_t operator denotes expectations based on the linear projections performed in the VAR estimation. Although not explicitly delineated, the operator conditions only on the set of regressors included in the estimation of each equation. Due to the restrictions set out above, this means that the relevant information set differs across variables.

and $TP_t^{n,i}$ is simply the residual from equation (7).

One key thing to note above is that, since our specification implies constant expectations over long-run levels of the real exchange rate when the VAR is stationary, the change in expectations regarding long-run nominal exchange rates, $s_{t+1,\infty}^{\Delta E}$, simply reflects changes in expectations over long-run relative price levels. Hence, this component can be computed from VAR-based expectations of the future path of inflation rates. Note that, in the subsequent sections, the results for the NOK-JPY and NZD-JPY pairs are not presented because the VAR estimates in this case violate the requirement that all eigenvalues of $\mathbf{\Gamma}$ be strictly inside the unit circle in the complex plane.

4 Empirical Results – Decomposition

In order to further disentangle the results from Section 2.3.1, we use the VAR-based short-term rate expectations combined with yields data to calculate \widetilde{EH}_{t+1}^n , $\Delta\widetilde{TP}_{t+1}^n$, $\Delta\widetilde{EH}_{t+1}^{n,n+4}$, and $\Delta\widetilde{TP}_{t+1}^{n,n+4}$. We regress jointly the exchange rate change on the term premia and expectations hypothesis terms for each maturity/horizon. The exact specifications are given by

$$\Delta s_{t+1} = \hat{\alpha}_{y,n} + \beta_{EHy,n} n \Delta\widetilde{EH}_{t+1}^n + \beta_{TPy,n} n \Delta\widetilde{TP}_{t+1}^n + error_{t+1} \quad (20)$$

$$\Delta s_{t+1} = \hat{\alpha}_{y,n} + \beta_{EHf,n} \Delta\widetilde{EH}_{t+1}^{n,n+4} + \beta_{TPf,n} \Delta\widetilde{TP}_{t+1}^{n,n+4} + error_{t+1}. \quad (21)$$

The results for both of the regressions above are reported in Figures 3 and 4.

The results from regression (20) paint quite a clear picture. The structural break at the short end of the curve that we observed in Figure 1 appears to be attributable to a structural break in the $\beta_{EHy,n}$ coefficient, while the structural break at the long end can be attributed to a structural break in the $\beta_{TPy,n}$ coefficient. Controlling for the term premia, a 1 percentage point large change in the relative expectations hypothesis component of yields/forwards prior

to the ZLB period led to a smaller appreciation of currency i vis-à-vis currency j compared with the ZLB period. This difference is largest at shorter maturities. In contrast, prior to the ZLB, a 1 percentage point increase in the relative term premia, controlling for movements in the expectations hypothesis term, led to an appreciation of currency i relative to j , while during the ZLB period, for both the short and the long end, it led to a depreciation in almost all cases or to a much smaller appreciation. This pattern is consistent across all currency pairs.

Next, consider the results from regression (21). They confirm the patterns observed in regression (20), where the only exception is that, at longer horizons, for the GBP base, the point estimates for $\beta_{EHf,n}$ are equally negative in both periods.

This result is interesting in light of the growing literature in the United States arguing that QE worked by lowering the term premium at longer horizons (See Abrahams et al. (2015)). Given our results above, this suggests that holding the effect of QE on interest rate expectations fixed, the lower term premia due to QE was associated with an appreciation rather than with a depreciation of the dollar.

In order to identify which component of the exchange rate change can explain the observed patterns as defined in equation (6), we further decompose the estimates from regressions (14), (16), (20), and (21) as follows. Denote by β the vector of coefficients from any one of these four regressions and by \mathbf{X} the time by variable matrix of independent variables used in the respective regression. Then, we can use our decomposition of exchange rate changes to decompose the coefficients as follows:

$$\begin{aligned}\beta &= (\mathbf{X}'\mathbf{X})^{-1} \mathbf{X}' \Delta s_{t+1}^{i,j} \\ &= \beta_{\tilde{u}_t, \mathbf{X}} + \beta_{-\varphi_{t+1}^{EH}, \mathbf{X}} + \beta_{\sigma_t - \sigma_{t+1}^F, \mathbf{X}} + \beta_{s_{t+1, \infty}^{\Delta E}, \mathbf{X}} \\ \text{where } \beta_{y, \mathbf{X}} &\equiv (\mathbf{X}'\mathbf{X})^{-1} \mathbf{X}' y \text{ for } y \in \{\tilde{u}_t, -\varphi_{t+1}^{EH}, \sigma_t - \sigma_{t+1}^F, s_{t+1, \infty}^{\Delta E}\}.\end{aligned}$$

In the case of a single, non-constant regressor, as is the case in regressions (14) and (16),

$$\beta_{y,\mathbf{x}} = \frac{Cov(y, x)}{Var(x)} \text{ for } y \in \{\tilde{v}_t, -\varphi_{t+1}^{EH}, \sigma_t - \sigma_{t+1}^F, s_{t+1, \infty}^{\Delta E}\} \text{ and } x \in \{n\Delta\tilde{y}_{t+1}^n, \Delta\tilde{f}_{t+1}^{n, n+m}\}.$$

By looking at the combined term $\sigma_t - \sigma_{t+1}^F$, rather than at σ_t and σ_{t+1}^F separately, the exact specification of the exchange rate equation of the VAR is not important, since $\sigma_t - \sigma_{t+1}^F$ can be calculated as a residual from equation (6), using only expectations based on the interest rate and inflation equations of the VAR. We chose to use such an approach, given that it is difficult to discipline the exchange rate forecast equation due to lack of consensus on the model used by traders. Exchange rate forecast data are often noisy and unreliable.¹⁴

First let us focus on the decomposition of the coefficients in regressions (14) and (16). Figures 5 through 12 report the results. Figure 5 reveals that $\frac{Cov(-\varphi_{t+1}^{EH}, n\Delta\tilde{y}_{t+1}^n)}{Var(n\Delta\tilde{y}_{t+1}^n)}$ is always negative but slightly more negative prior to 2009:Q1 for all maturities and all currency bases. This pattern can be seen much more clearly for $\frac{Cov(-\varphi_{t+1}^{EH}, \Delta\tilde{f}_{t+1}^{n, n+4})}{Var(\Delta\tilde{f}_{t+1}^{n, n+4})}$ for the USD base and at the short end of the curve for the DEM/EUR and GBP bases. In terms of magnitudes, $\frac{Cov(-\varphi_{t+1}^{EH}, n\Delta\tilde{y}_{t+1}^n)}{Var(n\Delta\tilde{y}_{t+1}^n)}$ ranges from -13.7 (at the short end) to -0.17 (at the long end), while $\frac{Cov(-\varphi_{t+1}^{EH}, \Delta\tilde{f}_{t+1}^{n, n+4})}{Var(\Delta\tilde{f}_{t+1}^{n, n+4})}$ ranges from -4.9 at the short end to 0.4 at the long end, where the positive point estimates are not statistically different from zero.

Similarly, one can see in Figures 7 and 8 that both $\frac{Cov(\tilde{v}_t, n\Delta\tilde{y}_{t+1}^n)}{Var(n\Delta\tilde{y}_{t+1}^n)}$ and $\frac{Cov(\tilde{v}_t, \Delta\tilde{f}_{t+1}^{n, n+4})}{Var(\Delta\tilde{f}_{t+1}^{n, n+4})}$ are negative for almost all n and more negative prior to the ZLB period. However, the regression coefficients are all less than 1 in absolute value and hence do not contribute much to the coefficients in regressions (14) and (16).

$\frac{Cov(\sigma_t - \sigma_{t+1}^F, n\Delta\tilde{y}_{t+1}^n)}{Var(n\Delta\tilde{y}_{t+1}^n)}$ and $\frac{Cov(\sigma_t - \sigma_{t+1}^F, \Delta\tilde{f}_{t+1}^{n, n+4})}{Var(\Delta\tilde{f}_{t+1}^{n, n+4})}$ also exhibit interesting patterns as can be seen in Figures 9 and 10. Prior to the crisis, all of the coefficients, $\frac{Cov(\sigma_t - \sigma_{t+1}^F, n\Delta\tilde{y}_{t+1}^n)}{Var(n\Delta\tilde{y}_{t+1}^n)}$, are positive

¹⁴Dominguez (1986), Frankel and Froot (1987), and Ito (1988) were some of the first studies testing rationality using survey data on exchange rate forecasts and they all find evidence of violations of rational expectations, including bias. Macdonald and Marsh (1996) finds evidence of a large degree of heterogeneity among exchange rate forecasts made by different individuals. Using more recent data, Ince and Molodtsova (2015) confirms that forecasts for major currencies still tend to fail rationality tests.

and many are statistically significant. After the crisis, all but one of the coefficients becomes very negative and all but three are statistically different from zero. The same pattern emerges for $\frac{Cov(\sigma_t - \sigma_{t+1}^F, \Delta \hat{f}_{t+1}^{n, n+4})}{Var(\Delta \hat{f}_{t+1}^{n, n+4})}$ at the short end of the curve. These coefficients are also large and economically significant relative to the size of the estimated coefficients in regressions (14) and (16).

This result implies that prior to the ZLB period, an increase in the yield of country i relative to the base country led to a depreciation of currency i through its correlation with the currency risk premia term, while during the ZLB period the same increase led to an appreciation in currency i due to a reversal in the relationship between yields and the currency risk premia term. In results not presented here, we confirm that this reversal is driven predominantly by the σ_{t+1}^F term. This implies that prior to the ZLB, an increase in the yield of country i relative to j was correlated with market expectations of lower future excess returns from being long the country j bond. However, this relationship flipped during the ZLB period.

One possible explanation for this is that prior to the ZLB period, the expected excess return was driven largely by flight-to-safety dynamics, where yields were elevated in non-base countries and further increases in rates implied a higher probability of future currency crashes in those countries (see Brunnermeier, Nagel, and Pedersen (2009)). Therefore, higher yields in country i would coincide with higher future expected excess returns from investing in bonds denominated in currency i . However, during the ZLB period, yields in non-base countries were more subdued. To the extent that there are nonlinearities in the relationship between the levels of an individual country's yields and the risk of a crash, the relationship between relative yields and crash risk in non-base countries may have weakened in the ZLB period. Additionally, investors engaging in reaching-for-yield behavior during the ZLB period could also have demanded lower expected excess returns for a given increase in crash risk associated with a change in relative yields. Further work will be required in order to test these potential explanations.

Finally, we consider the last component of the decomposition—changes in expectations over long-run nominal exchange rates, $s_{t+1,\infty}^{\Delta E}$. Recall that our empirical model features constant expectations over long-run real exchange rates, so this term captures changes in expectations over the path of future inflation rates in country i relative to the base country. Figures 11 and 12 reveal the following patterns. Prior to the ZLB, $\frac{Cov(s_{t+1,\infty}^{\Delta E}, n\Delta\tilde{y}_{t+1}^n)}{Var(n\Delta\tilde{y}_{t+1}^n)}$ and $\frac{Cov(s_{t+1,\infty}^{\Delta E}, \Delta\tilde{f}_{t+1}^{n,n+4})}{Var(\Delta\tilde{f}_{t+1}^{n,n+4})}$ are positive and in some cases statistically significant, while during the ZLB period they are even more positive and almost always statistically significant. The size of the estimated coefficients is also large and has a significant impact on the coefficients in regressions (14) and (16). This implies that in both subsamples, an increase of the yield in country i relative to the base country implied higher inflation expectations in the future in country i relative to the base country.

This is consistent with the idea that yields, which are highly influenced by policy rates, can contain information relevant to future inflation to the extent that policy rates respond to economic fundamentals and the central bank has private information regarding inflation. The fact that the coefficient is much more positive during the ZLB period can be potentially attributed to a stronger effect from this type of signaling in the ZLB period. Campbell et al. (2012) provides evidence showing that both forward guidance and QE statements had these effects (which they call "Delphic forward guidance"). In fact, their point estimates indicate that forward guidance and QE statements had even stronger signaling effects on inflation and unemployment forecasts in the ZLB period, although the estimates are rather imprecise in this period due to their analyses using only four years of data. Our results also indirectly support such a hypothesis. Another potential cause of this structural break is that the ZLB period is characterized by a greater prevalence of demand shocks that drive a positive conditional correlation between expected future inflation and expected future short-term rate increases due to a policy response. A third possibility is that monetary policy being constrained by the zero lower bound is itself the cause of a change in the relationship between interest rate changes and inflation expectations due to a change in the monetary

policy response to expected inflation.

To summarize, the decomposition presented in section 3 allows us to see that the more negative $\hat{\beta}_{y,n}$ and $\hat{\beta}_{f,n}$ estimates (from regressions (14) and (16)) at the short end of the yield curve during the ZLB period are due to a structural break in the relationship between the expected excess returns term, $\sigma_t - \sigma_{t+1}^F$, and changes in relative yields/forwards. The subsample estimates with respect to the other components of decomposition (6) all push in the opposite direction, suggesting that $\hat{\beta}_{y,n}$ and $\hat{\beta}_{f,n}$ should have become more positive at the short maturities/horizons. The structural breaks in these latter components are dominated by the break with respect to the expected excess return component, thus generating the overall pattern at the short end that we observed in Figures 1 and 2.

The opposite is true when one considers the long end of the yield curve or long-horizon forward rates. In that case, the fact that the $\hat{\beta}_{y,n}$ and $\hat{\beta}_{f,n}$ estimates are less negative or even positive during the ZLB period can be attributed to the structural breaks with respect to the expectations hypothesis term, $-\varphi_{t+1}^{EH}$, the term capturing changes in expectations over long-run relative prices, $s_{t+1,\infty}^{\Delta E}$, and the UIRP term, $\tilde{\iota}_t$, to a lesser extent.

Finally, we turn to the role played by individual components of exchange rate changes in the estimates of (20) and (21). Figures 13 through 20 present the results. The patterns here largely echo the decomposition of $\hat{\beta}_{y,n}$ and $\hat{\beta}_{f,n}$. The more negative estimates of $\beta_{EHy,n}$ and $\beta_{EHf,n}$ in the ZLB period for small n are attributable to changes in the relationship between the expected excess currency returns term, $\sigma_t - \sigma_{t+1}^F$, and changes in the expectations hypothesis and term premia components of yields/forwards as measured by $\beta_{\sigma_t - \sigma_{t+1}^F, \mathbf{X}}$. The changes with respect to the remaining components of exchange rates generally push in the other direction. For the long end, the more positive estimates of $\beta_{EHy,n}$ and $\beta_{EHf,n}$ in the ZLB period are explained by breaks in the relationship between changes in expectations over long-run prices, $s_{t+1,\infty}^{\Delta E}$, and changes in the expectations hypothesis and term premia components of yields/forwards. The expectations hypothesis term, $-\varphi_{t+1}^{EH}$, plays a role that

is qualitatively similar, but of a much smaller magnitude.

5 Conclusion

In this paper, we examine the contemporaneous relationship between quarterly nominal exchange rate changes and changes in yields and forward rates at different maturities/horizons. We study 10 developed countries with government debt that is considered to be free of default risk and pairs with respect to three base currencies: the USD, GBP, and DEM/EUR. We estimate both bilateral regressions and panel regressions where data are pooled across pairs for each base.

