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Essays on Exchange Rates and Term Structures

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A dissertation
submitted in partial fulfillment of the
requirements for the degree of

Doctor of Philosophy

University of Washington

2018

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Program Authorized to Offer Degree:
Economics

University of Washington

Abstract

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The overall theme of this dissertation is the explanation of the relationship between the exchange rates and the term structures of assets. **Chapter 1**, "Theoretical Exposition of Exchange Rate and Term Structures,"develop a theoretical framework that links the term structures of assets to the exchange rate. After setting up the net present value (NPV) representation of the exchange rate in terms of stochastic discount factors (SDF) under no-arbitrage conditions, it describes how the current and expected future economic variables are incorporated into the exchange rate by referring to three major asset pricing approaches. Then, the term structures of two asset classes – sovereign credit default swap (CDS) and yield curves – are proposed to measure the market's expectations and perception of risk in driving the exchange rate dynamics.

Chapter 2, "Currency returns, Credit Risk and its Proximity: Evidence from Sovereign Credit Default Swap", examines whether credit risk and its proximity are priced in currency returns by making use of information in the term structure of sovereign CDS. Building upon and modifying a CDS pricing model, I construct two risk measures explaining different aspects of risk perception: (i) "risk level", measured by the level of the CDS curve, represents whether the expected loss given credit events is high or low, and (ii) "risk proximity", measured by the slope of the CDS curve, captures how soon a specific credit event is likely to be materialized. Combined with the NPV representation of exchange rate, I set up a model

where the exchange rate is determined by credit risk level and proximity. Using a broad data set between 2004 and 2017 for twenty countries, I show that risk level and proximity individually can explain a considerable amount of variation in currency returns and two risk measures together improve the predictive ability over a single CDS spread. Comparing the two, risk level broadly plays a stronger role during normal times, while risk proximity gains significance when financial crisis nears. These findings suggest that not only the credit risk level but also its proximity should be considered to assess the market's perception of risk driving currency movements.

Chapter 3[1], "Global Financial Crisis and the Exchange Rate – Yield Curve Connection", examines how the recent crisis and associated policy responses affect the relationship between market expectations, risk, and macro-fundamentals in driving exchange rate dynamics. To construct measures for expected macroeconomic conditions and perceived risk over future horizons, we decompose information in the term structure of interest rates across countries using several well-established yield-curve models. Data for eight major country pairs from 1995 to 2016 shows strong evidence that both expectations and risk premiums can explain subsequent exchange rate changes, with signs broadly consistent with theoretical predictions. We also observe clear structural changes, likely induced by unconventional monetary policy and the markets changing risk attitude since 2008. Specifically, while expectations play a consistent role over the full sample period, risk premiums pick up their significance mostly after the crisis. Taylor-rule macro-fundamentals at first provide little-to-no marginal explanatory power for currency movements over the yield-curve components, but do become important during the zero-lower-bound period. These findings suggest a joint macro-finance approach to modeling yield curve, macro fundamentals, and exchange rates, to better encapsulate changing market conditions and policy responses.

[1]This chapter is based on co-authored work with Yu-chin Chen and Kwok Ping Tsang

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ACKNOWLEDGMENTS

I would like to express my sincere gratitude to my dissertation committee chair, Professor Yu-chin Chen, for her constant guidance and support throughout my dissertation. Professor Chen provided invaluable insights and encouraged me to improve my research. This dissertation would never have been accomplished without her advice.

It is with great pleasure that I offer thanks to the rest of my committee members, Professor Chang-Jin Kim and Professor Ji Hyung Lee, for their helpful comments and suggestions for my study.

Last but not the least, I would also like to thank my family. My wife supported and trusted me at all times, and my son always made me happy. Completing this work would have been all the more difficult were it not for their love. My parents have given me unconditional love and have always stood by me throughout my life.

DEDICATION

To my wife and son

Chapter 1

THEORETICAL EXPOSITION OF EXCHANGE RATES AND TERM STRUCTURES

1.1 Introduction

Forward-looking nature of exchange rates has been documented by international finance literature following the "asset-market approach". According to a standard asset pricing theory, an asset price can be expressed as a discounted sum of current and expected future fundamentals (Campbell and Shiller, 1987; Cochrane, 2009). As quoted in Obstfeld and Rogoff (1996), "Nominal exchange rate must be viewed as an asset price." Then, like other assets, the current nominal spot exchange rate should theoretically reflect the market's expectations concerning present and future economic conditions (Frenkel and Mussa, 1985; Engel and West, 2005).

A major difficulty in examining the present value framework of exchange rates stems from the fact that there rarely exists a direct measure of the market's expectations. Several means of measuring expectations have been proposed. Earlier works make use of survey data to examine the response of exchange rates to announcements of economic news which induce economic agents to revise their expectations.[1] Although these studies show that a wide range of macroeconomic announcements leads to the currency movements as the theory predicts, there are fundamental limitations associated with the survey data. The surveys usually contain expectation data up to relatively short-term horizons, while the present value model requires expectations far into the future. Also, the observations are infrequent, untimely and missing for many countries. Another strand of research infers the expectations from

[1] Andersen et al. (2003), Faust et al. (2003), and Clarida and Waldman (2008) show that an announcement of news about economic expansion or higher-than-expected inflation leads to an appreciation of the currency. See Engel et al. (2007) for discussion.

the estimates of a VAR or other statistical models (Engel and West, 2006; Mark, 2009). However, the statistical estimates of expectations are not free from the measurement issues as economic agents use many other unobserved sources of information to form expectations.

My dissertation deviates from previous literature by directly using the readily observable term structures of assets, which incorporate the information about the market's expectations about the macro-fundamentals and perceptions of risk over time. Compared to the survey data and estimated expectations, a set of asset prices may incorporate more accurate and more up-to-date economic data. In this regard, this so-called "macro-finance approach" has proven fruitful in jointly modeling financial asset prices and economic variables. The yield curve literature repeatedly finds the close link between the shape of yield curve and expected future macroeconomic variables (Ang et al., 2006; Rudebusch and Wu, 2007). Extending to an open-economy model, Chen and Tsang (2013) show that cross-country yield curves are examined to explain the currency behavior well. As financial market deepening has proceeded, the attempts to extract the expectations and risk measures from other financial instruments with various maturities have increased. For example, Xu and Taylor (1994) find that the currency options have information about volatility expectation, and Clarida et al. (2003) show that the term structure of FX forward premiums succeeds in predicting future spot exchange rates. Recently, Augustin (2018) documents that the term structure of credit default swap (CDS) spreads can deliver the information about global and domestic risk factors. This chapter aims to demonstrate the theoretical frameworks how the term structures of financial assets can effectively explain the forward-looking nature of the exchange rate.

I first set up the net present value (NPV) representation of exchange rates in terms of the stochastic discount factors (SDF) in two countries under general no-arbitrage condition. In the following section, surveying three modeling strategies of finance literature, I link the current and future expected economic variables to the exchange rates. Final section demonstrates that the expected macro-fundamentals and time-varying risks can be measured by term structures of assets by taking examples of two asset classes – sovereign credit default swap (CDS) spreads and yield curves.

1.2 Net Present Value Framework of Exchange Rates

1.2.1 Risk-adjusted Uncovered Interest Rate Parity

In the absence of arbitrage, there is a stochastic discount factor (M_{t+1}) that prices returns in domestic currency (R_{t+1}) and equivalently a pricing kernel (M^*_{t+1}) that prices returns in foreign currency such that

$$\begin{aligned} 1 &= E_t(M_{t+1}R_{t+1}) \\ 1 &= E_t(M^*_{t+1}R^*_{t+1}) \end{aligned} \tag{1.1}$$

Under the assumption that each SDF follows a log-normal distribution, domestic and foreign interest rates of one-period default-free risk-free bonds (i_t and i^*_t) can be expressed as

$$\begin{aligned} i_t &= -E_t m_{t+1} - \frac{1}{2} Var_t m_{t+1} \\ i^*_t &= -E_t m^*_{t+1} - \frac{1}{2} Var_t m^*_{t+1} \end{aligned} \tag{1.2}$$

where $m_{t+1} \equiv log M_{t+1}$, $m^*_{t+1} \equiv log M^*_{t+1}$.

Define the nominal spot exchange rate at time t (S_t) as home currency per foreign currency. The domestic currency return of investing in a foreign asset is given by $R_{t+1} = (S_{t+1}/S_t)R^*_{t+1}$. As presented in Backus et al. (2001) and Brandt et al. (2006), this relations in combinations with eq.(1.1) under complete markets implies that

$$\frac{S_{t+1}}{S_t} = \frac{M^*_{t+1}}{M_{t+1}} \tag{1.3}$$

By taking logarithms for both sides of eq.(1.3) and then conditional expectations, I obtain the following exchange rate dynamic equation.

$$E_t \Delta s_{t+1} = E_t m^*_{t+1} - E_t m_{t+1} \tag{1.4}$$

where $s_t \equiv S_t$ and $\Delta s_{t+1} = s_{t+1} - s_t$. The expected depreciation of the exchange rate is given by the difference in the expectations of foreign and domestic pricing kernels in log. Combining eq.(1.4) with (1.2) gives us

$$E_t \Delta s_{t+1} = i_t - i_t^* + \frac{1}{2}\left(Var_t m_{t+1} - Var_t m_{t+1}^*\right) \tag{1.5}$$

The expected change of exchange rate consists of two parts, the interest rate differential and a term called the currency risk premium. By denoting the risk premium $\rho_t = -\frac{1}{2}\left(Var_t m_{t+1} - Var_t m_{t+1}^*\right)$, this equation is essentially same as the risk-adjusted uncovered interest rate parity (UIP) such as

$$E_t \Delta s_{t+1} = i_t - i_t^* - \rho_t \tag{1.6}$$

For better economic interpretation of the currency risk premium, I further assume that the exchange rate also follows the log-normal distribution and derive the equation using covariance terms.[2]

$$E_t \Delta s_{t+1} = i_t - i_t^* - \frac{1}{2}\left[Cov_t(m_{t+1}, \Delta s_{t+1}) - Cov_t(m_{t+1}^*, -\Delta s_{t+1})\right] \tag{1.7}$$