We find that, in general, the coefficients from univariate regressions of exchange rate changes on bond yield changes at maturities of less than three years are generally consistent with those predicted by a model where: (1) short-term interest rates follow an AR(1) process and (2) there is no omitted variable bias on net from correlations of yield changes with expected excess currency returns, changes in term premia, or changes in expectations over long-run nominal exchange rates. That is to say, we find that a relative increase in a country's short-term bond yield coincides with a contemporaneous, statistically significant appreciation in that country's currency, as folk wisdom dictates. However, a relative increase in a country's long-term bond yield results in a much smaller appreciation than predicted by the simple model. It even leads to a slight depreciation for certain currency pairs. Using subsample analysis, we find that this pattern in the coefficients becomes more pronounced during the ZLB period starting in 2009:Q1—in other words, the coefficient curve twists. In this period, increases in relative one-year forward rates at horizons of five years or more lead to statistically significant depreciation in our panel regressions for all three currency bases.

To better understand the drivers of these patterns, we use expectations implied by an estimated VAR to decompose both yields/forwards and exchange rate changes. Each yield or forward rate is decomposed into an expectations hypothesis component, which is a function

of expectations of future short-term interest rates and a term premium component. Exchange rates are decomposed into an analogous expectations hypothesis component, a component that is a function of expected one-period excess returns at different points in time, and a term reflecting changes in expectations over long-run nominal exchange rate levels. When the real exchange rate is stationary, this last term is equivalent to changes in expectations over long-run relative price levels.

Using the bond yield and forward rate decompositions, we find that the structural break between the pre-ZLB and ZLB periods in the relationship between changes in exchange rates and yields/forwards is driven by a break in the relationship of exchange rates with the expectations hypothesis component of yields/forwards at short maturities and with the term premia components at longer maturities. In other words, during the ZLB period, an increase in the relative expectations hypothesis term was associated with an even larger exchange rate appreciation, while an increase in the relative term premia, especially at long maturities, coincides with an exchange rate depreciation.

Using the exchange rate decomposition, we offer an alternate accounting of the drivers behind the twisting of the coefficient curve. More specifically, the structural break at the short end of the curve can be attributed to the fact that prior to the ZLB period, a cross-currency bond position that enjoyed an increase in yields was associated with higher contemporaneous and future expected excess currency returns, which lower the current value of the currency. This relationship flips signs during the ZLB period. At the long end of the curve, an increase in relative yields/forwards was associated with higher expected relative long run price levels and smaller increases in expected future relative short-term rates during the ZLB period.

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Table 1: Change in Relative Short-Term Yields, 1990:Q1–2014:Q3

§ base currency	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	GBP
$\Delta_{t+1}^{s,i,\$}$	-8.617***	1.624	-9.053	-4.539	-7.186	-7.089**	-4.337	-5.991	-15.795***
	[2.822]	[1.866]	[7.496]	[4.878]	[7.018]	[3.472]	[2.860]	[4.255]	[4.869]
constant	-0.323	-0.044	-0.609	-0.182	-0.410	-0.295	-0.316	-0.259	-0.300
	[0.581]	[0.365]	[0.569]	[0.554]	[0.791]	[0.684]	[0.557]	[0.586]	[0.434]
# Observations	99	99	89	84	66	65	97	86	99
Adjusted R^2	0.041	-0.006	0.030	0.001	0.004	0.051	0.012	0.026	0.199
RMSE	5.658	3.582	5.403	4.997	6.414	5.472	5.471	5.185	4.376

£ base currency	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	USD
$\Delta_{t+1}^{s,i,\pounds}$	-17.946***	-5.131*	-10.064	-12.894*	-32.038***	-5.105**	-10.130***	-7.292	-15.795***
	[2.356]	[2.813]	[8.727]	[7.358]	[9.765]	[2.415]	[2.383]	[4.725]	[4.869]
constant	-0.160	0.018	-0.613	-0.064	0.023	-0.257	-0.234	-0.123	0.300
	[0.510]	[0.465]	[0.551]	[0.463]	[0.766]	[0.490]	[0.522]	[0.448]	[0.434]
# Observations	99	99	89	84	66	65	97	86	99
Adjusted R^2	0.228	0.030	0.055	0.066	0.323	0.035	0.107	0.029	0.199
RMSE	5.072	4.643	5.189	4.115	6.229	3.916	5.151	4.164	4.376

DEM/€ base currency	AUD	CAD	CHF	JPY	NOK	NZD	SEK	GBP	USD
$\Delta_{t+1}^{s,i,DEM/\pounds}$	-11.189**	6.078	0.600	-10.583	-1.817	-8.161**	0.349	-12.894*	-4.539
	[5.221]	[5.341]	[3.298]	[10.213]	[4.246]	[3.952]	[4.746]	[7.358]	[4.878]
constant	-0.153	-0.154	-0.410*	0.246	-0.006	-0.163	-0.059	0.064	0.182
	[0.634]	[0.562]	[0.246]	[0.714]	[0.452]	[0.579]	[0.359]	[0.463]	[0.554]
# Observations	84	84	84	63	65	84	84	84	84
Adjusted R^2	0.036	0.018	-0.011	0.019	-0.011	0.051	-0.012	0.066	0.001
RMSE	5.814	5.145	2.280	5.778	3.645	5.257	3.282	4.115	4.997

Frequency: Quarterly

Dependent variable: $\Delta s_{t+1}^{i,j}$ Quarterly log exchange rate change (increase implies base currency appreciation)

Independent variable: Change in relative three-month yields (not annualized).

Heteroskedasticity-robust standard errors are in brackets.

***/**/* Statistically significant at 1, 5, and 10 percent, respectively.

Table 2: Change in Relative 5-Year Yields, 1990:Q1–2014:Q3

§ base currency	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	GBP
$n\Delta\tilde{y}_{t+1}^{20,i,\S}$	-0.623*	0.171	-1.008***	-1.012***	-0.795***	-0.245	-0.087	-0.436**	-0.836***
	[0.327]	[0.170]	[0.291]	[0.310]	[0.284]	[0.220]	[0.192]	[0.191]	[0.239]
constant	-0.289	-0.042	-0.592	-0.289	-0.395	-0.323	-0.295	-0.248	-0.159
	[0.574]	[0.365]	[0.535]	[0.520]	[0.755]	[0.691]	[0.563]	[0.564]	[0.448]
# Observations	99	99	89	84	66	65	97	86	99
Adjusted R^2	0.041	-0.003	0.164	0.111	0.089	0.008	-0.009	0.031	0.135
RMSE	5.659	3.578	5.017	4.715	6.134	5.594	5.528	5.171	4.547
£ base currency	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	USD
$n\Delta\tilde{y}_{t+1}^{20,i,\pounds}$	-1.114***	-1.026***	-0.968*	-0.635	-1.016	-0.128	-0.376	-0.413	-0.836***
	[0.188]	[0.273]	[0.488]	[0.384]	[0.615]	[0.164]	[0.270]	[0.309]	[0.239]
constant	-0.190	0.024	-0.575	-0.024	-0.185	-0.309	-0.294	-0.080	0.159
	[0.515]	[0.434]	[0.533]	[0.458]	[0.872]	[0.491]	[0.551]	[0.449]	[0.448]
# Observations	99	99	89	84	66	65	97	86	99
Adjusted R^2	0.211	0.153	0.097	0.030	0.094	-0.005	0.012	0.022	0.135
RMSE	5.126	4.338	5.075	4.195	7.208	3.997	5.420	4.180	4.547
DEM/€ base currency	AUD	CAD	CHF	JPY	NOK	NZD	SEK	GBP	USD
$n\Delta\tilde{y}_{t+1}^{20,i,DEM/\pounds}$	-1.045***	-0.436	-0.466*	-0.636	-0.095	-0.851***	0.552*	-0.635	-1.012***
	[0.313]	[0.406]	[0.238]	[0.579]	[0.125]	[0.295]	[0.321]	[0.384]	[0.310]
constant	-0.163	-0.088	-0.366	0.134	0.003	-0.095	-0.028	0.024	0.289
	[0.609]	[0.562]	[0.235]	[0.711]	[0.455]	[0.581]	[0.344]	[0.458]	[0.520]
# Observations	84	84	84	63	65	84	84	84	84
Adjusted R^2	0.101	0.006	0.051	0.034	-0.010	0.065	0.067	0.030	0.111
RMSE	5.614	5.177	2.209	5.735	3.643	5.218	3.151	4.195	4.715

Frequency: Quarterly; $n = 20$

Dependent variable: $\Delta s_{t+1}^{i,j}$ Quarterly log exchange rate change (increase implies base currency appreciation).

Independent variable: Change in relative 5-year quarterly yields (not annualized).

Heteroskedasticity-robust standard errors are in brackets.

***/**/* Statistically significant at 1, 5, and 10 percent, respectively.

Table 3: Change in Relative 10-Year Yields, 1990:Q1–2014:Q3

\$ base currency	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	GBP
$n\Delta\tilde{y}_{t+1}^{40,i,\$}$	-0.179 [0.175]	0.266* [0.135]	-0.384** [0.175]	-0.385* [0.199]	-0.491** [0.185]	-0.018 [0.140]	0.013 [0.119]	-0.149 [0.117]	-0.220 [0.150]
constant	-0.234 [0.581]	-0.013 [0.365]	-0.516 [0.565]	-0.233 [0.546]	-0.434 [0.754]	-0.331 [0.701]	-0.287 [0.563]	-0.206 [0.576]	-0.090 [0.480]
# Observations	99	99	89	84	66	65	97	86	99
Adjusted R^2	0.003	0.031	0.063	0.038	0.102	-0.016	-0.010	0.002	0.017
RMSE	5.770	3.517	5.308	4.905	6.092	5.660	5.532	5.247	4.847
£ base currency	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	USD
$n\Delta\tilde{y}_{t+1}^{40,i,£}$	-0.493*** [0.137]	-0.447** [0.201]	-0.133 [0.253]	0.089 [0.203]	-0.541* [0.314]	0.020 [0.089]	0.115 [0.124]	0.063 [0.147]	-0.220 [0.150]
constant	-0.208 [0.551]	0.003 [0.459]	-0.603 [0.565]	0.012 [0.470]	-0.180 [0.878]	-0.322 [0.497]	-0.347 [0.557]	0.001 [0.446]	0.090 [0.480]
# Observations	99	99	89	84	66	65	97	86	99
Adjusted R^2	0.091	0.054	-0.005	-0.010	0.081	-0.015	-0.003	-0.009	0.017
RMSE	5.503	4.585	5.353	4.280	7.257	4.016	5.461	4.246	4.847
DEM/€ base currency	AUD	CAD	CHF	JPY	NOK	NZD	SEK	GBP	USD
$n\Delta\tilde{y}_{t+1}^{40,i,DEM/€}$	-0.362* [0.195]	-0.022 [0.236]	-0.077 [0.149]	-0.406 [0.326]	-0.007 [0.072]	-0.247 [0.167]	0.370** [0.168]	0.089 [0.203]	-0.385* [0.199]
constant	-0.219 [0.633]	-0.104 [0.568]	-0.393 [0.246]	0.077 [0.722]	-0.004 [0.454]	-0.184 [0.598]	-0.012 [0.339]	-0.012 [0.470]	0.233 [0.546]
# Observations	84	84	84	63	65	84	84	84	84
Adjusted R^2	0.028	-0.012	-0.007	0.048	-0.016	0.015	0.097	-0.010	0.038
RMSE	5.836	5.224	2.276	5.691	3.654	5.357	3.099	4.280	4.905

Frequency: Quarterly; $n = 40$

Dependent variable: $\Delta s_{t+1}^{i,j}$ Quarterly log exchange rate change (increase implies base currency appreciation).

Independent variable: Change in relative 10-year quarterly yields (not annualized).

Heteroskedasticity-robust standard errors are in brackets.

***/**/* Statistically significant at 1, 5, and 10 percent, respectively.

Table 4: Change in Relative 7-Year Forward Rates 3 Years Ahead, 1990:Q1–2014:Q3

\$ base currency	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	GBP
$\Delta \hat{f}_{t+1}^{12,40,i,\$}$	-0.028 [0.231]	0.429** [0.208]	-0.292 [0.218]	-0.217 [0.274]	-0.567** [0.259]	0.120 [0.221]	0.075 [0.169]	-0.056 [0.186]	0.031 [0.205]
Constant	-0.171 [0.582]	-0.011 [0.360]	-0.477 [0.579]	-0.169 [0.554]	-0.451 [0.770]	-0.328 [0.702]	-0.281 [0.562]	-0.153 [0.581]	-0.038 [0.492]
# Observations	99	99	89	84	66	65	97	86	99
Adjusted R^2	-0.010	0.049	0.012	-0.003	0.072	-0.008	-0.009	-0.011	-0.010
RMSE	5.808	3.482	5.451	5.007	6.191	5.639	5.527	5.281	4.914
£ base currency	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	USD
$\Delta \hat{f}_{t+1}^{12,40,i,\pounds}$	-0.227 [0.216]	-0.115 [0.269]	0.199 [0.288]	0.542** [0.217]	-0.531 [0.380]	0.114 [0.130]	0.336** [0.149]	0.321* [0.163]	0.031 [0.205]
Constant	-0.157 [0.576]	-0.017 [0.475]	-0.609 [0.567]	0.029 [0.457]	-0.219 [0.905]	-0.317 [0.496]	-0.354 [0.545]	0.041 [0.442]	0.038 [0.492]
# Observations	99	99	89	84	66	65	97	86	99
Adjusted R^2	0.001	-0.008	-0.003	0.038	0.035	-0.004	0.034	0.026	-0.010
RMSE	5.769	4.732	5.347	4.177	7.435	3.994	5.360	4.170	4.914
DEM/€ base currency	AUD	CAD	CHF	JPY	NOK	NZD	SEK	GBP	USD
$\Delta \hat{f}_{t+1}^{12,40,i,DEM/\pounds}$	-0.157 [0.275]	0.220 [0.357]	0.042 [0.165]	-0.438 [0.424]	0.067 [0.149]	-0.128 [0.205]	0.515** [0.209]	0.542** [0.217]	-0.217 [0.274]
Constant	-0.277 [0.646]	-0.108 [0.568]	-0.411 [0.247]	0.041 [0.741]	-0.015 [0.449]	-0.248 [0.598]	-0.008 [0.340]	-0.029 [0.457]	0.169 [0.554]
# Observations	84	84	84	63	65	84	84	84	84
Adjusted R^2	-0.008	-0.007	-0.011	0.027	-0.012	-0.007	0.101	0.038	-0.003
RMSE	5.944	5.211	2.280	5.755	3.647	5.415	3.093	4.177	5.007

Frequency: Quarterly; $n = 12$, $m = 28$

Dependent variable: $\Delta s_{t+1}^{i,j}$ Quarterly log exchange rate change (increase implies base currency appreciation).

Independent variable: Change in relative 1-year forward rates 5 years ahead.

Heteroskedasticity-robust standard errors are in brackets.

***/**/* Statistically significant at 1, 5, and 10 percent, respectively.

Table 5: Correlation between VAR-Implied and Market-Based Expected Monetary Policy Changes

	AUD	CAD	CHF	DEM/EUR	NOK	NZD	SED	GBP	USD
# Observations	0.72	0.39	0.62	0.74	0.21	0.74	0.66	0.78	0.75
	104	102	97	87	79	102	102	109	114

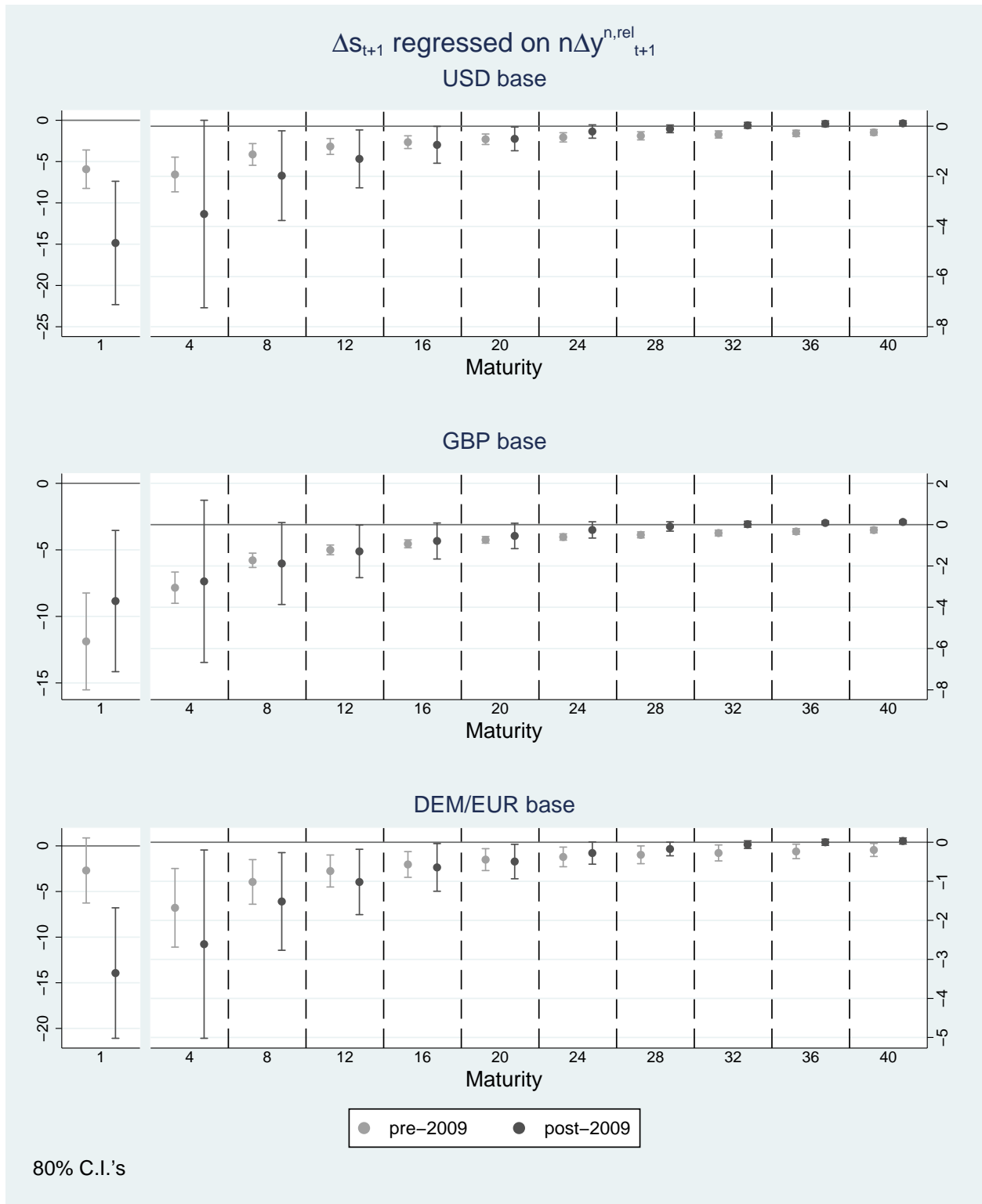
Frequency: Quarterly

Table 6: Correlation between VAR-Implied and Market-Based Expected Monetary Policy Surprises

	AUD	CAD	CHF	DEM/EUR	NOK	NZD	SED	GBP	USD
# Observations	0.83	0.65	0.49	0.84	-0.11	0.78	0.79	0.69	0.77
	104	103	97	88	80	103	103	109	114

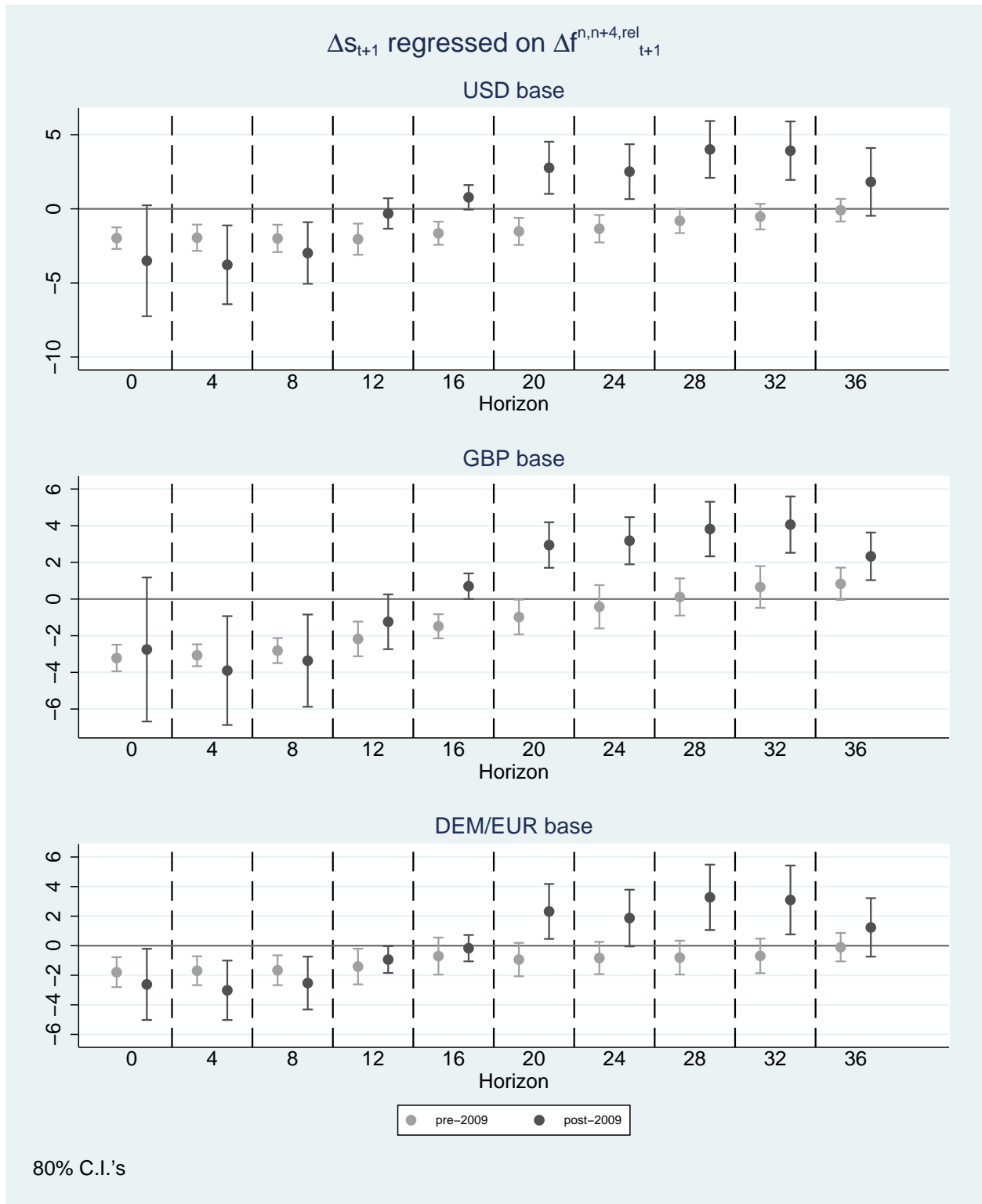
Frequency: Quarterly

Figure 1: Coefficients from Regressing Exchange Rate Changes on Relative Long-Term Yield Changes, 1990:Q1–2014:Q3



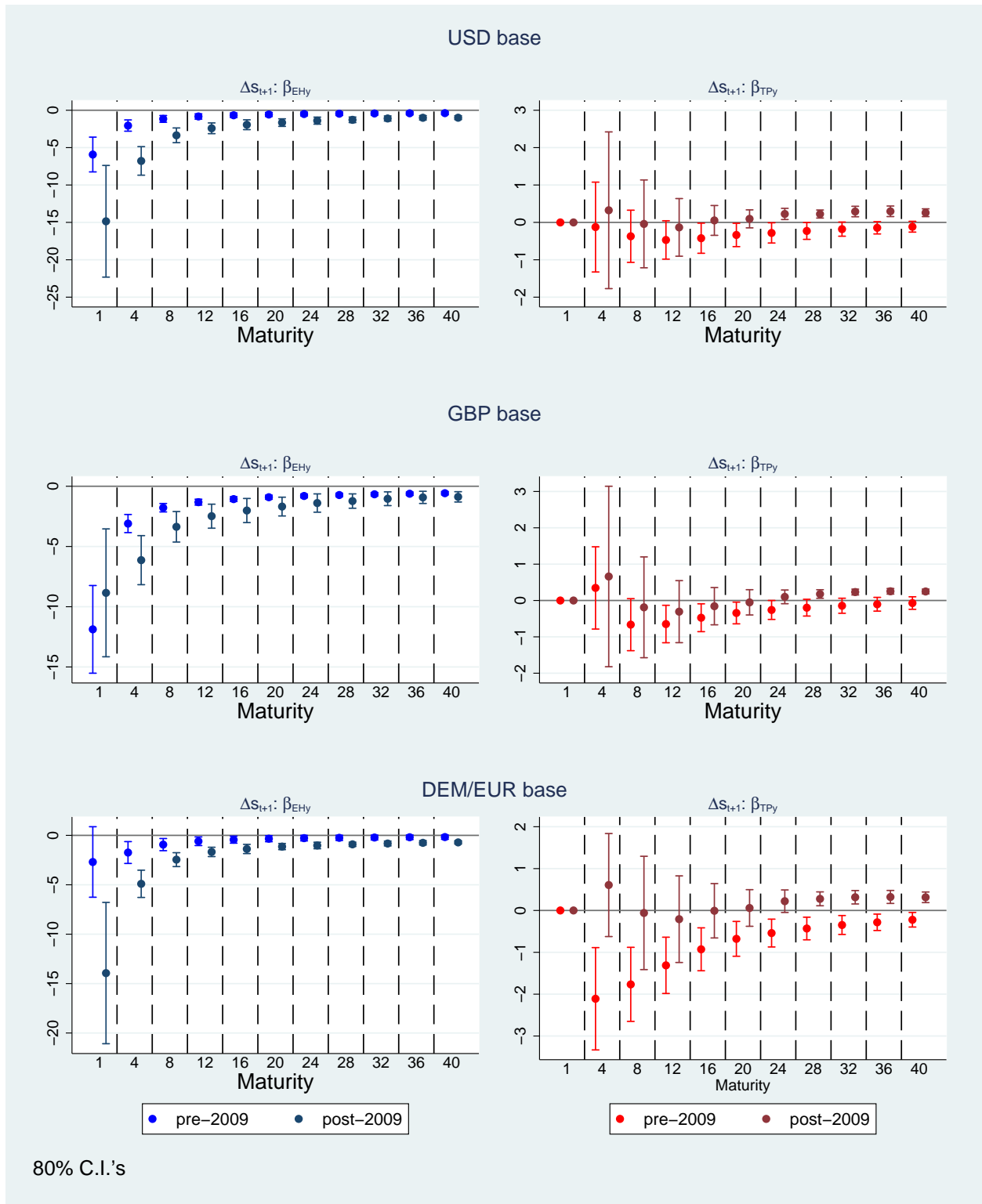
Sources: Authors' calculations.

Figure 2: Coefficients from Regressing Exchange Rate Changes on Relative 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



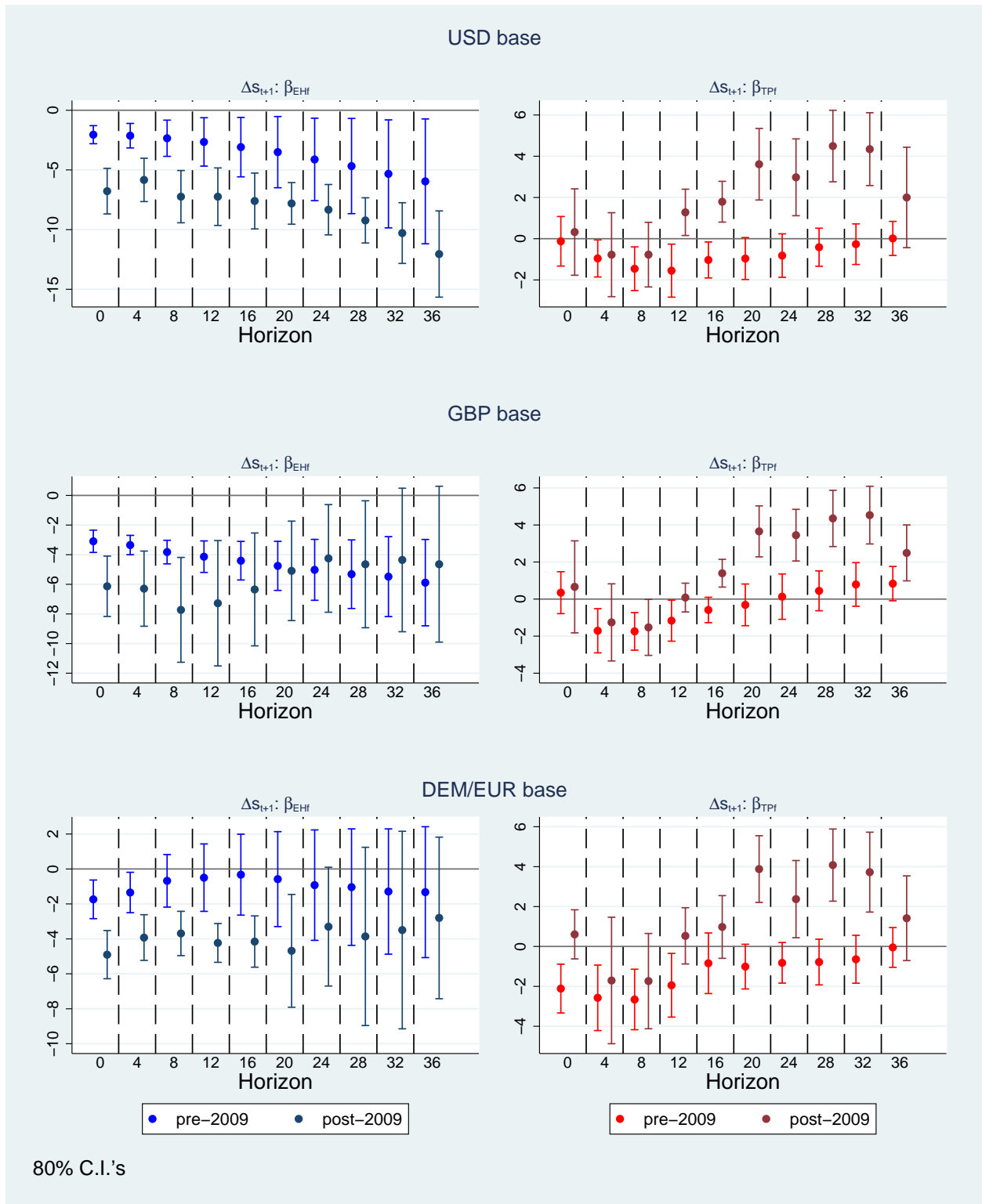
Sources: Authors' calculations.