In this representation, $\rho_t = \frac{1}{2}\left[Cov_t(m_{t+1}, \Delta s_{t+1}) - Cov_t(m_{t+1}^*, -\Delta s_{t+1})\right]$. The risk premium is determined by the covariance between the currency value and the SDF in domestic and foreign country. One standard principle in asset pricing theory is that an asset is less desirable and risky when it does not pay well when money is really needed. In this case, investors command a premium to hold it. Suppose m_{t+1} and m_{t+1}^* are high due to a global economic shock. Considering that the SDF is often interpreted as the representative agent's inter-temporal marginal rate of substitution, agent's marginal utility and appetite

[2]Specifically, first I rearrange the eq.(1.3) to $M_{t+1}\frac{S_{t+1}}{S_t} = M_{t+1}^*$ and $M_{t+1}^*\frac{S_t}{S_{t+1}} = M_{t+1}$. Then taking logarithms for both sides and conditional expectations, applying the log-normality assumption gives $E_t\Delta s_{t+1} = -E_t m_{t+1} - \frac{1}{2}Var_t m_{t+1} + E_t m_{t+1}^* + \frac{1}{2}Var_t m_{t+1}^* - \frac{1}{2}\left[Cov_t(m_{t+1}, \Delta s_{t+1}) - Cov_t(m_{t+1}^*, -\Delta s_{t+1})\right]$. Finally, by plugging in eq.(1.2), I get the above equation.

for wealth are high in the future period. In this circumstance, if $Cov_t(m_{t+1}, \Delta s_{t+1}) > 0$, $Cov_t(m^*_{t+1}, -\Delta s_{t+1}) < 0$ or as long as if $Cov_t(m_{t+1}, \Delta s_{t+1}) - Cov_t(m^*_{t+1}, -\Delta s_{t+1}) > 0$, then the home currency relative to foreign currency is a bad hedge, and thus risk premium is required to compensate the investors for bearing risk. As a result, the home currency is expected to appreciate as the negative sign in front of risk premium implies.

1.2.2 Net Present Value Representations

Iterating forward the eq.(1.4), I get the NPV equation that the exchange rate is the sum of cross-country differences in current and expected future SDFs.

$$s_t = E_t \sum_{j=0}^{\infty} \left(m_{t+j+1} - m^*_{t+j+1} \right) + E_t s_{t+\infty} \tag{1.8}$$

The term $E_t s_{t+\infty}$ can be thought of as the long-run nominal exchange rate, or the long-run real exchange rate assuming the purchasing power parity holds. Although large literature attempts to study the equilibrium real exchange rate, this chapter focuses on the nominal exchange rate dynamics, regarding the long-run exchange rate as a constant.

I also iterate three alternative representations of the risk-adjusted UIP to get

$$\begin{aligned}
s_t &= -E_t \sum_{j=0}^{\infty} \left(i_{t+j} - i^*_{t+j} \right) + E_t \sum_{j=0}^{\infty} \rho_{t+j} + E_t s_{t+\infty} && (1.9) \\
&= -E_t \sum_{j=0}^{\infty} \left(i_{t+j} - i^*_{t+j} \right) - \frac{1}{2} \sum_{j=0}^{\infty} \left(Var_t m_{t+j+1} - Var_t m^*_{t+j+1} \right) + E_t s_{t+\infty} && (1.10) \\
&= -E_t \sum_{j=0}^{\infty} \left(i_{t+j} - i^*_{t+j} \right) \\
&\quad + \frac{1}{2} \sum_{j=0}^{\infty} \left[Cov_t(m_{t+j+1}, \Delta s_{t+j+1}) - Cov_t(m^*_{t+j+1}, -\Delta s_{t+j+1}) \right] + E_t s_{t+\infty} && (1.11)
\end{aligned}$$

Nominal exchange rate can be expressed as two summations, the sum of current and expected future short-term interest rates and the sum of expected risk premiums over time.

1.3 Linking Current and Expected Economic Conditions to Exchange Rates

The previous NPV equations hold under general no-arbitrage conditions, but do not provide explicit links to macroeconomic variables. In this section, I demonstrate how market's expectations about current and future economic variables can be incorporated into the exchange rate by surveying three well-known modeling strategies in the finance literature.

1.3.1 Consumption-based Asset Pricing

Suppose that there are two endowment economies. In each economy, a representative agent maximizes

$$E_t \sum_{j=0}^{\infty} \beta^j U(C_{t+j}) \tag{1.12}$$

where $U(.)$ is the utility function, and C_t is the consumption. The SDF, in other words, the inter-temporal marginal rate of substitution is defined as

$$M_{t+1} = \beta \frac{U'(C_{t+1})}{U'(C_t)} \tag{1.13}$$

The SDF can be further specified in terms of macro-fundamentals by taking particular utility functions.

Consider a standard constant relative risk aversion (CRRA) utility function. A representative agent maximizes

$$E_t \sum_{j=0}^{\infty} \beta^j \frac{C_{t+j}^{1-\gamma} - 1}{1-\gamma} \tag{1.14}$$

where γ is the coefficient of relative risk aversion. The log of SDF of the home country is expressed as

$$m_{t+1} = log\beta - \gamma \Delta c_{t+1} \tag{1.15}$$

where $c_t \equiv logC_t$. Assume $c_{t+1} = \mu_{t+1} + c_t + \epsilon_{t+1}$, where $\epsilon_{t+1} \sim N(0, \sigma_{t+1})$. μ_{t+1} is a conditional expectation of the consumption growth rate and σ_{t+1} is a conditional volatility measure of consumption growth. The log of SDF of foreign country is assumed to have the same representation with superscript $*$ on variables and parameters. For simplicity, further assume symmetric parameters in home and abroad, $\beta = \beta^*, \gamma = \gamma^*$. Applying the NPV equation of exchange rate, eq.(1.8) and (1.10),[3]

$$\begin{aligned} s_t &= -\gamma \sum_{j=0}^{\infty} (\mu_{t+j+1} - \mu^*_{t+j+1}) \\ &= -E_t \sum_{j=0}^{\infty} \left(i_{t+j} - i^*_{t+j} \right) - \frac{\gamma^2}{2} \sum_{j=0}^{\infty} (\sigma_{t+j+1} - \sigma^*_{t+j+1}) \end{aligned} \tag{1.16}$$

Nominal exchange rate is determined by cross-country difference in expected consumption growth rates, and/or difference in expected consumption volatility over time. But, power utility function is often empirically examined to fail in generating sizeable asset returns. For example, if $\mu_{t+j} \approx \mu^*_{t+j} \approx \mu, \forall j$ and $\sigma_{t+j} \approx \sigma^*_{t+j} \approx \sigma, \forall j$, then it is difficult to explain the variation in exchange rate.

Many attempts have been taken to solve the problem. One approach is to try different utility functions such as non-time-separable utilities.[4] Consider that a representative agent is characterized by external habit preferences as in Campbell and Cochrane (1999) and Verdelhan (2010). The external habit level corresponds to a subsistence level or social externality. Now, the agent maximizes

$$E_t \sum_{j=0}^{\infty} \beta^j \frac{(C_{t+j} - H_{t+j})^{1-\gamma} - 1}{1-\gamma} \tag{1.17}$$

where H_t is external habit level. Defining the surplus consumption ratio, which is the percentage gap between consumption and habit, $X_t \equiv \frac{C_t - H_t}{H_t}$, the log of SDF is given by

[3] The long-run real exchange rate is ignored as being regarded as a constant in this chapter.

[4] The other approaches include the general equilibrium models and factor pricing models. See Cochrane (2009) for discussion

$$m_{t+1} = ln\beta - \gamma(\Delta x_{t+1} + \Delta c_{t+1}) \tag{1.18}$$

where $x_t \equiv logX_t$. Assuming that

$$\begin{aligned} c_{t+1} &= \mu_{t+1} + c_t + \epsilon_{t+1}, \epsilon_{t+1} \sim N(0, \sigma_{t+1}) \\ x_{t+1} &= (1-\phi)\bar{x} + \phi x_t + \lambda(x_t)(\Delta c_{t+1} - \mu_{t+1}) \end{aligned}$$

where $\lambda(x_t) = \frac{\sqrt{1-2(x_t-\bar{x})}}{\bar{X}} - 1$ for $x_t \leq x_{max}$, 0 otherwise, $\bar{X} = \frac{\gamma\sigma}{\sqrt{\gamma(1-\phi)-\delta}}$, and $\gamma(1-\phi) > \delta$. I rewrite the SDF of the home country as

$$m_{t+1} = ln\beta - \gamma\left[\mu_{t+1} + (\phi-1)(x_t - \bar{x}) + (1+\lambda(x_t))(\Delta c_{t+1} - \mu_{t+1})\right] \tag{1.19}$$

Similarly, the log of SDF of the foreign country is defined with superscript $*$ on variables and parameters. Again, assume symmetric parameters across countries, $\beta = \beta^*, \gamma = \gamma^*, \bar{x} = \bar{x}^*$. Substituting the SDFs into the NPV equation, eq.(1.8) and (1.10),

$$\begin{aligned} s_t &= -\gamma \sum_{j=0}^{\infty}(\mu_{t+j+1} - \mu^*_{t+j+1}) + \gamma(1-\phi)\sum_{j=0}^{\infty}(E_t x_{t+j} - E_t x^*_{t+j}) \qquad (1.20) \\ &= -E_t \sum_{j=0}^{\infty}\left(i_{t+j} - i^*_{t+j}\right) \\ &\quad -\frac{\gamma^2}{2\bar{X}}\sum_{j=0}^{\infty}\left[(\sigma_{t+j+1} - \sigma^*_{t+j+1})(1+2\bar{x}) - 2(\sigma_{t+j+1}x_{t+j} - \sigma^*_{t+j+1}x^*_{t+j})\right] \end{aligned}$$

Nominal exchange rate is determined by cross-country difference in expectations and variance of consumption growth, surplus consumption ratio, and interaction between them. Even with $\mu_{t+j} \approx \mu^*_{t+j} \approx \mu, \forall j$ and $\sigma_{t+j} \approx \sigma^*_{t+j} \approx \sigma, \forall j$, this model can potentially produce the variation in the exchange rate as long as two countries have heterogeneity in habit formation.