Figure 3: Coefficients from Regressing Exchange Rate Changes on Relative Expectations Hypothesis and Term Premia Components of Long-Term Yield Changes, 1990:Q1–2014:Q3



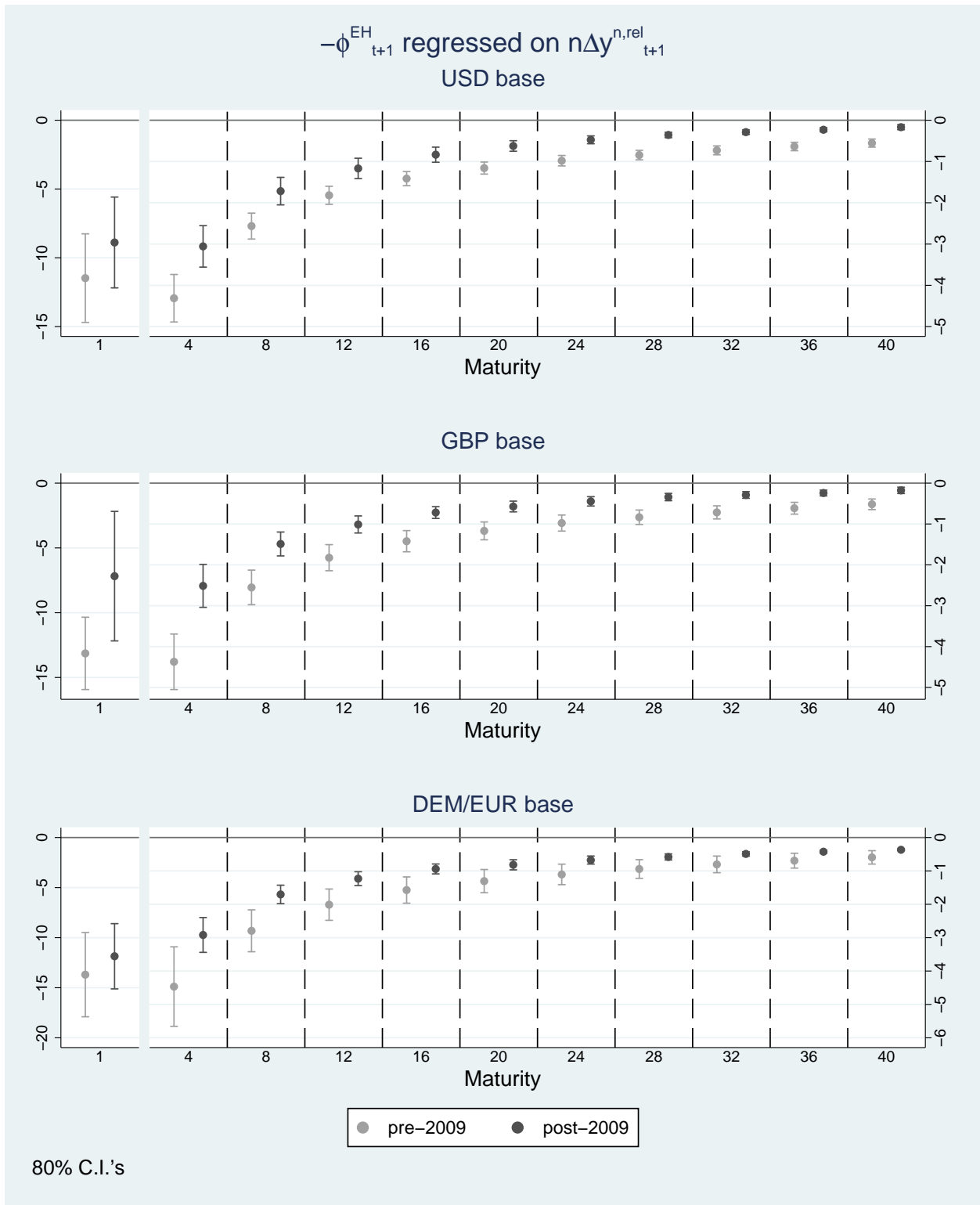
Sources: Authors' calculations.

Figure 4: Coefficients from Regressing Exchange Rate Changes on Relative Expectations Hypothesis and Term Premia Components of 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



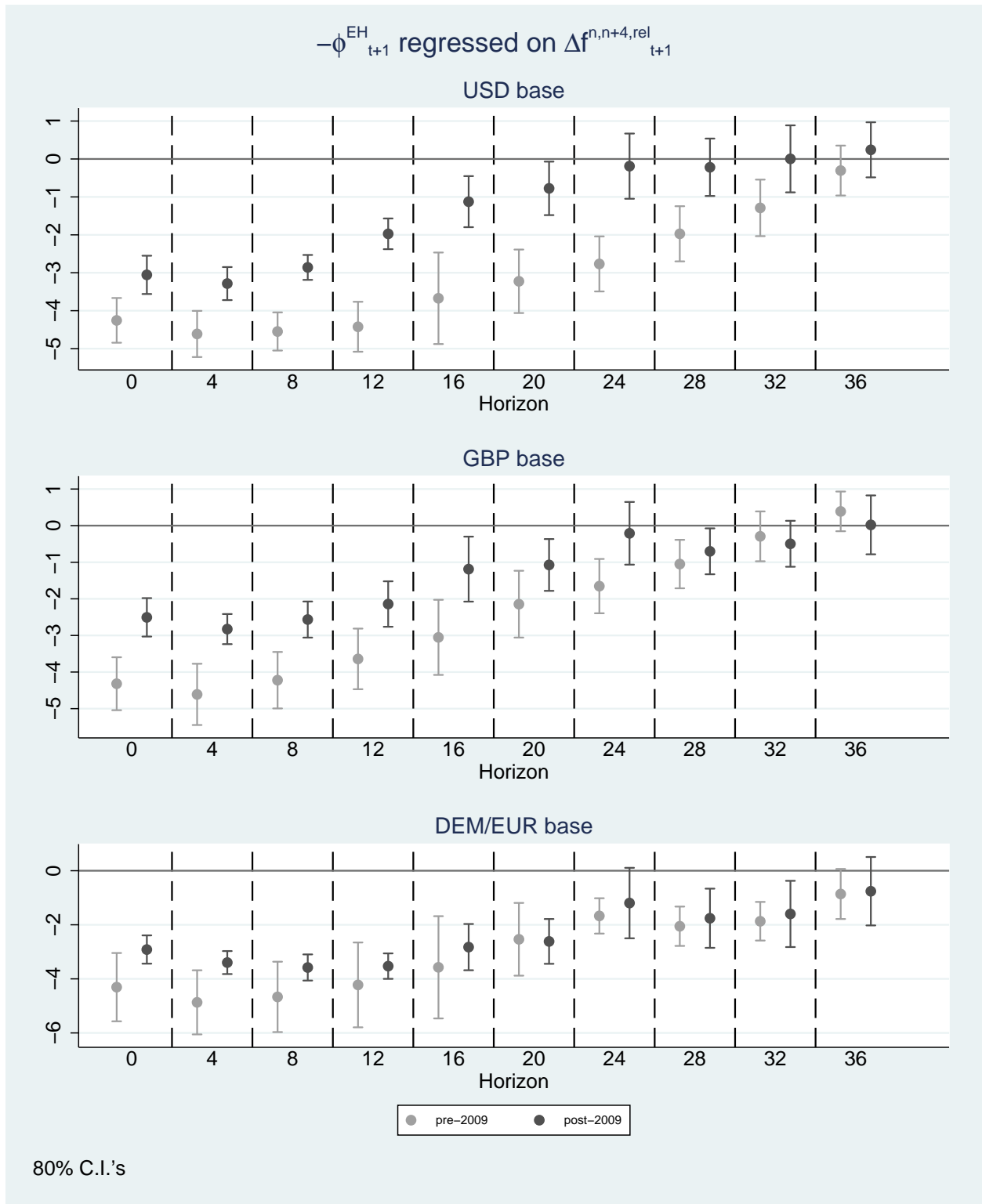
Sources: Authors' calculations.

Figure 5: Coefficients from Regressing $-\varphi_{t+1}^{EH}$ on Relative Long-Term Yield Changes, 1990:Q1–2014:Q3



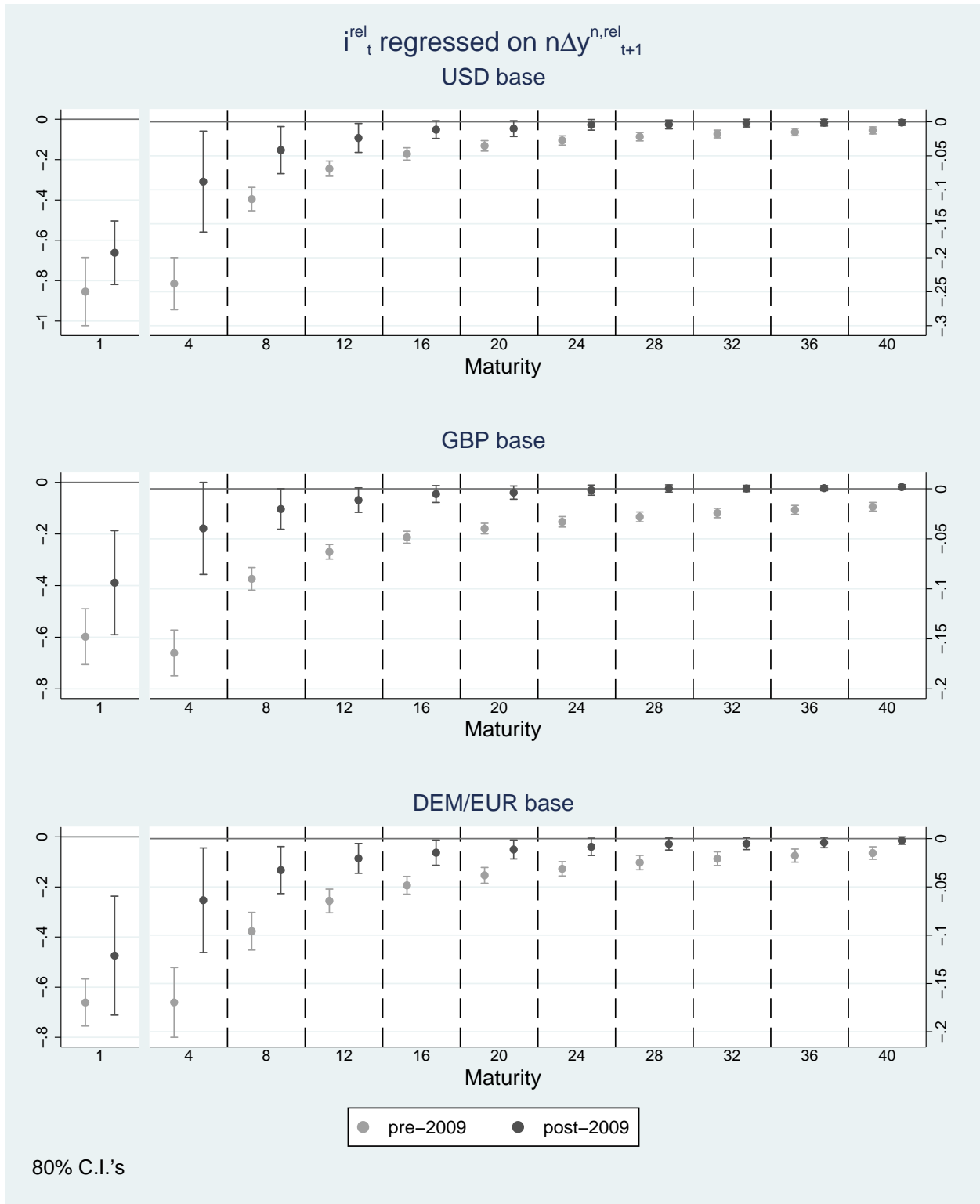
Sources: Authors' calculations.

Figure 6: Coefficients from Regressing $-\phi_{t+1}^{EH}$ on Relative 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



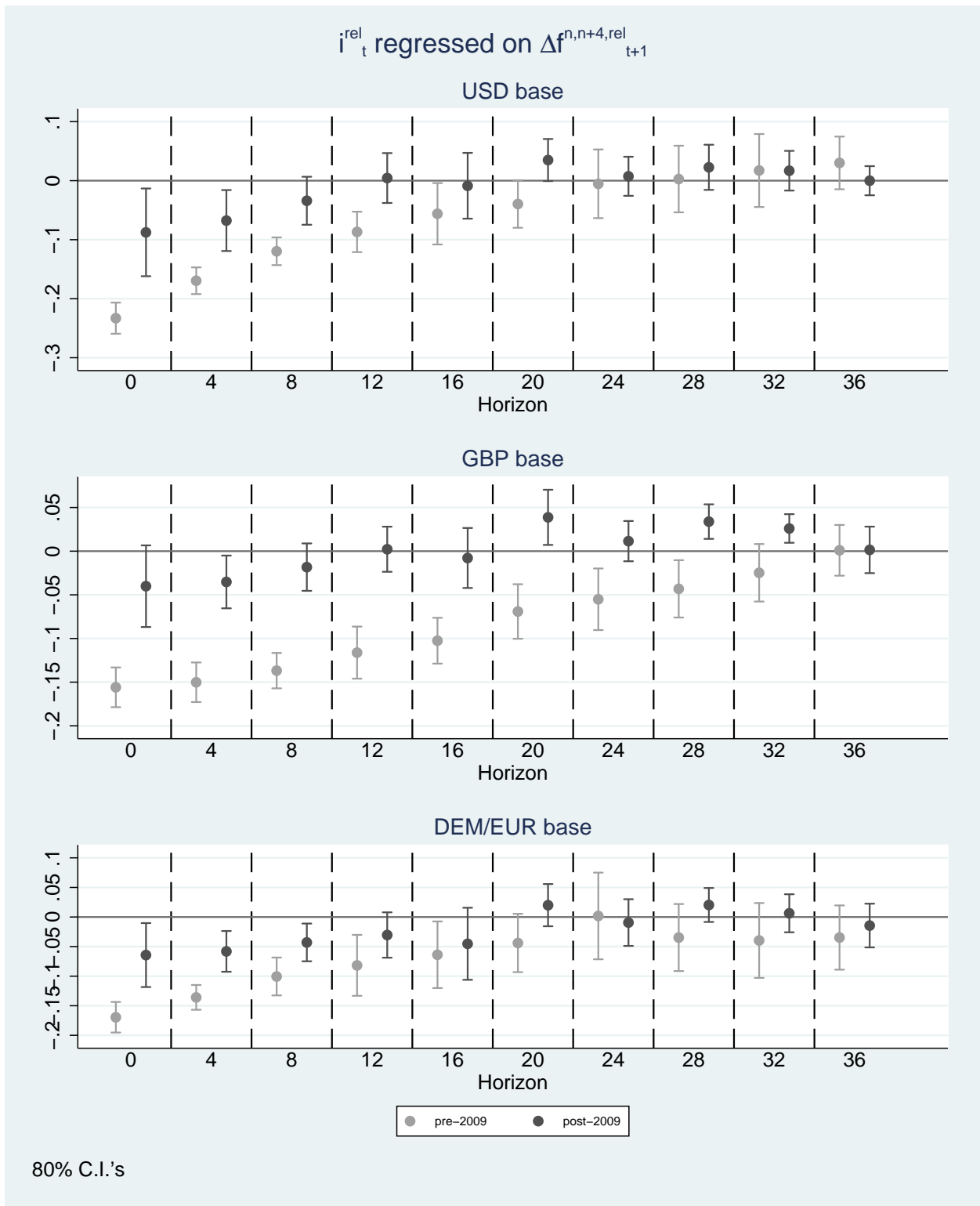
Sources: Authors' calculations.

Figure 7: Coefficients from Regressing \tilde{i}_t^{rel} on $n\Delta y_{t+1}^{\text{n,rel}}$, 1990:Q1–2014:Q3



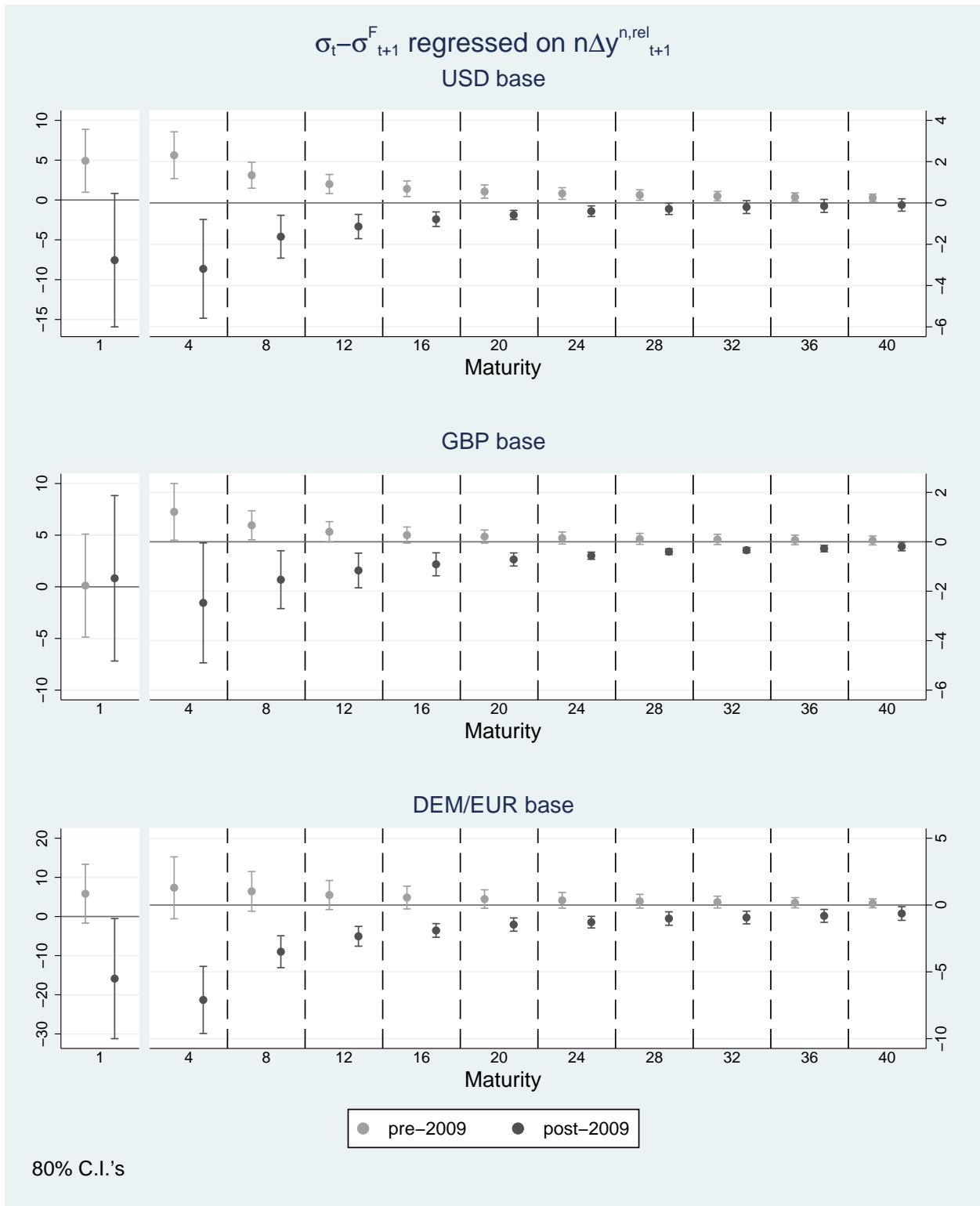
Sources: Authors' calculations.

Figure 8: Coefficients from Regressing \tilde{i}_t^{rel} on Relative 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



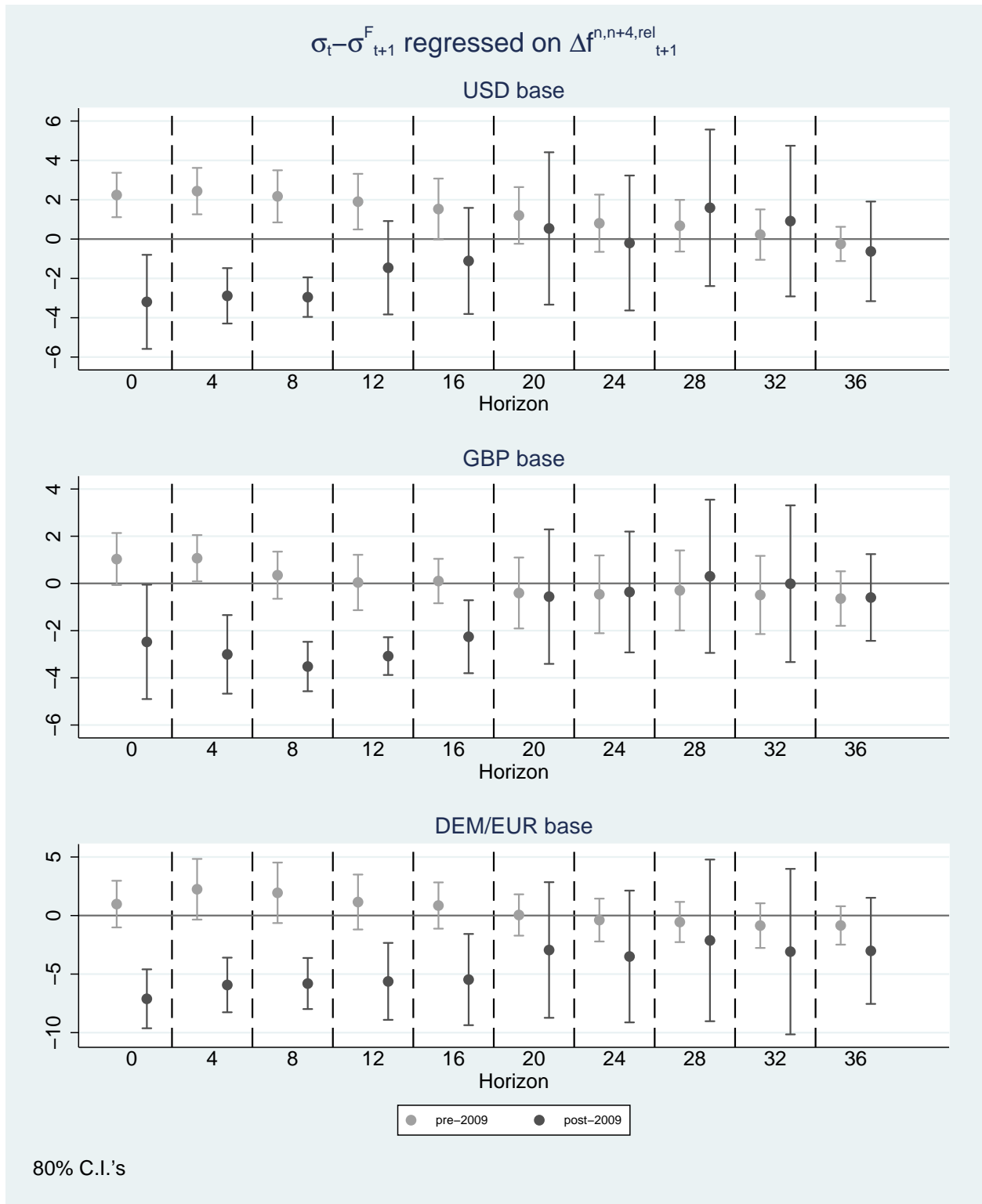
Sources: Authors' calculations.

Figure 9: Coefficients from Regressing $\sigma_t - \sigma_{t+1}^F$ on Relative Long-Term Yield Changes, 1990:Q1–2014:Q3



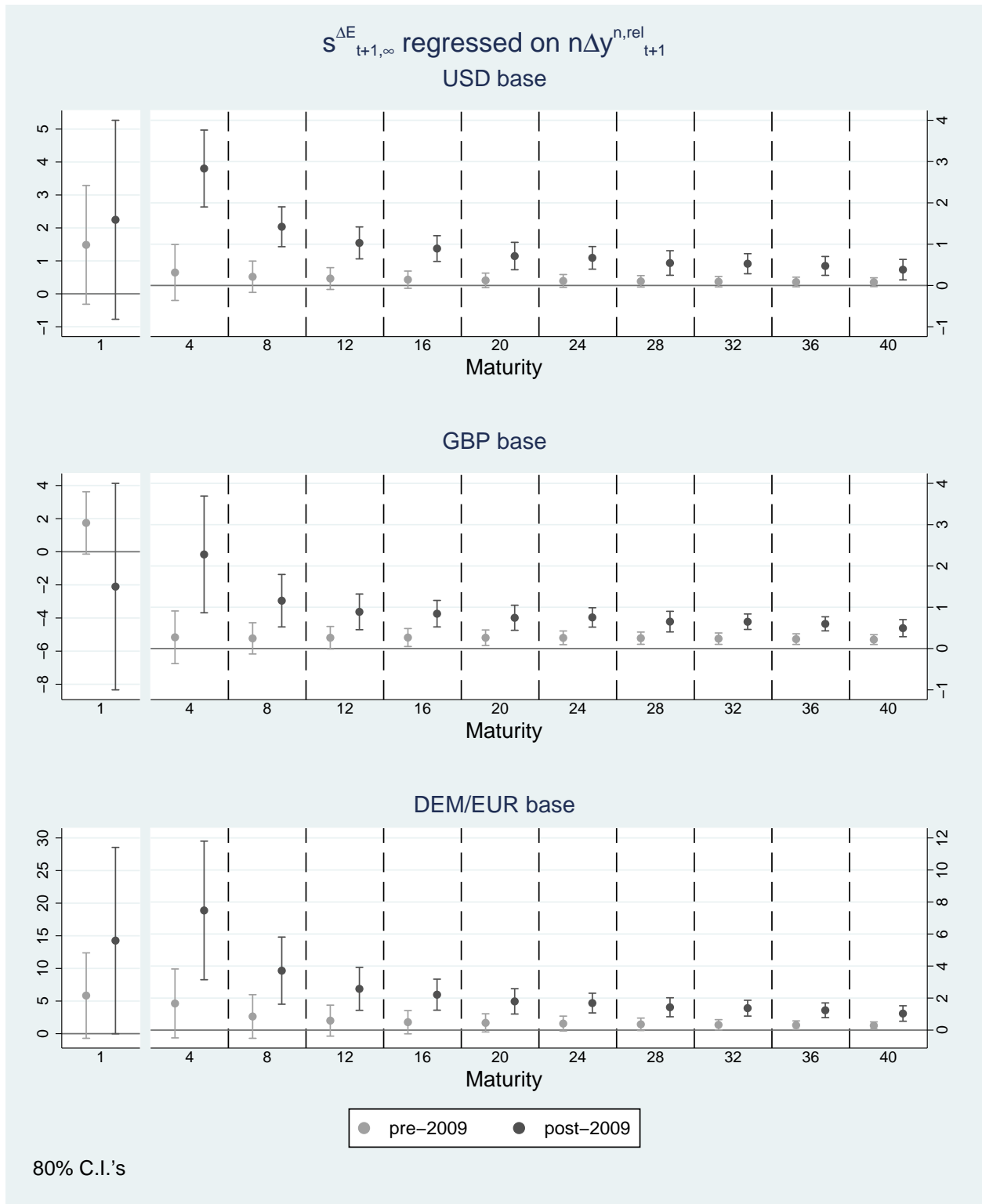
Sources: Authors' calculations.

Figure 10: Coefficients from Regressing $\sigma_t - \sigma_{t+1}^F$ on Relative 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



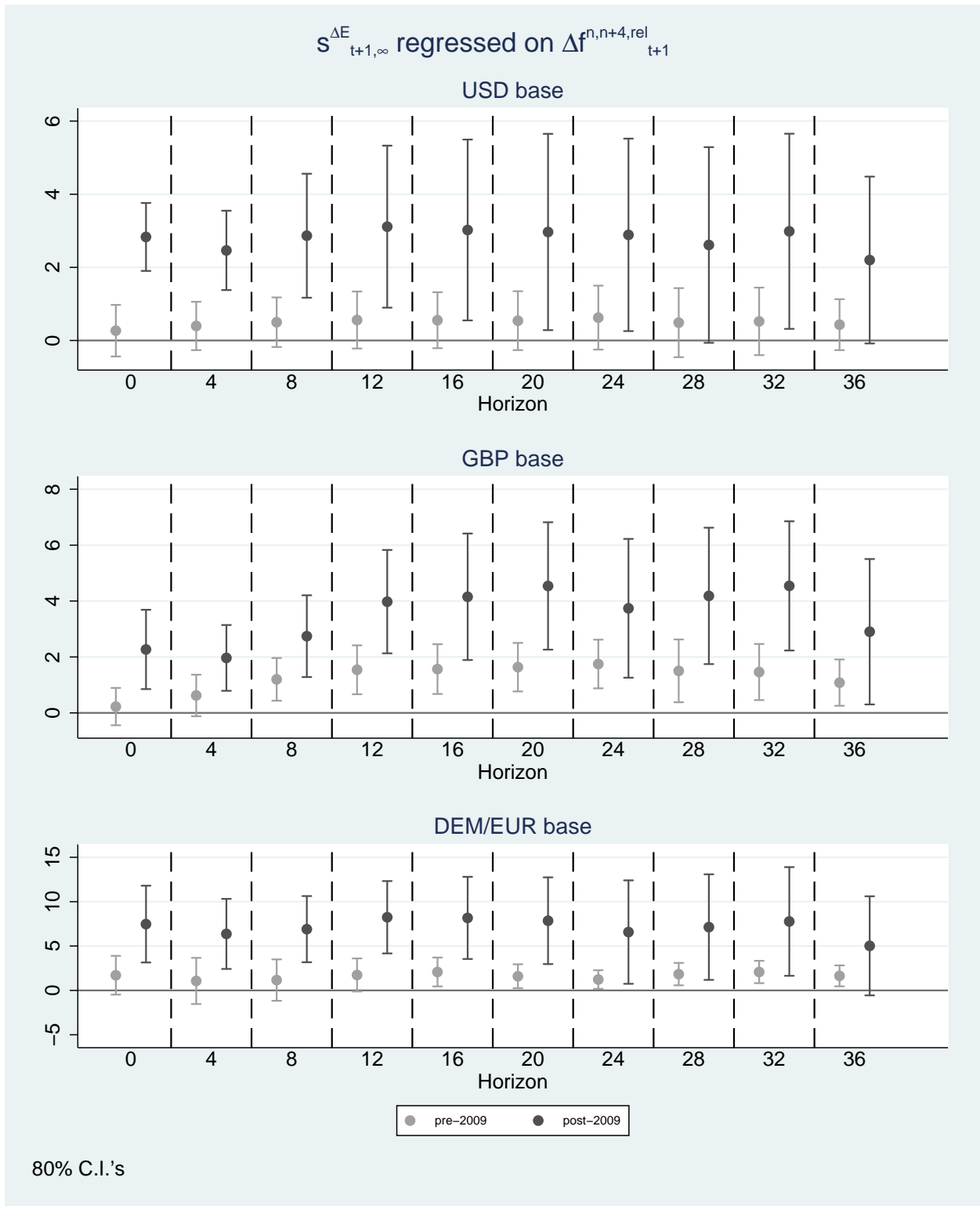
Sources: Authors' calculations.