Verdelhan (2010) shows that under the external habit preference and with the assumptions on parameter values, countercyclical risk premia and procyclical interest rates can account for the uncovered interest rate parity (UIP) puzzle. Epstein and Zin (1989) recursive preference has also been widely applied to the model as an alternative non-time-separable utilities. For example, Bansal and Shaliastovich (2012) demonstrate how models based on the Epstein-Zin preference and the preference for the early resolution of uncertainty can generate the stylized fact that high-interest rate currencies are expected to appreciate. However, none of these papers based on the habit formation and recursive preferences explicitly explores the forward-looking nature of exchange rates, which this chapter attempts to explain.

1.3.2 Factor-based Asset Pricing

Many asset pricing studies use flexible factor models under the no-arbitrage condition. The SDF that prices all assets, assuming the log-normal distribution, can be expressed as

$$M_{t+1} = exp(-i_t - \frac{1}{2}\lambda_t^{'}\lambda_t - \lambda_t\epsilon_{t+1}) \tag{1.21}$$

where λ_t is market prices of risk, which is a function of state variables. For example, the affine Gaussian dynamic term structure model pioneered by Ang and Piazzesi (2003) has the following specification: A vector of risk factors or state variables, X_t, follows a first-order Gaussian VAR.

$$X_t = \mu + \Phi X_{t-1} + \Sigma\epsilon_t$$

where $\epsilon_t \sim N(0, I)$ and Σ is lower triangluar. The short-term interest rate, i_t, and a vector of prices of risk, λ_t, are an affine function of the factors.

$$\begin{aligned} i_t &= \delta_0 + \delta_1^{'} X_t \\ \lambda_t &= \lambda_0 + \lambda_1^{'} X_t \end{aligned}$$

Making use of the flexibility associated with the choice of state variables, macro-fundamentals including inflation rate, real activity measures, output gap, policy rates have been employed in the literature (Ang and Piazzesi, 2003; Ang et al., 2006; Dewachter and Iania, 2011).

It is straightforward to show that the NPV representation of exchange rate can also be linked to the underlying economic variables by combining eq.(1.21) and its foreign counterpart with eq.(1.8).

$$\begin{aligned} s_t &= E_t \sum_{j=0}^{\infty} \left(m^*_{t+j+1} - m_{t+j+1} \right) \\ &= -E_t \sum_{j=0}^{\infty} \left(i_{t+j} - i^*_{t+j} \right) - \frac{1}{2} E_t \sum_{j=0}^{\infty} \left(\lambda'_{t+j} \lambda_{t+j} - \lambda^{*,'}_{t+j} \lambda^*_{t+j} \right) \\ &= f(X_t, E_t X_{t+1}, ..., E_t X_{t+\infty}, X^*_t, E_t X^*_{t+1}, ..., E_t X^*_{t+\infty}) \end{aligned} \tag{1.22}$$

Without further developing the expression, it is clear that the nominal exchange rate depends on a function of current and expected future values of state variables in two countries.

1.3.3 Taylor-rule Monetary Policy

Another body of finance literature links the macro-fundamentals to asset prices through Taylor-rule type monetary policies. Consider a standard two-country model where the domestic and foreign central banks sets their interest rates, i_t and i^*_t. I assume that the central bank follows a standard Taylor rule, addressing inflation and output (or unemployment) deviations from their target levels, but the domestic country targets the real exchange rate, or purchasing power parity, in addition. This captures the notion that central banks often raise interest rates when their currency depreciates, as discussed in Taylor (2001).[5] The monetary policy rules can be expressed as

$$\begin{aligned} i_t &= \mu_t + \beta_y \widetilde{y}_t + \beta_\pi (\pi_t - \bar{\pi}_t) + \delta q_t + u_t \\ i^*_t &= \mu^*_t + \beta^*_y \widetilde{y}^*_t + \beta^*_\pi (\pi^*_t - \bar{\pi}^*_t) + u^*_t \end{aligned} \tag{1.23}$$

[5]It is common in the literature to assume that the Fed reacts only to inflation and output gap, yet other central banks put a small weight on the real exchange rate. See Clarida, Gali, and Gertler (1999), Engel et al. (2007), and Molodtsova and Papell (2009).

where $\widetilde{y}_t$ is the output gap, π_t is the inflation rate, $\bar{\pi}_t$ is the inflation rate target and $q_t(= s_t - (p_t - p_t^*))$ is the real exchange rate, defined as the nominal exchange rate, s_t, adjusted by the CPI-price level difference between home and abroad, $p_t - p_t^*$. μ_t is the intercept including the equilibrium interest rate, and the stochastic shock u_t represents policy errors, which are assumed to be white noise. The corresponding foreign variables are denoted with superscript "$*$" and all variables except for the interest rates in these equations are in logged form. For simplicity, I assume the home and foreign central banks to have the same policy weights, and that $\beta_y = \beta_y^* > 0$, $\beta_\pi = \beta_\pi^* > 1$ and $\delta > 0$. Then, approximating the monetary policy rules, eq.(3.1), with $m = 1$, I specify the home relative to foreign monetary policy rules as

$$i_t^1 - i_t^{1,*} = \beta f_t^R + \delta s_t \tag{1.24}$$

where $f_t^R = [\mu_t - \mu_t^*,\ \widetilde{y}_t - \widetilde{y}_t^*,\ (\pi_t - \bar{\pi}_t) - (\pi_t^* - \bar{\pi}_t^*),\ -(p_t - p_t^*),\ u_t - u_t^*]'$ and $\beta = [1,\ \beta_y,\ \beta_\pi,\ \delta, 1]$.

Relating to the NPV equation, eq.(1.9),

$$\begin{aligned} s_t &= -E_t \sum_{j=0}^{\infty} (i_{t+j} - i_{t+j}^*) + E_t \sum_{j=0}^{\infty} \rho_{t+j} \\ &= -E_t \sum_{j=0}^{\infty} \left(\frac{1}{1+\delta}\right)^{j+1} \beta f_{t+j}^R + E_t \sum_{j=0}^{\infty} \left(\frac{1}{1+\delta}\right)^{j+1} \rho_{t+j}^1 \end{aligned} \tag{1.25}$$

Nominal exchange rate depends on expected macro-fundamentals and the risks over time.

1.4 Term Structure of Asset as Proxy for Expectations and Risk Determining Exchange Rates

The NPV representations of exchange rate evidently show that the nominal exchange rate should reflect the current and expected future economic fundamentals and/or risks. But, empirically it is not of trivial tasks to measure the market's expectations about future economic situations, especially over long horizons. This section demonstrates how the term structures of two asset classes can provide relevant information about either expectations and/or risks which drives the currency movements.

1.4.1 Sovereign Credit Default Swap

(1) Exchange Rates and Sovereign Credit Risk

Recall the NPV equation of exchange rate based on the factor-based asset pricing.

$$\begin{aligned} s_t &= E_t \sum_{j=0}^{\infty} \left[-(i_{t+j} - i^*_{t+j}) - \frac{1}{2} \left(\lambda'_{t+j}\lambda_{t+j} - \lambda^{*\prime}_{t+j}\lambda^*_{t+j} \right) \right] \\ &= f(X_t, E_t X_{t+1}, ..., E_t X_{t+\infty}) \end{aligned} \tag{1.26}$$

This NPV representation implies that the exchange rate at time t depends on the current and expected future path of state variables. Note that this equation is derived under the assumption of default-free assets. Now, relaxing the assumption to allow the possibility of default, I present the modified NPV equation of exchange rate.

Consider a price at time t of a one-period zero-coupon bond (P_t^1) which is subject to default risk (Duffie and Singleton, 1999) given by

$$P_t^1 = exp(-r_t - h_t l_t) \tag{1.27}$$

where r_t is risk-free short rate at time t, h_t is default probability between time t and $t+1$ conditional on information up to time t, and l_t is expected fractional loss at $t+1$ conditional on information up to time t. Compared to a default-free bond, a defaultable bond is like being priced using the default-adjusted discount rate, which is essentially the nominal short rate: $i_t = r_t + h_t l_t$.[6] Under the absence of arbitrage and the assumption of a pricing kernel of the home country (M_{t+1}) which prices all assets is conditionally log-normal,

$$M_{t+1} = exp\left(-r_t - h_t l_t - \frac{1}{2}\lambda'_t\lambda_t - \lambda'_t\epsilon_{t+1}\right) \tag{1.28}$$

Assuming that the sources of uncertainty are r_t, h_t, and l_t, I specify that λ_t is a function of these state variables. Equivalently defining the pricing kernel of the foreign country with

[6] See Duffie and Singleton (1999) for discussion.

corresponding variables with superscript $*$, and substituting two SDF equations into (1.26),

$$\begin{aligned} s_t &= E_t \sum_{j=0}^{\infty} \left[-(r_{t+j} + h_{t+j} l_{t+j} - r^*_{t+j} - h^*_{t+j} l^*_{t+j}) - \frac{1}{2} \left(\lambda'_{t+j} \lambda_{t+j} - \lambda^{*'}_{t+j} \lambda^*_{t+j} \right) \right] \\ &= f(R_t, H_t, L_t, R^*_t, H^*_t, L^*_t) \end{aligned} \quad (1.29)$$

where $R_t = (r_t, E_t r_{t+1}, ..., E_t r_{t+\infty})$, $H_t = (h_t, E_t h_{t+1}, ..., E_t h_{t+\infty})$, $L_t = (l_t, E_t l_{t+1}, ..., E_t l_{t+\infty})$, $R^*_t = (r^*_t, E_t r^*_{t+1}, ..., E_t r^*_{t+\infty})$, $H^*_t = (h^*_t, E_t h^*_{t+1}, ..., E_t h^*_{t+\infty})$ and $L^*_t = (l^*_t, E_t l^*_{t+1}, ..., E_t l^*_{t+\infty})$. This NPV representation shows that the exchange rate in the existence of potential default risk relies on the current and expected expected future path of risk-free short rates, default probabilities, and fractional losses given default.

(2) Price of Credit Default Swap[7]

In order to show that the informational contents contained in the term structure of CDSs are exactly the state variables associated with default risk, I start by evaluating a price at time t of an m-tenor CDS (D^m_t).[8] Building upon and modifying continuous time pricing of CDS in Duffie et al. (2003), I set up the pricing model in a discrete time framework to analytically examine the meaning of CDS spreads and their term structure. The premium leg, which is the present value of protection buyer's cash flow (V^{PB}_t), is given by[9]

$$V^{PB}_t = D^m_t + h_t exp(-r_t)0 + (1 - h_t) exp(-r_t) E_t(V^{PB}_{t+1}) \quad (1.30)$$

The CDS premium (D^m_t) is paid at the beginning of time t for protection. If a credit event is triggered between time t and $t + 1$, the contract is terminated and no more premium is

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[7] Please refer to **Chapter 2** and Appendix for details about CDS contract.

[8] I use the term "tenor" instead of "maturity" for CDS contract because the length of CDS contracts is shorter than or equal to underlying bond maturity and the cash flow of CDSs is fundamentally different from that of bonds.

[9] I assume without loss of generality that the CDS premium is paid on a yearly basis. This assumption enables us to relate annualized CDS spreads to annualized exchange rate changes and interest rates. However, the model can easily be translated into bi-annual or quarterly payment schemes.

paid. If not, the contract goes on to the next period. Iterating forward the eq.(1.30),[10]

$$V_t^{PB} = D_t^m E_t \left[\sum_{j=0}^{m} exp \left(- \sum_{k=0}^{j} (r_{t+k-1} + h_{t+k-1}) \right) \right] \tag{1.31}$$

On the other hand, the contingent payment leg, which is the present value of protection seller's cash flow (V_t^{PS}), is given by

$$V_t^{PS} = h_t exp(-r_t) l_t + (1 - h_t) exp(-r_t) E_t(V_{t+1}^{PS}) \tag{1.32}$$

If a credit event happens between time t and $t+1$, the protection seller pays the notional amount equivalent to the loss given event at the end of time $t+1$, terminating the contract. Iterating forward the eq.(1.32),

$$V_t^{PS} = E_t \left[\sum_{j=1}^{m} exp \left(- \sum_{k=1}^{j} (r_{t+k-1} + h_{t+k-2}) \right) h_{t+j-1} l_{t+j-1} \right] \tag{1.33}$$

Granted that a fairly priced CDS at time t equates V_t^{PB} and V_t^{PS}, a price at time t of an m-tenor CDS (D_t^m) is expressed as

$$D_t^m = \frac{E_t \left[\sum_{j=1}^{m} exp \left(- \sum_{k=1}^{j} (r_{t+k-1} + h_{t+k-2}) \right) h_{t+j-1} l_{t+j-1} \right]}{E_t \left[\sum_{j=0}^{m} exp \left(- \sum_{k=0}^{j} (r_{t+k-1} + h_{t+k-1}) \right) \right]} \tag{1.34}$$

A single CDS spread of tenor m is a function of risk-free short rates, default probabilities, and fractional losses given default over the same tenor. So, I rewrite the equation using the general functional form ($g(.)$) as

$$D_t^m = g(R_t^m, H_t^m, L_t^m) \tag{1.35}$$

$R_t^m = (r_t, E_t r_{t+1}, ..., E_t r_{t+m})$, $H_t^m = (h_t, E_t h_{t+1}, ..., E_t h_{t+m})$, $L_t^m = (l_t, E_t l_{t+1}, ..., E_t l_{t+m})$.

[10] I define $r_{t-1} = h_{t-1} = 0$ and use the approximation of $exp(c) \simeq 1 + c$, for small c.

(3) Linking Term Structure of CDS and Exchange Rates

Nominal exchange rate is determined by state vectors $R_t, H_t, L_t, R_t^*, H_t^*, L_t^*$ (eq.(1.29)). As each CDS spread of tenor m contains information about the same state variables up to its tenor (eq.(1.35)), the term structures of home and foreign CDS ($D_t = (D_t^1, D_t^2, ..., D_t^n), D_t^* = (D_t^{1,*}, D_t^{2,*}, ..., D_t^{n,*})$) can be used as proxies for these state vectors. Therefore, the eq.(1.29), using the different general function ($h(.)$), can be approximated by

$$\begin{aligned} s_t &= f(R_t, H_t, L_t, R_t^*, H_t^*, L_t^*) \\ &\approx h(D_t, D_t^*) \end{aligned} \tag{1.36}$$

The term structures of sovereign CDS spreads of home and foreign countries provide information about the current and expected future state variables associated with default risk, and thus can explain the forward-looking nature of exchange rate determination.

1.4.2 Yield Curves

(1) Exchange Rates and Taylor-rule Monetary Policy

Recall the NPV equation of exchange rate when central banks follow asymmetric Taylor-rules, where they set the policy rates reacting to inflation and output deviations from their targets with the home country adding the target of the real exchange rate.

$$s_t = \underbrace{-E_t \sum_{j=0}^{\infty} \left(\frac{1}{1+\delta} \right)^{j+1} \beta f_{t+j}^R}_{\text{Expectations}} + \underbrace{E_t \sum_{j=0}^{\infty} \left(\frac{1}{1+\delta} \right)^{j+1} \rho_{t+j}^1}_{\text{Risk}} \tag{1.37}$$

These formulations show that the exchange rate depends on both expected macro fundamentals (expectations) and the perceived risks over time (risk).

(2) The Yield Curve: Proxy for both Expectations and Risk

According to the NPV representation, the nominal exchange rate depends on the net present value of cross-country differences in expected future macroeconomic fundamentals and country-specific risk premiums over different horizons. Since exchange rates in this formulation rely more on the future expectations than on current variables, properly measuring expectations and time-varying risk becomes especially important in empirical testing.[11] I deviate from previous literature by making use of information embedded in the yield curves as proxies for both expectations about future macro fundamentals and the time-varying risks. I show that the Taylor-rule fundamentals are exactly the macroeconomic indicators the yield curves appear to embody information for, and currency risk factors are correlated with the bond term premiums. This can be viewed as an open-economy extension of the so-called joint "macro-finance" approach which has proven fruitful in modeling other financial assets such as the yield curves themselves. As stated in Diebold, Piazzesi, and Rudebusch (2005), the joint approach captures both the macroeconomic perspective that the short rate is a monetary policy instrument used to stabilize the economy, and the financial perspective that yields of all maturities are determined by the expected future path of short rates and the term premiums. The yield curves with short to long yields therefore contain market expectations about future macroeconomic conditions and perceived future uncertainty. Applying the same principles to the forward-looking exchange rate determination, cross-country yield differences of all maturities can proxy the expected future relative macro fundamentals and difference in the underlying risks across countries.

Traditional models of the yield curve postulate that the shape of the yield curve reflects the expected future paths of interest rates and future risk perception. According to the expectations hypothesis (EH), a long yield of maturity m can be written as the average of the current one-period yield and the expected one-period yields for the coming $m-1$ periods,

[11] Previous literature often ignores risk or makes overly simplistic assumptions about these expectations, such by using simple VAR forecasts of macro fundamentals as proxies for expectations (Mark, 1995; Engel and West, 2005). See discussion in Chen and Tsang (2013).

plus a term premium

$$i_t^m \equiv \frac{1}{m} E_t \sum_{j=0}^{m-1} i_{t+j}^1 + \theta_t^m \tag{1.38}$$

where θ_t^m represents the term premium perceived at t associated with holding a long bond until $t+m$. Extending this formulation to two-country model, the difference in the yields of maturity m across countries can be expressed as

$$i_t^m - i_t^{m,*} \equiv \frac{1}{m} E_t \sum_{j=0}^{m-1} (i_{t+j}^1 - i_{t+j}^{1,*}) + (\theta_t^m - \theta_t^{m,*}) \tag{1.39}$$

where I define the *relative yield* as the cross-country yields difference ($i_t^{m,R} \equiv i_t^m - i_t^{m,*}$), the *relative expected yields* as the average of the current and expected one-period yield differences across countries ($E_t i_t^{m,R} \equiv E_t(i_t^m - i_t^{m,*}) \equiv \frac{1}{m} E_t \sum_{j=0}^{m-1} (i_{t+j}^1 - i_{t+j}^{1,*})$), and the *relative term premiums* as the home minus foreign term premiums ($\theta_t^{m,R} \equiv \theta_t^m - \theta_t^{m,*}$).

The *relative expected yields* reflect market expectations about future macroeconomic fundamentals. A large body of research over the past decades has convincingly demonstrated that the yield curve contains information about expected future economic conditions such as output gap and inflation.[12] The recent macro-finance yield curve literature connects the observation that the short rate is a monetary policy instrument with the idea that yields of all maturities are risk-adjusted averages of expected short rates.[13] As the *relative expected yields* are constructed as the expected path of future short-term interest rate differentials, which in turn are set by monetary policy rules of home and foreign central banks, it is there-

[12] Two empirical strategies are typically adopted in the literature to test this macro-finance view of the yield curve. The first, more atheoretical approach captures the joint dynamics of the macroeconomic fundamentals and the yield curve using a general VAR. Ang and Piazzesi (2003) show that the yield curve predicts GDP growth and Diebold, Rudebusch, and Aruoba (2006) demonstrates strong dynamic interactions between the macroeconomy and the yield curves. Another body of studies model the macroeconomic variables structurally. Rudebusch and Wu (2007, 2008) find that the *level* factor incorporates long-term inflation expectations, and the *slope* factor captures the central bank's dual mandate of stabilizing the real economy and keeping inflation close to its target. Similarly, Dewachter and Lyrio (2006) find that the *level* factor reflects agents' long run inflation expectation, the *slope* factor captures the business cycle, and the *curvature* represents the monetary stance of the central bank.