Figure 11: Coefficients from Regressing $s_{t+1,\infty}^{\Delta E}$ on Relative Long-Term Yield Changes, 1990:Q1–2014:Q3



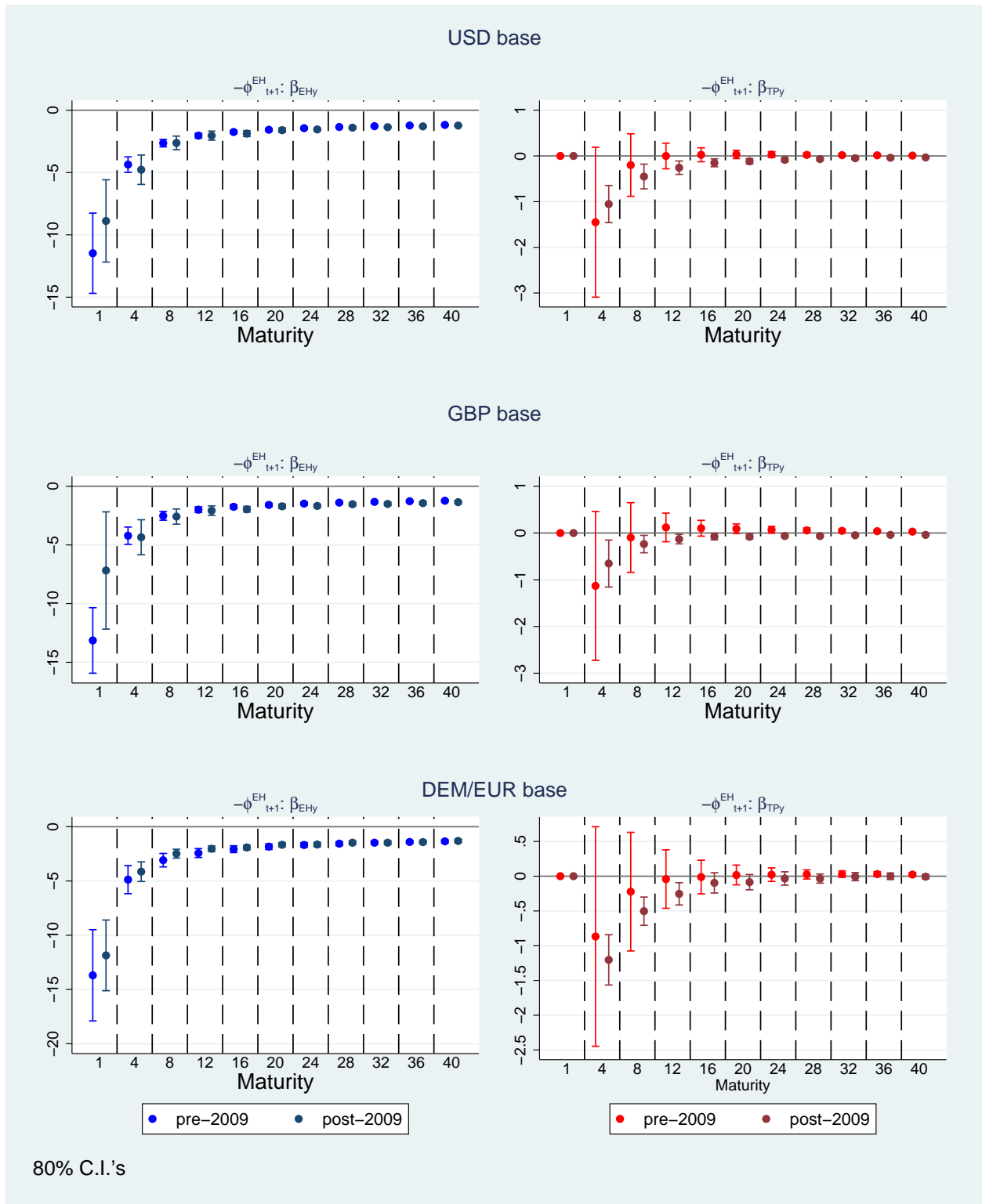
Sources: Authors' calculations.

Figure 12: Coefficients from Regressing $s_{t+1,\infty}^{\Delta E}$ on Relative 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



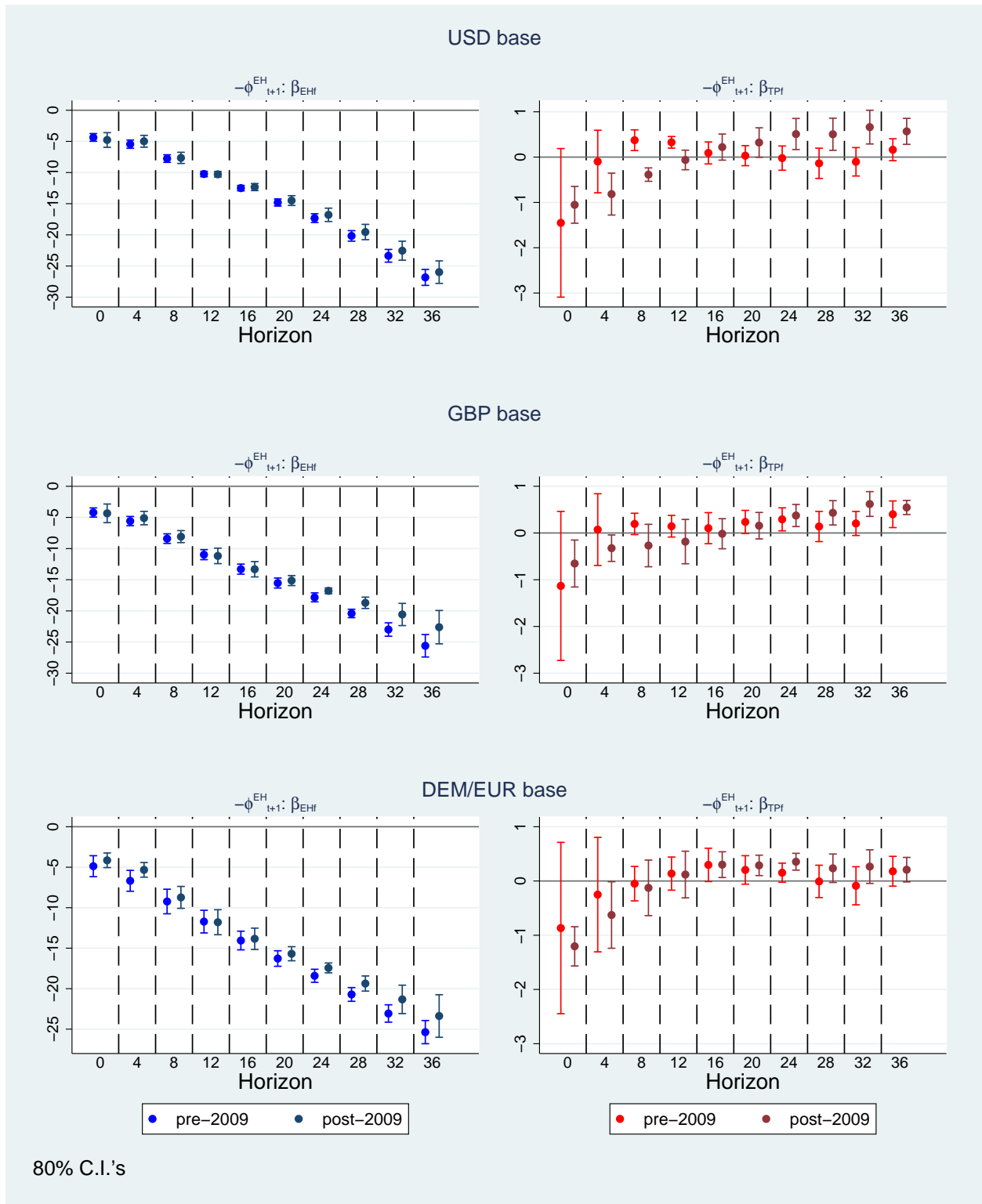
Sources: Authors' calculations.

Figure 13: Coefficients from Regressing $-\varphi_{t+1}^{EH}$ on Relative Expectations Hypothesis and Term Premia Components of Long-Term Yield Changes, 1990:Q1–2014:Q3



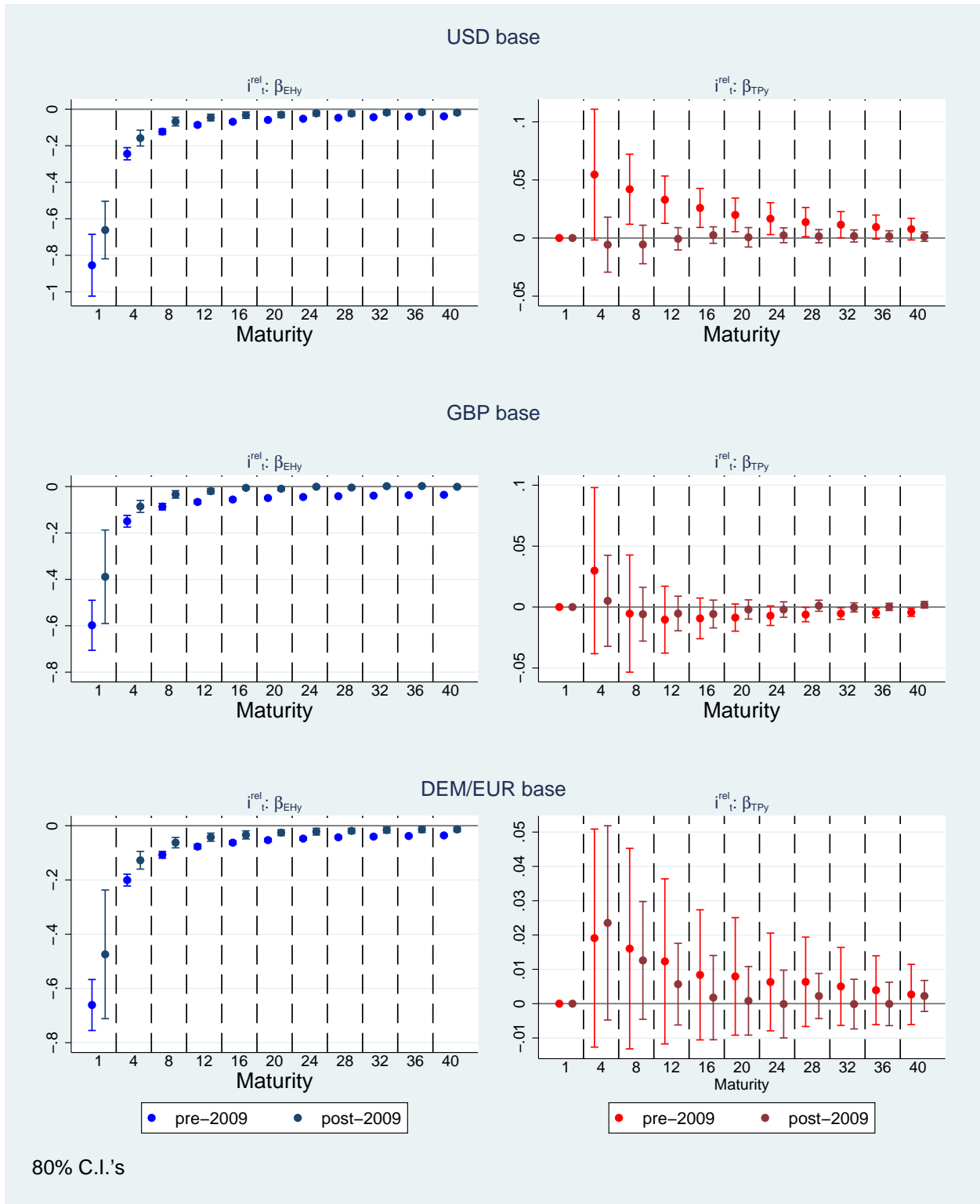
Sources: Authors' calculations.