[13] See Diebold, Piazzesi, and Rudebusch (2005) among others.

fore theoretically clear that the cross-country yield curves reflect market expectations about future macroeconomic fundamentals. By the equation, the *relative expected yields* proxy the first summation on the right hand side of eq.(1.37) since

$$E_t(i^1_{t+j} - i^{1,*}_{t+j}) = \beta E_t f^R_{t+j} + \delta E_t s_{t+j}, \forall j \tag{1.40}$$

The *relative term premiums* link to the currency risk premiums. Empirically, both the currency market and the bond market exhibit significant deviations from their respective risk-neutral efficient market conditions - the UIP and the EH - with the presence of time-varying risk being the leading explanation for both empirical patterns.[14] Assuming that a small number of underlying risk factors affect all asset prices, the bond term premiums would then be correlated with the currency risk premiums. This relationship can be shown with a simple modification of the UIP and EH equations.

$$\begin{aligned} E_t s_{t+m} &= i^m_t - i^{m,*}_t - \rho^m_t \\ &= \frac{1}{m} E_t \sum_{j=0}^{m-1} (i^1_{t+j} - i^{1,*}_{t+j}) - E_t \sum_{j=0}^{m-1} \rho^1_{t+j} \\ &= \frac{1}{m} E_t \sum_{j=0}^{m-1} (i^1_{t+j} - i^{1,*}_{t+j}) + (\theta^m_t - \theta^{m,*}_t) - \rho^m_t \end{aligned} \tag{1.41}$$

From this relationship, I show that the *relative term premiums* of maturity m is correlated with the currency risk premiums of the same maturity.

$$\rho^m_t - \frac{1}{m} \sum_{j=0}^{m-1} E_t \rho^1_{t+j} = \theta^m_t - \theta^{m,*}_t \tag{1.42}$$

[14] Fama (1984) and subsequent literature documented significant deviations from uncovered interest parity. In the bond markets, the failure of the expectation hypothesis is well-established. Wright (2011) and Rudebusch and Swanson (2012) are recent examples of research that studies.

Since the term premiums are required to compensate investors for holding the longer-maturity bond and these risks include systematic inflation, liquidity, and other risks over that maturity, the *relative term premiums* represent the cross-country difference in time-varying risks. Summing up, the *relative yield curves*, which are the sum of *relative expected yields* and *relative term premiums*, can proxy both expectations and risk portions in the NPV equation based on the Taylor-rule monetary policy.

1.5 Conclusions

This chapter provides the theoretical background for the forward-looking nature of exchange rate. Starting from very general net present value (NPV) representation of exchange rate under no-arbitrage conditions, I demonstrate how the nominal spot exchange rate reflects the current and expected future fundamentals by three major approaches in finance literature – consumption-based asset pricing, factor-based asset pricing, and Taylor-rule monetary policy. I argue that the fundamental limitations of the present value model of exchange rate about measuring market's expectations could be addressed by making use of direct measures from the term structures of assets. By taking examples of two asset classes – sovereign CDS spreads and yield curve – I show that the information embedded in the term structures are proxy for market's expectations about macro-fundamentals and perception of risk over time which drive the currency movements. In **Chapter 2** and **3**, I empirically examine these relationship with a broad data set.

Chapter 2

CURRENCY RETURNS, CREDIT RISK AND ITS PROXIMITY: EVIDENCE FROM SOVEREIGN CREDIT DEFAULT SWAP

2.1 Introduction

A large literature has documented the role of risk premiums in explaining currency returns. As currency can be viewed as a class of financial assets in international portfolios, systematic sources of risk drive currency returns both across currencies and over time. Since investors with risky currencies should be compensated for bearing risk, whether risk is high or low (the "risk level") forecasts returns to holding that currency. Another aspect of risk that investors are aware of is whether risk is near or far in time (the so-called "risk proximity").[1] If a specific risky event is likely to be realized anytime soon, withdrawal of investments and portfolio rebalancing cause changes in the value of a currency. Although both aspects of risk are perceived by investors and thus priced in currency returns, little attention has been paid to risk proximity. This chapter aims to evaluate the roles of these two different aspects of risk in explaining currency movements by using the information embedded in the term structure of sovereign credit default swaps.

A sovereign credit default swap (CDS) is a bilateral Over-the-Counter (OTC) insurance contract offering protection against the default of a referenced sovereign government. Protection buyer purchases insurance against contingent credit events by paying an annuity premium quarterly or bi-annually and protection seller compensates buyer for the losses given credit events.[2] Similar to other insurance contracts, CDS premiums or spreads naturally

[1] The term "risk proximity" is often used in risk management literature.

[2] In practice, there is no default in government bonds. Instead, the International Swap and Derivative Association (ISDA) references four types of credit events: acceleration, failure to pay, restructuring, and

provide information about the riskiness of a referenced entity, in this case, a sovereign.

Sovereign credit risk is closely related to currency risk. Intuitively, one of the most significant risks when holding assets issued in foreign currency is credit risk. Credit events such as a default may trigger a collapse of the banking system of a country, causing enormous losses on the value of assets denominated in that currency. Under floating exchange rate regime and free capital flow, apparently the price of insurance against credit events measures sovereign currency risk well. Many studies empirically find that sovereign CDS reflect the market pricing of time-varying systematic risks from various origins. Pan and Singleton (2008) and Longstaff et al. (2011) show that global risk is a main driver of CDS, while Remolona et al. (2008) argue that local risk is also an important determinant. The stability of the domestic financial system is found to affect CDS spreads (Acharya et al., 2014).

A sovereign CDS spread has an advantage over other risk measures because it is directly observable. As currency risk is not observable, there have been several alternative approaches to circumvent this issue. Pioneered by Lustig and Verdelhan (2007), much of the finance literature on the carry-trade strategy considers systematic exchange risk as an unobserved common factor.[3] They empirically show that portfolios constructed by the carry-trade strategy yield high returns due to latent risk measures. Another approach estimates currency risk by borrowing risk information from the other asset classes. Bekaert et al. (2007) point out that risk factors driving the premiums in the term structure of interest rates may also drive the risk premium in currency returns. Similarly, Chen et al. (2016) theoretically and empirically show that bond term premiums, which are separated out from the yield curve, are linked to currency risk premiums. Compared to these risk measures, a sovereign CDS spread, by nature, is free from issues related to nonobservance and estimation because it

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repudiation. Sovereign CDSs have been increasing in use from the early 2000s and as such, policymakers, regulators and investors monitor CDS spreads to gauge national financial stability. See the Appendix for details about contractual provisions and statistics.

[3]The carry-trade is a strategy under which investors take long positions on high-yield currencies and short positions on low-interest rate currencies. Other papers following this approach are Brunnermeier et al. (2008), Farhi et al. (2009), Lustig et al. (2011), Menkhoff et al. (2012) and many others.

is an observable price of credit risk determined by the interaction of protection buyers and sellers in the market. Recent researches in the aftermath of major global crises consider a sovereign CDS as a risk measure and relate it to carry-trade returns (Coudert and Mignon, 2013; Della Corte et al., 2016).[4]

Another useful feature of sovereign CDSs is their term structure. There exist different sovereign CDS contract tenors from 1- to 10-years, with each actively traded in the market, unlike corporate CDSs which are mostly concentrated in 5-year tenor contracts.[5] Borrowing the concept of the bond yield curve, I construct the term structure of sovereign CDSs and define the CDS curve as CDS spreads against tenors. I pay attention to the term structure because it delivers more useful information over a single asset price as proven in recent studies. These studies attempt to explain currency movements in the net present value (NPV) framework by using the information from the term structures. Clarida et al. (2003) show that the term structure of FX forward premiums succeeds in predicting future spot exchange rates, while a single forward rate fails. Similarly, Chen and Gwati (2014) find that FX option term structure contains information about higher moments in exchange rate dynamics and helps forecast currency returns. For the term structure of interest rates, which includes information about time-varying risk as well as the expectation of future macroeconomic fundamentals, cross-country yield curve differences are examined to better explain currency behavior (Chen and Tsang, 2013; Chen et al., 2016). While this line of research has shown that the measures extracted from the term structures improve the forecasting ability, studies on what exactly these measures represent and how they drive currency returns are scant. This chapter explicitly investigates the implications of risk measures from CDS term structure and links to exchange rates by building a CDS pricing model and combining with insights from previous studies on exchange rates and defaultable bonds modeling.

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[4]The earlier work of Reinhart (2002) also explores the effect of sovereign default events on currency crises, using credit rating data.

[5]See the Appendix for details about statistics.

I first construct two risk measures from CDS term structure. In the spirit of Nelson and Siegel (1987), the term structure of CDS spreads or the shape of the CDS curve is summarized into the level and slope factors. Graphically, the level captures the co-movement of various tenors of CDS spreads, and the slope reflects the gap between shortest- and longest-tenor CDS spreads. Interpreting the expressions for the level and slope of the CDS curve based on the CDS pricing model, I term them “risk level” and “risk proximity”, respectively. The level of the CDS curve is expressed as the weighted average of expected losses given default over different horizons. As an increase in the level implies that expected losses, regardless of contract tenors, grows due to either higher default probabilities or larger loss rates, I name it “risk level”. The slope, defined as shortest- minus longest-tenor CDS spreads, provides information about the timing of credit events. I prove that if default probability in the long-run is higher than in the short-run, the CDS curve is upward-sloping and the slope is negative. Conversely, the inverted CDS curve and positive slope result from relatively higher default probability in the near term. In this regard, a higher value of slope captures how soon the credit event is likely to be materialized – the so-called “risk proximity”.

Next, the net present value (NPV) representation of exchange rate presented in **Chapter 1** states that the exchange rate is determined by the cross-country differences in current and expected future path of risk-free short rates, default probabilities, and fractional losses given a default. Since the information is also contained in CDS spreads from short- to long-tenors and effectively summarized by the level and slope of the CDS curve, the connection between credit risk measures and currency returns can be well-established. Therefore, theoretically, currency returns are determined by both credit risk level and proximity.

For the empirical test, I look at a broad data set of quarterly exchange rate changes against the USD, three-month zero-coupon yields, and the term structures of sovereign CDSs for twenty countries from 2004 to 2017. Sample countries mostly consist of emerging and developing countries because credit risk is typically higher in these countries, compared to advanced countries. I first check if the level and slope factors capture most of the variation in entire sovereign CDS spreads of various tenors. Principal Component Analysis reveals that the first two principal components, which explain almost all of the variation in CDS curve, actually represent graphical level and slope, respectively. Then, I proceed to examine the roles of different sets of sovereign credit risk measures including a single CDS spread, level, and slope factors in explaining the exchange rate changes. The results confirm the existing findings in the recent finance literature as follows: 1) sovereign credit risk forecasts a large share of subsequent currency returns (Coudert and Mignon, 2013); 2) the effect of risk on currency movements is state-dependent (Clarida et al. (2009) among others) – an increase in sovereign credit risk results in positive currency returns in a low volatility state, while leading to low returns or losses in a high volatility state[6]; 3) the UIP puzzle is mitigated by reducing the omitted variable problem once incorporating credit risk measures in the regression.

The findings are distinctive in the following ways: 1) level and slope factors individually can explain a considerable amount of the variation in currency movements; 2) the model with both level and slope factors outperforms the model with a single CDS in explanatory power. These findings suggest that not only risk level but also risk proximity matters for currency returns and should not be ignored, results that are similar to, yet differentiated from the previous term structure literature. While confirming that the term structure of an asset delivers more useful information about risk in addition to a single asset price, I make progress by clearly interpreting that new information from the term structure is about different aspects of the risk perception of market participants – both risk level and proximity. To my best knowledge, there has been no attempt to investigate the role of risk proximity in driving asset prices. Considering that investors are sensitive to how soon a bad event is actually likely to happen as well as how much the expected loss is, the concept of risk proximity may be one of the missing pieces in asset pricing model in general; and 3) comparing the role of risk level and proximity, the former explains currency behavior more consistently, while the latter becomes more important as crises approach.

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[6]Clarida et al. (2009) divide the sample periods into high, medium, low volatility and find that high-interest rate currencies pay high returns during low volatility times, but low returns during high volatility times. These results coincide with the findings in Brunnermeier et al. (2008) that under adverse financial markets, as the carry-trades are unwound, dramatic depreciation happens for the high-interest rate currency.

2.2 Theoretical Framework

2.2.1 Credit Risk Level and Proximity

Recall a price time t of an m-tenor CDS (D_t^m) is expressed as

$$D_t^m = \frac{E_t\left[\sum_{j=1}^m exp\left(-\sum_{k=1}^j (r_{t+k-1} + h_{t+k-2})\right) h_{t+j-1} l_{t+j-1}\right]}{E_t\left[\sum_{j=0}^m exp\left(-\sum_{k=0}^j (r_{t+k-1} + h_{t+k-1})\right)\right]} \tag{2.1}$$

A single CDS spread of tenor m is a function of the current and expected future path of risk-free short rates, default probabilities, and fractional losses given default over the same tenor. To make use of information embedded in the term structure of CDSs in an efficient way, I describe the shape of the CDS curve with level and slope factors, borrowing the idea from the bond yield curve model (Nelson and Siegel (1987) among others).[7] Suppose a sovereign CDS has tenors of $m = 1, 2, ..., n$. The level factor is defined as a CDS spread of the longest tenor, $L(D_t) = D_t^n$, and the slope factor as the difference in CDS spreads between the shortest- and the longest-tenor, $S(D_t) = D_t^1 - D_t^n$. The meaning of the level factor is straightforward. From eq.(2.1), it represents the weighted average of expected losses given default over short to long horizons. The weights for each period are just different from conventional discount rates. Since expected loss measures whether the risk embedded in that asset is high or low, I refer to it as "risk level". An increase in level implies that expected losses, regardless of contract tenors, are escalated due to either higher default probabilities or larger loss rates.

Next, I examine the implication of the slope factor. To motivate the idea with a simple example, I assume there are only two CDS spreads, D_t^1 and D_t^2, determined by:

$$D_t^1 = \frac{exp(-r_t) h_t l_t}{exp(-r_t - h_t)} \tag{2.2}$$

$$D_t^2 = \frac{exp(-r_t) h_t l_t + E_t[exp(-r_t - r_{t+1} - h_t) h_{t+1} l_{t+1}]}{exp(-r_t - h_t) + E_t[exp(-r_t - h_t - r_{t+1} - h_{t+1})]} \tag{2.3}$$

[7] In **Section 2.3**, from Principal Component Analysis, I show that level and slope factors are actually the first two principal components which explain 99% of variation in entire CDS spreads at each time t.

If the slope is negative or the CDS curve is upward-sloping, then $D_t^1 < D_t^2$. Simplifying the eq.(2.2) $<$ (2.3) with the assumption of $l_t = l_{t+1}$, I obtain $h_t < E_t h_{t+1}$.[8] This implies that the default probability next period (between $t+1$ and $t+2$) is expected to be higher than this period (between t and $t+1$). Conversely, for the positive slope factor $D_t^1 > D_t^2$, or inverted CDS curve, default probability is relatively higher in the near-term, $h_t > E_t h_{t+1}$.

I derive the general expression which compares two adjacent CDS spreads, D_t^{m-1} and D_t^m for $m = 2, ..., n$. $D_t^{m-1} < D_t^m$ if and only if

$$h_t - E_t h_{t+m} + E_t \left[\sum_{j=1}^{m} exp \left(- \sum_{k=1}^{j} (r_{t+k} + h_{t+k-1}) \right) (h_{t+j} - h_{t+m}) \right] < 0 \qquad (2.4)$$

And $D_t^{m-1} > D_t^m$ if and only if

$$h_t - E_t h_{t+m} + E_t \left[\sum_{j=1}^{m} exp \left(- \sum_{k=1}^{j} (r_{t+k} + h_{t+k-1}) \right) (h_{t+j} - h_{t+m}) \right] > 0 \qquad (2.5)$$

If I assume monotonically increasing CDS spreads over tenors, $D_t^1 < D_t^2 < ... < D_t^n$, then I can prove that it must be $h_t < E_t h_{t+n}$.[9] I interpret that if the slope factor is negative, then the default probability in the long-run is higher than that in the short-run. Similarly, with the assumption of a monotonically decreasing CDS curve, $D_t^1 > D_t^2 > ... > D_t^n$, the positive slope reflects that default is more likely to happen in the short-run compared to the long-run, $h_t > E_t h_{t+n}$. In this regard, the higher value of the slope measures how near in time a specific credit event is likely to be realized, which I term "risk proximity".[10]

[8] In the credit default swap literature, fractional loss is conventionally assumed to be constant over time because it is determined by fundamentals of a referenced entity and thus estimated by the historical recovery rate ($r = 1 - l$) in practice.

[9] Starting from $D_t^1 < D_t^2$ and using the obtained condition $h_t < E_t h_{t+1}$, I obtain the condition for $D_t^2 < D_t^3$ to be $h_t < E_t h_{t+2}$. Continuing this process sequentially to $D_t^{n-1} < D_t^n$, I end up with $h_t < E_t h_{t+n}$.

[10] The term "risk proximity" is often used in the risk management literature.

2.2.2 Exchange Rates and Credit Risk Factors

As shown in **Chapter 1**, the nominal exchange rate can be expressed by state vectors over the short- to long-term, $(R_t, H_t, L_t, R_t^*, H_t^*, L_t^*)$, and the vectors of home CDS and foreign CDS over entire tenors ($D_t = [D_t^1, D_t^2, ..., D_t^n]'$, $D_t^* = [D_t^{1,*}, D_t^{2,*}, ..., D_t^{n,*}]'$) contains the same information up to reasonable long-term. Then, the spot exchange rate can be explained by home and foreign CDS term structures:

$$s_t = h(D_t, D_t^*) \tag{2.6}$$

Since the level and slope measures can effectively summarize the entire term structure of CDS spreads and provide a more interesting interpretation of risk level and proximity, I propose to use the level and slope instead of the entire set of CDS spreads. Then, the eq.(2.6), using the different general function ($F(.)$), becomes

$$s_t = F(L(D_t), S(D_t), L(D_t^*), S(D_t^*)) \tag{2.7}$$

Based on this theoretical framework, I conduct empirical studies to show how the level and slope factors of sovereign CDS curves explain the variation in the exchange rate in the following sections.

2.3 Data and Principle Component Analysis

2.3.1 Data Description

Twenty sample countries are chosen on the basis of two criteria. First, sovereign CDSs should be actively traded in the market. I select candidate countries in descending order by the trading volume reported by the Depository Trust and Clearing Corporation (DTCC).[11]

[11] The Depository Trust and Clearing Corporation(DTCC) runs a warehouse for CDS trade confirmations accounting for around 90% of the total market and releases market data on the outstanding notional of CDS trades on a weekly basis.

Second, among the candidates, I exclude the countries whose exchange rate regime is not floating, based on the IMF Annual Report on Exchange Arrangements and Exchange Restrictions, 2016. Sample countries are Australia (AU), Brazil (BR), Chile (CL), Colombia (CO), Hungary (HU), Iceland (IS), Indonesia (ID), Israel (IL), Japan (JP), Korea (KR), Mexico (MX), Norway (NO), Peru (PE), Philippines (PH), Poland (PL), Romania (RO), South Africa (ZA), Sweden (SE), Thailand (TH), Turkey (TR) and the United States of America (US). Although samples are mostly emerging and developing countries where sovereign credit risk is especially potential, I have a complete set of both advanced and emerging economies over various continents including America, Asia, Europe, and Africa.