Figure 14: Coefficients from Regressing $-\varphi_{t+1}^{EH}$ on Relative Expectations Hypothesis and Term Premia Components of 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



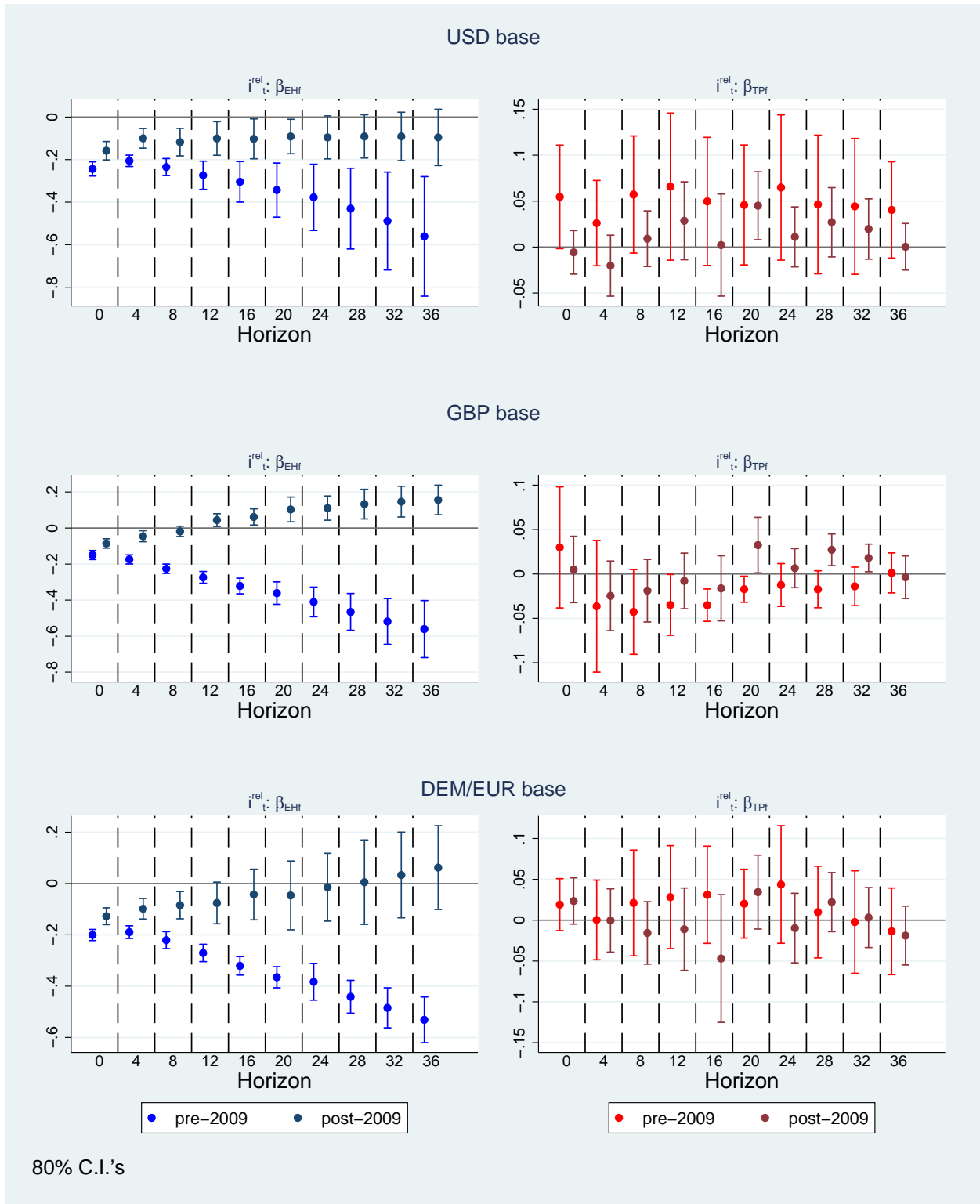
Sources: Authors' calculations.

Figure 15: Coefficients from Regressing \tilde{i}_t on Relative Expectations Hypothesis and Term Premia Components of Long-Term Yield Changes, 1990:Q1–2014:Q3



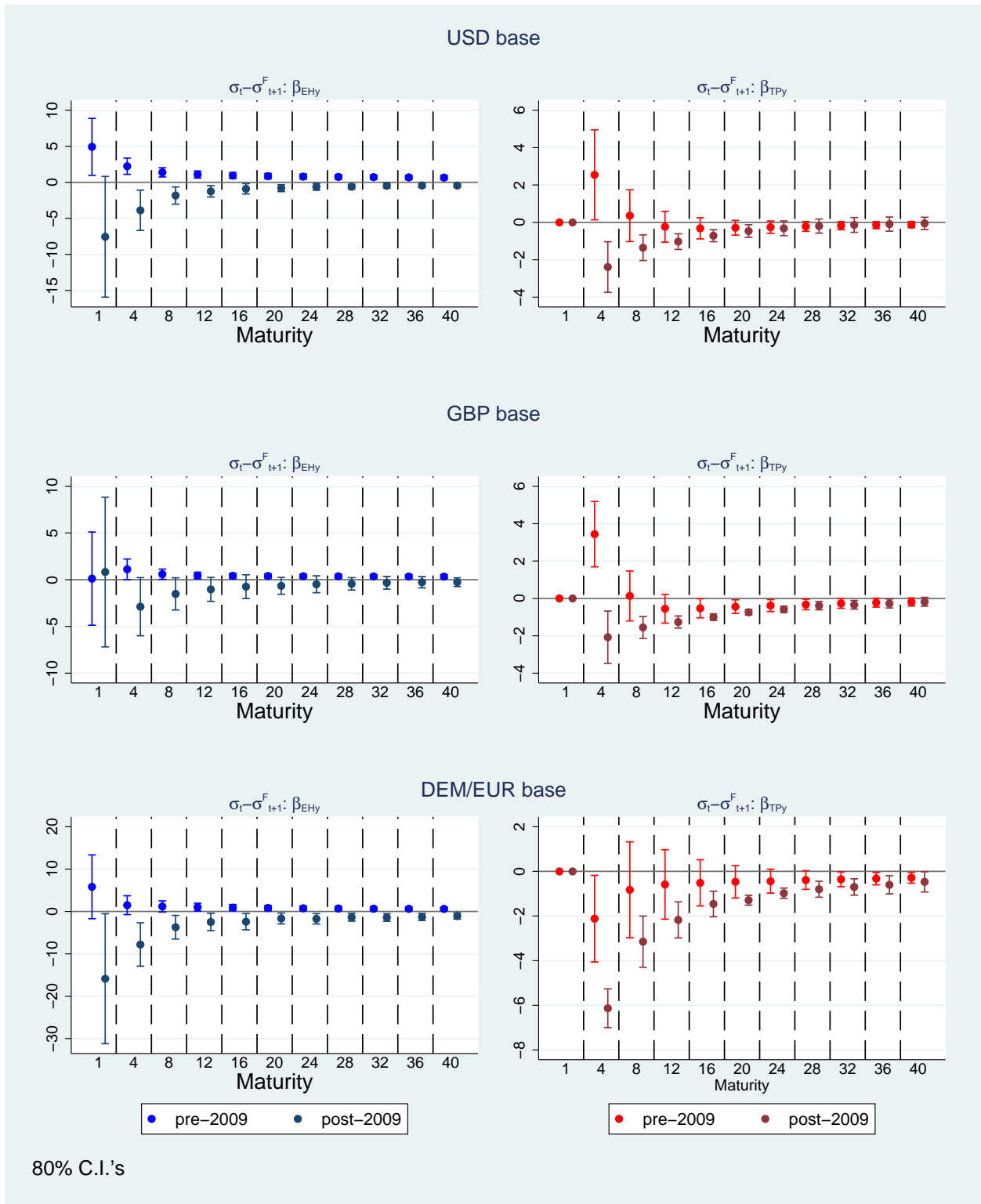
Sources: Authors' calculations.

Figure 16: Coefficients from Regressing \tilde{i}_t on Relative Expectations Hypothesis and Term Premia Components of 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



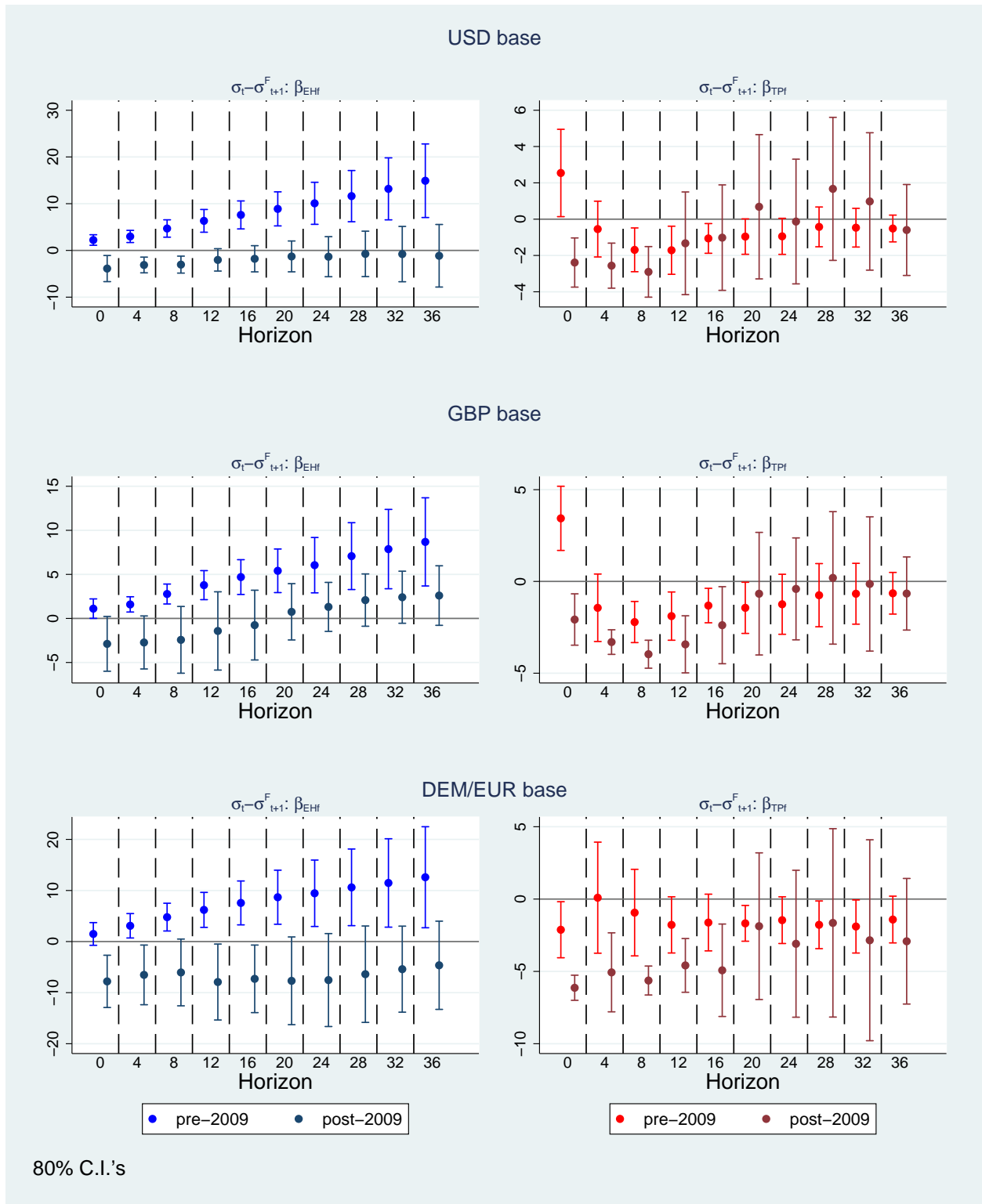
Sources: Authors' calculations.

Figure 17: Coefficients from Regressing $\sigma_t - \sigma_{t+1}^F$ on Relative Expectations Hypothesis and Term Premia Components of Long-Term Yield Changes, 1990:Q1–2014:Q3



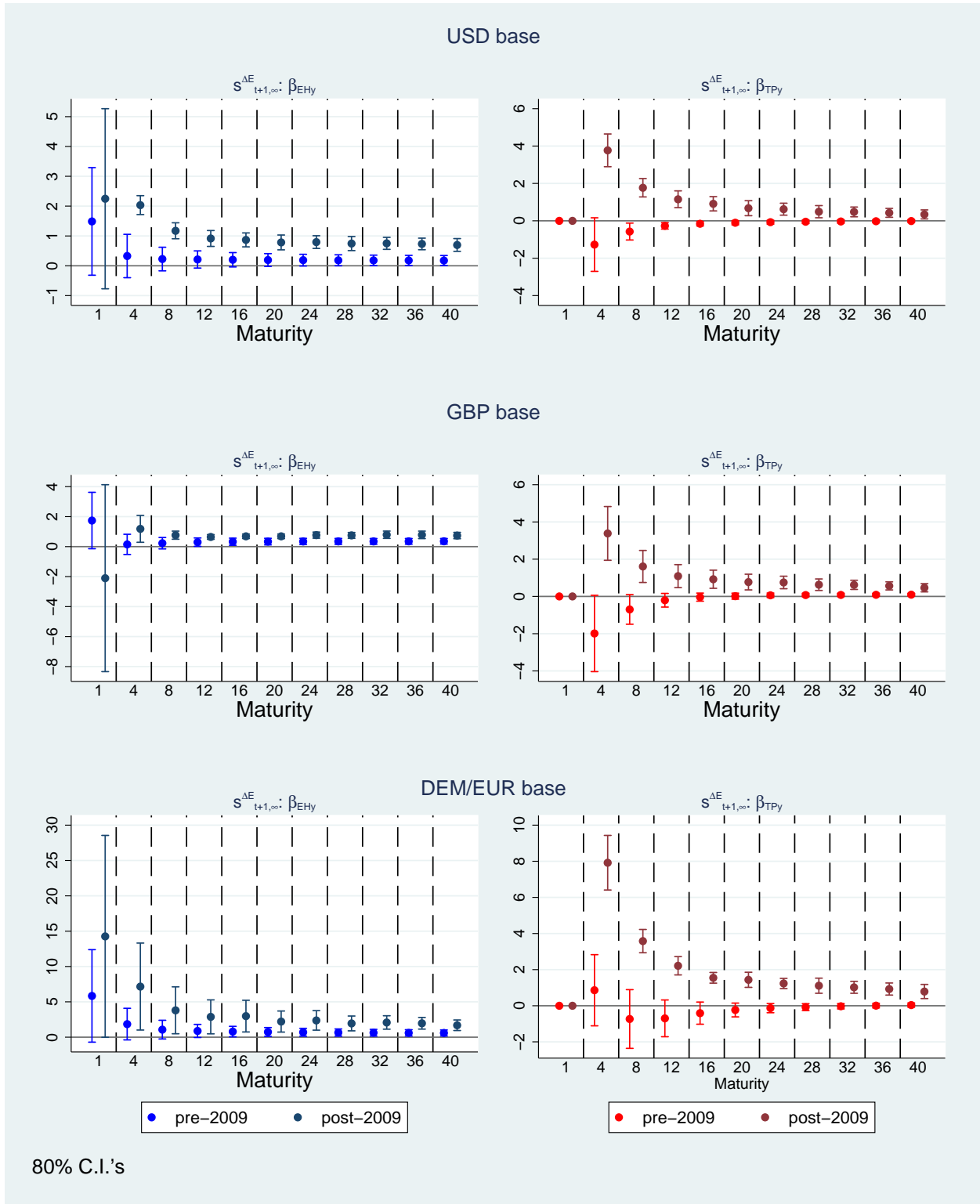
Sources: Authors' calculations.