The main data I examine consists of monthly observations from January 2004 to June 2017 of the following series:[12] 1) spot exchange rate data: End-of-month exchange rates are obtained from Bloomberg. I use the logged exchange rate, measured as a per-dollar rate. The quarterly exchange rate change is expressed as $\Delta s_{t+3} = s_{t+3} - s_t$ and indicated as an annualized percentage. A positive Δs_{t+3} denotes depreciation of the home currency against the US dollar, and a negative Δs_{t+3} denotes appreciation; 2) zero coupon yield of three-month maturity: End-of-month zero coupon yields as an annualized percentage are obtained from Bloomberg;[13] and 3) sovereign CDS data: Data on sovereign CDS spreads is collected from Bloomberg and Datastream. I use sovereign CDS spreads with tenors of 1, 2, 3, 5, 7, 10 years, with USD as the currency of denomination and in annuity basis point. CDS spreads are from the last trading day of each month.

Table 2.1 reports the summary statistics of sample data. Considering potential structural breaks due to the Great Recession, the sample period is divided by two preliminary break dates – November 2007 and June 2009.[14] For interpretation purposes, I label three sub-periods divided by two breaks as pre-crisis, crisis, and post-crisis. For the quarterly

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[12]Sample periods are shorter for some countries due to data availability.

[13]For countries with no zero coupon yield data, three-month interbank rates are obtained instead.

[14]According to NBER, the recession in the US was from December 2007 to June 2009. Although the recession periods in sample countries are not identical, I choose these breaks since no country was free from the massive impact of the Global Financial Crisis.

exchange rate change Δs_{t+3} in **Panel A**, I observe that all currencies appreciated before the crisis except for IDR, JPY and ZAR, and that all currencies except JPY depreciated during the crisis. This would be consistent with the idea that the US dollar (along with the Japanese Yen) is commonly considered as a safe haven currency. Behavior after the crisis differs across countries due to local and domestic events. For example, BRL, COP, HUF, and ZAR depreciated further, while ISK, ILS, KRW appreciated. The volatility of exchange rates increased during the crisis. After the crisis, the standard deviation has decreased but still is higher than pre-crisis levels for some countries, reflecting persisting uncertainties. From **Figure 2.1**, I see episodes of exchange rate volatility, with spikes during the Great Financial Crisis and the European Fiscal Crisis.

Panel B describes statistics on interest rate differentials measured by cross-country differences in zero-coupon yield of three-month maturity as home minus US yield, $i_t^3 - i_t^{3,*}$. The interest rate of the home country is higher than that of the US and the gap widened during the crisis. The volatility is very low, compared to that of exchange rates, implying that the interest rate differential is not enough to generate the variation in currency movements and lending support to the view that more volatile variables are needed as explanatory variables in addition.

2.3.2 Principal Component Analysis of Term Structure of CDS

I describe the evolution of the term structure of sovereign CDSs over the sample period. **Figure 2.2** graphically shows sovereign CDS spreads of six different contract tenors. One immediately noticeable feature present in all countries is that all the CDS spreads co-move. During major global events when sovereign credit risk is mounted, CDS spreads increase sharply regardless of tenors. What is more interesting in this figure is that the gap between short- and long-tenor CDS spreads varies over time. Usually, the long-tenor CDS spread is higher than the short-tenor CDS spread due to longer exposure and higher uncertainty, but the gap becomes narrower or even inverted during the Global Financial Crisis. The level of the CDS curve explains the co-movement of CDS spreads, while the slope describes the difference between the shortest- and the longest-tenor CDS spreads. From this observation, I draw two lessons: 1) the term structure of sovereign CDSs contains more information than a single CDS spread; and 2) the level and slope of the CDS curve summarize the shape of the CDS curve well. In the previous section, I theoretically demonstrated that the level reflects "risk level "and that the slope captures "risk proximity". For example, as Iceland experienced the financial crisis that involved the actual default of three major commercial banks between 2008 and 2010, instant sovereign default was highly anticipated in the market, which was reflected in a positive slope as well as a very high level.

Next, I perform Principal Component Analysis and analyze that the level and slope are indeed important two factors that determine the term structure of sovereign CDSs. Let D_t denote the $N \times 1$ vector of sovereign CDS spreads at each time t, and Ω be the $N \times N$ covariance matrix of D_t. The principal components are the linear combinations of D_t which account for as much variation in D_t as possible. That is,

$$\begin{aligned} PC_{1t} &= p_1^{'} D_t, \; where \; p_1 : eigenvector \; with \; the \; largest \; eigenvalue \; from \; \Omega \\ PC_{2t} &= p_2^{'} D_t, \; where \; p_2 : eigenvector \; with \; the \; second \; largest \; eigenvalue \; from \; \Omega \end{aligned}$$

The results are reported in **Table 2.2**. The first two principal components explain around 99% of the variation in entire sovereign CDS spreads, of which the first constitutes 84 - 98%, while the second accounts for 2 - 15%. Due to strong co-movement of six CDS spreads, the large proportion explained by the first principal component is not surprising. Rather, the second principal component explains a small but meaningful share of variation in the data. Turning to factor loadings on CDS spreads, the first principal component has roughly constant factor loadings across tenors, indicating that it is "level". On the other hand, the second principal component has positive factor loadings on short tenors and negative factor loadings on long tenors, implying that it is "slope", defined as short minus long tenor CDSs. These findings suggest that the two principal components coincide with the geometrical level and slope of the CDS curve as defined in **Section 2.2**. The correlation between the first principal component and level is over 0.9 for all the countries and the correlation between the second principal component and slope is mostly 0.8 or higher for most countries. In the following empirical studies, I use the two principal components instead of geometrical level and slope and denote them as "level" factor, $L(D_t)$, and "slope" factor, $S(D_t)$, respectively.[15]

Table 2.3 describes the statistics for level and slope factors with potential structural breaks. During the Financial Crisis, the level greatly increases while the slope became flatter or even inverted. This implies that investors perceived that huge losses from the credit event were highly likely to occur in the near future due to the impact from the Global Financial Crisis. In contrast, the slope during the post-crisis period has quickly become as steep as in the pre-crisis period, while the level remained high. It can be interpreted that sovereign risk level is still high due to the sluggish recovery of real economies and ongoing financial uncertainties, but actual credit events are not anticipated shortly thanks to international efforts in securing financial safety nets.

2.4 Sovereign Credit Risk and its Proximity in Exchange Rate Determination

2.4.1 Explaining Currency Returns with Sovereign Credit Risk

I empirically examine whether sovereign credit risk factors perceived at a particular point in time can explain quarterly currency returns as predicted from the reduced-form model developed in **Section 2.2**. I also compare the explanatory power of different sets of risk measures: 1) one-year CDS spread, D_t^1; 2) level factor, $L(D_t)$, which captures risk level; 3) slope factor, $S(D_t)$, which implies risk proximity; and 4) both level and slope factors, $L(D_t)$ and $S(D_t)$.[16] Specifically, the following regressions are estimated with structural breaks for

[15] Since $PC_{1t} \approx D_t^{10}$ and $PC_{2t} \approx D_t^1 - D_t^{10}$, whether I use two principal components or literal level and slope does not make a big difference to empirical studies.

[16] In empirical studies, I focus on home sovereign CDS spreads only. One reason is that I assume no default risk in the US, the foreign country. This assumption seems realistic in that market participants rarely expect a default in US assets. The other reason is that data on US sovereign CDS spreads only becomes available in recent years. However, in the Appendix, I conduct robustness checks by using cross-country differences in credit risk factors.

each currency pair:[17]

$$\text{Model 1: } \Delta s_{t+3} = \beta_0 + \beta_1 D_t^1 + \epsilon_{t+3} \quad (2.8)$$

$$\text{Model 2: } \Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \epsilon_{t+3} \quad (2.9)$$

$$\text{Model 3: } \Delta s_{t+3} = \beta_0 + \beta_1 S(D_t) + \epsilon_{t+3} \quad (2.10)$$

$$\text{Model 4: } \Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \beta_2 S(D_t) + \epsilon_{t+3} \quad (2.11)$$

Table 2.4 presents p-values from a joint Wald test and adjusted R^2s.[18] As shown in the table, p-values are all below 5% with a few exceptions for Israel and Thailand. The hypothesis that sovereign credit risk factors have no information about three-month exchange rate change is strongly rejected, regardless of which set of risk measures is employed. The regressions generate high adjusted R^2s up to 60%. This is quite an impressive portion in light of the near-zero R^2 typical in this literature. I confirm the existing finding that sovereign credit risk accounts for a large share of currency returns (Coudert and Mignon, 2013).

Comparing the predictive ability of four different models, I first notice that the model with both level and slope factors (Model 4) can explain currency movements better than the model with one-year CDS spread (Model 1) in most countries. This finding lines up with previous literature that the term structures provide more useful information than a single asset price (Clarida et al., 2003; Chen and Tsang, 2013). One-year CDS spread may capture the amount of credit risk in the shortest run, but can say nothing about the timing of risky events. In contrast, the level and slope from the term structure of CDSs, which capture both credit risk level and proximity, can together help more comprehensively evaluate the investors' perception of risk and thereby better forecast currency returns. Another prominent

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[17] The test for endogenous structural breaks in the regression is performed based on Bai and Perron (2003) multiple break tests(with 15% trimming and 5 - 10% significance level). After identifying zero to two breaks, structural break dummy variables for each sub-period are incorporated into the regression.