Figure 18: Coefficients from Regressing $\sigma_t - \sigma_{t+1}^F$ on Relative Expectations Hypothesis and Term Premia Components of 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



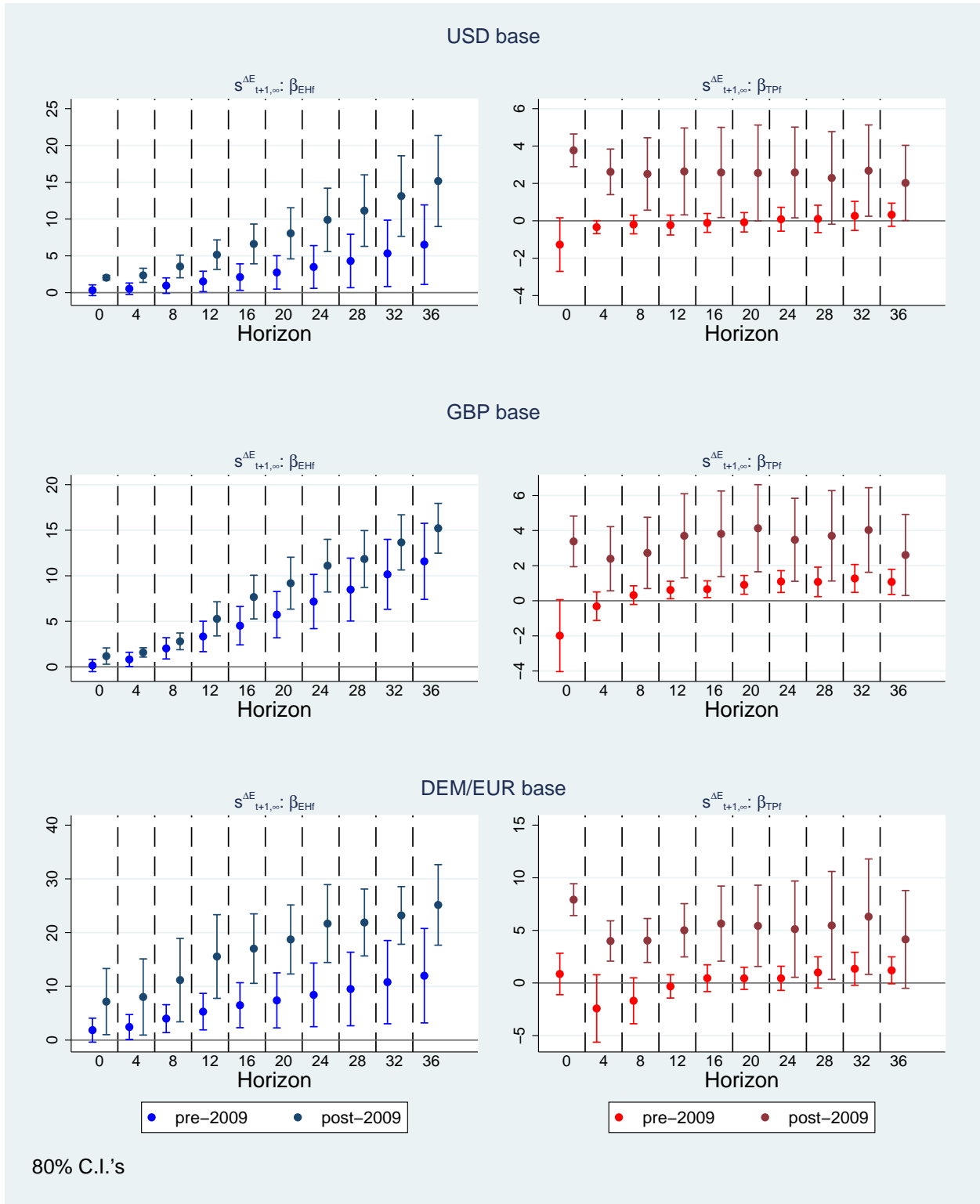
Sources: Authors' calculations.

Figure 19: Coefficients from Regressing $s_{t+1,\infty}^{\Delta E}$ on Relative Expectations Hypothesis and Term Premia Components of Long-Term Yield Changes, 1990:Q1–2014:Q3



Sources: Authors' calculations.

Figure 20: Coefficients from Regressing $s_{t+1,\infty}^{\Delta E}$ on Relative Expectations Hypothesis and Term Premia Components of 1-Year Forward Rate Changes, 1990:Q1–2014:Q3



Sources: Authors' calculations.

6 Appendix

6.1 Model

Consider a trader or financial institution that can invest in risk-free sovereign bonds of different maturities denominated in any currency. We will use no-arbitrage conditions seen from the trader's viewpoint to derive our expressions for exchange rate dynamics. As is common in the asset pricing literature, we start from no-arbitrage conditions. We assume that the trader's nominal stochastic discount factor (SDF), $M_{t+1}^{\$}$, is denominated in U.S. dollars and we do not take a stand on whether markets are complete or incomplete.¹⁵ Then, the no-arbitrage condition that guarantees the trader's indifference between investing in n -period bonds denominated in U.S. dollars, currency i , or currency j is

$$1 = E_t \left[M_{t+1}^{\$} R_{t+1}^{n,\$} \right] = E_t \left[M_{t+1}^{\$} R_{t+1}^{n,i} \frac{S_t^{i,\$}}{S_{t+1}^{i,\$}} \right] = E_t \left[M_{t+1}^{\$} R_{t+1}^{n,j} \frac{S_t^{j,\$}}{S_{t+1}^{j,\$}} \right] \text{ for every } i, j. \quad (22)$$

For analytical tractability, we assume that the nominal SDF, bond prices, and exchange rate changes are log-normally distributed.

First, we focus on one-period, risk-free bonds. Denoting logs of variables with lowercase letters, the first equality in equation (22) gives

$$i_t^i = -E_t (m_{t+1}^i) - \frac{1}{2} \text{Var}_t (m_{t+1}^i), \quad (23)$$

where

$$m_{t+1}^i = m_{t+1}^{\$} - \Delta s_{t+1}^{i,\$} \quad (24)$$

¹⁵This assumption that the trader's nominal SDF is denominated in U.S. dollars can be relaxed. We can assume that the unit of accounting is any currency or basket of currencies, which would be a realistic alternative if one interprets this trader as a global financial institution. More precisely, assume that the SDF of the global bank is denominated in a fictitious currency called ξ . All the derivations that follow will still hold, where the dollar-denominated SDF is related to the ξ -denominated SDF by the following accounting identity:

$$M_{t+1}^{\$} = M_{t+1}^{\xi} \frac{S_t^{\$, \xi}}{S_{t+1}^{\$, \xi}}.$$

is the trader's nominal SDF in terms of currency i and $\Delta s_{t+1}^{\$, \$} = 0$. In our model, equation (24) is simply an accounting identity.¹⁶

The second equality in equation (22) delivers the following expressions for the expected and realized exchange rate changes.

$$E_t \Delta s_{t+1}^{i,j} = i_t^i - i_t^j + \sigma_t^{i,j}$$

$$\text{where } \sigma_t^{i,j} \equiv \frac{1}{2} \text{Var}_t \left(\Delta s_{t+1}^{i,\$} \right) - \frac{1}{2} \text{Var}_t \left(\Delta s_{t+1}^{j,\$} \right) - \text{Cov}_t \left(m_{t+1}^{\$, \$}, \Delta s_{t+1}^{i,j} \right).$$

In terms of long-term bonds, the no-arbitrage condition for holding an n -period bond for k periods gives the following:

$$P_t^{n,i} = E_t \left[M_{t,t+k}^i P_{t+k}^{n-k,i} \right] \quad \text{where } M_{t,t+k}^i \equiv \prod_{l=1}^k M_{t+l}^i.$$

Assuming log-normality and using the definition $y_t^{n,i} \equiv -\frac{1}{n} \ln \left(P_t^{n,i} \right)$ gives

$$\begin{aligned} y_t^{n,i} &= -\frac{1}{n} p_t^{n,i} = -\frac{1}{n} \left[E_t m_{t,t+n}^i + \frac{1}{2} \text{Var}_t \left(m_{t,t+n}^i \right) \right] \\ &= -\frac{1}{n} E_t \sum_{k=1}^n \left(E_{t+k-1} m_{t+k}^i + \frac{1}{2} \text{Var}_{t+k-1} \left(m_{t+k}^i \right) \right) \\ &\quad + \frac{1}{n} \left[E_t \sum_{k=1}^n \frac{1}{2} \text{Var}_{t+k-1} \left(m_{t+k}^i \right) + \frac{1}{2} \text{Var}_t \left(m_{t,t+n}^i \right) \right] \\ &= \frac{1}{n} \sum_{k=0}^{n-1} E_t i_{t+k}^i + \underbrace{\frac{1}{2n} \left[\sum_{k=1}^n E_t \text{Var}_{t+k-1} \left(m_{t+k}^i \right) - \text{Var}_t \left(m_{t,t+n}^i \right) \right]}_{TP_t^{n,i}}, \end{aligned}$$

where we used the short-term bond pricing equation with $n = 1$,

$$i_{t+1}^i = -E_t m_{t+1}^i - \frac{1}{2} \text{Var}_t \left(m_{t+1}^i \right).$$

¹⁶Equation (24) resembles the relationship that would hold under an assumption of complete markets linking nominal exchange rate changes to the nominal SDFs of traders in two separate countries—the distinction being that these traders could have different real SDFs. The assumption of complete markets amounts to an assumption about the degree of risk-sharing between the traders in different countries, but it is not necessary for our derivations, since these are all made from the perspective of one single trader/institution.

6.2 Derivations for the AR(1) Example

Since we have

$$E_t \tilde{i}_{t+k}^j = (1 + \rho_j + \dots + \rho_j^k) \bar{i}_j + \rho_j^k \tilde{i}_t^j,$$

this gives

$$E_{t+1} \tilde{i}_{t+k+1} - E_t \tilde{i}_{t+k} = \rho_i^k \Delta \tilde{i}_{t+1}^i - \rho_j^k \Delta \tilde{i}_{t+1}^j \text{ for } k \geq 0.$$

Thus, we have

$$\begin{aligned} \Delta s_{t+1} &= - \left(E_{t+1} \sum_{k=0}^{\infty} \tilde{i}_{t+1+k} - E_t \sum_{k=0}^{\infty} \tilde{i}_{t+k} \right) + \sigma_t - \sigma_{t+1}^F + s_{t+1, \infty}^{\Delta E} \\ &= - \frac{1}{1 - \rho_i} \Delta \tilde{i}_{t+1}^i + \frac{1}{1 - \rho_j} \Delta \tilde{i}_{t+1}^j + \sigma_t - \sigma_{t+1}^F + s_{t+1, \infty}^{\Delta E}. \end{aligned}$$

Changes in relative n -period yields can be expressed as

$$\begin{aligned} \Delta \tilde{y}_{t+1}^{n,i} &= \frac{1}{n} \sum_{k=0}^{n-1} (E_{t+1} \tilde{i}_{t+k+1} - E_t \tilde{i}_{t+k}) + \Delta \widetilde{TP}_{t+1}^n \\ &= \frac{1}{n} \sum_{k=0}^{n-1} (\rho_i^k \Delta \tilde{i}_{t+1}^i - \rho_j^k \Delta \tilde{i}_{t+1}^j) + \Delta \widetilde{TP}_{t+1}^n \\ &= \frac{1}{n} \frac{1 - \rho_i^n}{1 - \rho_i} \Delta \tilde{i}_{t+1}^i - \frac{1}{n} \frac{1 - \rho_j^n}{1 - \rho_j} \Delta \tilde{i}_{t+1}^j + \Delta \widetilde{TP}_{t+1}^n. \end{aligned}$$

6.3 Data Description

- *Exchange Rates*: End-of-quarter exchange rates are obtained using daily data from Global Financial Data.
- *Short-term rates*: End-of-quarter three-month bill rates were obtained from the following sources:
 - Australia, Canada, New Zealand, Norway, Sweden, Switzerland, United Kingdom, and United States: Central bank data obtained through Haver Analytics.
 - Germany: Reuters data obtained through Haver Analytics. German three-month

bill rates are replaced with three-month EONIA OIS swap rates starting in 1999:Q1.

- Japan: Bloomberg
- *Zero-Coupon Yields*: End-of-quarter zero-coupon yields were obtained from the following sources:
 - Canada, Germany, Sweden, Switzerland, and United Kingdom: Central banks
 - Norway: Data from Wright (2011) extended with data from the BIS
 - Australia, New Zealand: Data from Wright (2011) extended with data from central banks
 - Japan: Wright (2011)
 - United States: Gürkaynak, Sack, and Wright (2007)
- *GDP Deflator, Output Gap, and Current Account-to-GDP ratio*: All macro data are from the OECD Main Economic Indicators and Economic Outlook databases. The GDP gap is computed using the OECD’s annual estimates of potential GDP, which were log-linearly interpolated to the quarterly frequency. German data are replaced with euro-area data starting in 1999:Q1.
- *Market-Based Interest Rate Surprises and Expected Changes*: These are computed using prices of futures on three-month interest rates on the last trading day of each quarter. These expectations refer to the three-month rates on each contract’s last trading day, which typically falls within the second-to-last week of each quarter. When computing the surprises and expected changes in these interest rates, the actual rate used is the underlying rate of each futures contract. The futures data are all obtained from Bloomberg and are based on the following underlying rates:
 - Australia: Australian 90-day bank accepted bills
 - Canada: Canadian three-month bankers’ acceptance
 - Switzerland: three-month Euroswiss

- Germany/EU: ICE three-month Euribor
- Norway: three-month NIBOR
- New Zealand: New Zealand 90-day bank accepted bills
- Sweden: three-month Swedish T-bill (1992:Q4–2007:Q4); three-month STIBOR (2008:Q1-present)
- United Kingdom: three-month Sterling LIBOR
- United States: three-month Eurodollar

Table 7: Data Sample Ranges

	Exchange rates, short-term yields, and macroeconomic variables	Zero-coupon long-term yields
Australia	1990:Q1 – 2014:Q3	1990:Q1 – 2014:Q3
Canada	1990:Q1 – 2014:Q3	1990:Q1 – 2014:Q3
Germany	1993:Q2 – 2014:Q3	1993:Q2 – 2014:Q3
Japan	1992:Q3 – 2014:Q3	1992:Q3 – 2009:Q2
New Zealand	1990:Q1 – 2015:Q1	1990:Q1 – 2014:Q3
Norway	1995:Q2 – 2014:Q3	1998:Q1 – 2014:Q3
Sweden	1990:Q1 – 2014:Q3	1992:Q4 – 2014:Q3
Switzerland	1992:Q1 – 2014:Q3	1992:Q1 – 2014:Q3
United Kingdom	1990:Q1 – 2014:Q3	1990:Q1 – 2014:Q3
United States	1990:Q1 – 2015:Q1	1990:Q1 – 2015:Q1

6.4 Additional Tables

Table 8: Monetary Policy Persistence, 1990:Q1–2014:Q3

	AUD	CAD	CHF	DEM/EUR	JPY	NOK	NZD	SEK	GBP	USD
i_t^i	0.876*** [0.023]	0.937*** [0.027]	0.876*** [0.022]	0.951*** [0.023]	0.873*** [0.035]	0.957*** [0.032]	0.915*** [0.033]	0.936*** [0.032]	0.954*** [0.015]	0.968*** [0.015]
constant	0.593*** [0.129]	0.136 [0.091]	0.089* [0.046]	0.052 [0.052]	0.021 [0.017]	0.107 [0.085]	0.436** [0.177]	0.105 [0.089]	0.083 [0.069]	0.022 [0.047]
# Observations	99	99	89	84	66	65	97	86	99	99
Adjusted R^2	0.944	0.944	0.940	0.936	0.946	0.901	0.905	0.941	0.974	0.958
RMSE	0.554	0.675	0.406	0.406	0.180	0.706	0.773	0.567	0.529	0.456

Frequency: Quarterly

Dependent variable: i_{t+1}^i three-month quarterly yields

Independent variable: Lagged three-month quarterly yields

Heteroskedasticity-robust standard errors are in brackets.

***/**/* Statistically significant at 1, 5, and 10 percent, respectively.