[18] To complement the analysis of the link between sovereign credit risk and currency movements, I also regress excess currency returns, $xr_{t+3}(= i_t^3 - i_t^{3,*} - \Delta s_{t+3})$ on four sets of risk measures. The results remain qualitatively the same. The results can be provided upon request.

feature is the performance of the slope-only model (Model 3). Given that the slope factor explains an only small portion of the variation in entire CDS spreads as investigated in the Principal Component Analysis, its explanatory power is beyond expectation. This model performs better than the one-year CDS model (Model 1) in 9 out of 20 countries and even better than the level-only model (Model 2) in 8 out of 20 countries. These results lend support to my argument that foreign currency holders care not only how much sovereign risk is expected from that currency, but also how soon the risky event is anticipated.[19]

I further investigate the relationship between credit risk measures and currency returns by employing a two-state Markov-Switching model. Although the structural break model estimates coefficients in sub-sample periods divided by break dates, the state-dependent relationship between variables might be averaged out in long sub-sample periods. This model is also silent about what each state represents. Alternatively, the Markov-Switching model can show state-dependency more clearly by allowing the regime to switch endogenously by unobserved yet comprehensive state factors. The model specification is as follows.

$$\Delta s_{t+3} = \beta_{0,\xi_t} + \beta_{1,\xi_t} L(D_t) + \beta_{2,\xi_t} S(D_t) + \epsilon_{t+3}, \quad (2.12)$$
$$where\ \epsilon_{t+3}\ i.i.d.N(0, \sigma^2_{\xi_t})$$

$$\beta_{i,\xi_t} = \beta_{i,0}(1-\xi_t) + \beta_{i,1}\xi_t,\ for\ i = 0, 1$$
$$\sigma^2_{\xi_t} = \sigma^2_0(1-\xi_t) + \sigma^2_1\xi_t$$
$$\xi_t = 0, 1$$
$$Pr[\xi_t = 0|\xi_{t-1} = 0] = P_{00}$$
$$Pr[\xi_t = 1|\xi_{t-1} = 1] = P_{11}$$

where ξ_t is state variable.

[19] I also check the robustness over different horizons and with the relative credit risk measures, and find the consistent results. Relative credit risk measures are defined as $D^{1,R}_t = D_t - D^{*,1}_t$, $L(D^R_t) = L(D_t) - L(D^*_t)$, $S(D^R_t) = S(D_t) - S(D^*_t)$. See Appendix for the results

After estimating the parameters by Maximum Likelihood Estimation (MLE), the filtered probabilities are computed by the Hamilton filter (Hamilton, 1989).[20]

The regression results are reported in **Table 2.5**. I first observe that adjusted R^2s are very high up to 72%. Even though the Markov-Switching model is estimated by Maximum Likelihood Estimation (MLE), which maximizes the log-likelihood of the model, I can compute adjusted R^2s from the linear combinations of two fitted values weighted by filtered probabilities of each state.[21] [22]

What does each regime imply? The volatility of state 1 appears to be relatively high compared to that of state 1 for all countries ($\sigma_0^2 < \sigma_1^2$). Consequently, state 0 represents a low-volatility state, and state 1 stands for a high-volatility state. I also explore when the volatility has been historically high by graphically checking the filtered probability of state 1 (P_1) as illustrated in **Figure 2.3**. High-volatility states for sample countries commonly include major global crises such as the Global Financial Crisis and European Fiscal Crisis.[23] Moreover, it seems noteworthy that the persistence of a high-volatility state is country-dependent. The persistence is found to be low for countries like Korea, but high for countries such as Mexico and Peru. The value of transition probability from state 1 to state 1 ($P_{11} = Pr[s_t = 1|s_{t-1} = 1]$) compares the country-specific persistence as it relates to the expected

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[20] Accurately, the model is estimated by Quasi-MLE. The model is not correctly specified because of auto-correlation in the error term due to overlapping data. However, it turns out that MLE estimates are consistent, while standard errors should be estimated by the robust covariance matrix. Specifically, let $\hat{\theta} = arg\min_{\theta\in\Theta} -lnL(\theta) = arg\min_{\theta\in\Theta} -lnf(y_t;\theta)$ and define the Gradient matrix evaluated at $\hat{\theta}$ as $G(\hat{\theta}) = \sum_{t=1}^{T} \frac{\partial}{\partial\theta} lnf(y_t;\hat{\theta}) = \sum_{t=1}^{T} G_t$, the Hessian matrix evaluated at $\hat{\theta}$ as $H(\hat{\theta}) = \sum_{t=1}^{T} \frac{\partial^2}{\partial\theta\partial\theta'} lnf(y_t;\hat{\theta})$. The robust covariance matrix in the existence of auto-correlation can be computed by $var(\hat{\theta}) = H(\hat{\theta})^{-1} J_P(\hat{\theta}) H(\hat{\theta})^{-1}$, where $J_P(\hat{\theta}) = \sum_{t=1}^{T} G_t G_t' + \sum_{i=1}^{p} w_i \left(\sum_{t=i+1}^{T} G_t G_{t-i}' + \sum_{t=i+1}^{T} G_{t-i} G_t'\right)$. P indicates that the approximation is curtailed at P lags of the auto-correlation, and w_i represents the weights with $\sum_{i=1}^{p} w_i = 1$.

[21] I obtain the estimated $\Delta\hat{s}_{t+3} = [\hat{\beta}_{0,0} + \hat{\beta}_{1,0} L(D_t) + \hat{\beta}_{2,0} S(D_t)]\hat{P}_0 + [\hat{\beta}_{0,0} + \hat{\beta}_{1,0} L(D_t) + \hat{\beta}_{2,0} S(D_t)]\hat{P}_1$ and compute the adjusted R^2 by using the residual sum of squares (RSS).

[22] Here, I do not report the goodness of fit measures from four different sets of risk factors as in **Table 2.4**. But, the model with both level and slope factors is also found to be superior to other models in terms of adjusted R^2 and AIC. The results can be provided upon request.

[23] Since it is difficult to clearly define exact periods for these crises, I use the Chicago Board Options Exchange (CBOE) Volatility Index (VIX index) to identify a highly volatile global financial market. Specifically, I indicate the periods when VIX index is greater than 20 along with the filtered probabilities.

duration of high-volatility state by $1/(1-P_{11})$.[24] The lower transition probability (P_{11}) a country exhibits, the longer the high-volatility state persists. The economic implication is that countries with a relatively more persistent high-volatility state suffered longer from the impact of global crises, compared to countries with less persistence.

The coefficient estimates confirm the findings from empirical studies on the carry-trade strategy (Clarida et al., 2009; Brunnermeier et al., 2008; Menkhoff et al., 2012). They show that when volatility is low, currency with higher risk appreciates as investors demand compensation for holding risky currency. However, under high volatility, as investors abruptly unwind their portfolios in favor of safe haven currencies, the value of the currency with higher risk plummets. Since the level and slope factors capture two types of riskiness – risk level and proximity – an increase in either level or slope is accompanied by significant appreciation of the currency in a low volatility state (state 0) and by depreciation or less appreciation of the currency in a high volatility state (state 1).[25] The concepts of risk level and proximity give strong insights on the sign-switching relationship between credit risk and currency returns. Recall that risk level is associated with the expected path of loss given credit events and risk proximity delivers information about the timing of the actual credit event. During normal times when a credit event is not likely to be realized shortly, both credit risk level and proximity forecast positive currency returns, as investors hold that currency to be compensated for bearing risk. However, when a global crisis is about to be triggered, a country with weak economic fundamentals is vulnerable to credit risk with potential default. Market participants with international portfolios anticipate that a default in this country is near in time and that expected loss is also escalated due to an increased default probability. Accordingly, they withdraw their investment and rebalance their portfolios so as to avoid highly probable losses, causing depreciation in the currencies of these countries.

[24]Expected duration of state 1: $E(D_{s_t=1}) = \sum_{j=1}^{\infty} P_{11}^{j-1}(1-P_{11}) = 1/(1-P_{11})$. For Korea, $E(D_{s_t=1}) = 1/(1-0.780) = 4.54$ months. For Mexico, $E(D_{s_t=1}) = 1/(1-0.923) = 13.01$ months.

[25]I have similar results from the regression of the excess currency returns, $xr_{t+3} (= i_t^3 - i_t^{3,*} - \Delta s_{t+3})$. Higher level and slope factors give high positive returns during a low volatility state, and incur losses or low returns during a high volatility state. The results can be provided upon request.

2.4.2 Comparing the Role of Risk Level vs. Risk Proximity

I demonstrated that both level and slope factor individually explain subsequent currency movements well. The next question is which of them accounts for more of the variation in currency returns. From the structural break model, the explanatory power of the level-only model is found to be higher than that of the slope-only model in 12 out of 20 countries for the full sample period. Risk level broadly seems to matter more than risk proximity. However, is this true regardless of the state of the economy? In order to answer this question, I run the same regressions repeatedly with 36-month rolling windows.[26] **Figure 2.4** shows the time-varying adjusted R^2s from the level-only and slope-only models. I first observe that the predictive ability of the level factor is relatively higher for a longer period of time. However, during the Global Financial Crisis period of 2008 - 2009, the slope factor explains a greater share of subsequent currency returns than does the level factor. This pattern is particularly apparent in developing countries such as Chile, Hungary, Indonesia, Israel, Korea, Mexico, Romania, and South Africa.

I further compare the statistical significance of two factors from both the level and slope models with the same rolling regression method.[27] $P-values$ of level and slope factors over sample periods are visualized in **Figure 2.5**. The contrast between normal times and crisis is obvious. The level factor is relatively more statistically significant during non-crisis periods, while the slope factor increases in significance during the crises. Again, this is especially true for developing countries. These observations suggest that risk level is the main determinant during normal times and risk proximity plays more critical role during crises. An intuitive explanation can be made by summing up the empirical findings so far.

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[26]Specifically, I run the following regressions with 36-month rolling windows by OLS estimation: 1) Level-only model: $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \epsilon_{t+3}$; 2) Slope-only model: $\Delta s_{t+3} = \beta_0 + \beta_1 S(D_t) + \epsilon_{t+3}$. I choose the window size of 36-month to obtain a clear picture for the comparison of goodness of fit measures. However, the choice of window size does not affect the key interpretation.

[27]I run the following regressions with 36-month rolling windows by OLS estimation: $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \beta_2 S(D_t) + \epsilon_{t+3}$. Then, $P-values$ of each regressor is computed for each window. If $P-value \leq 0.01$, I give "3". If $0.01 < P-value \leq 0.05$, I give "2". If $0.05 < P-value \leq 0.1$, I give "1". Otherwise, "0".

Suppose the economy is in a low volatility state, which include the periods before and after the Global Financial Crisis. The level is relatively moderate and the slope is steep. Investors are aware of some level of expected loss embedded by nature in the foreign asset subject to credit risk, but do not expect the actual event anytime soon. As a result, the risk level mainly drives currency returns, while the risk proximity has little role. However, in a high volatility state or near the Global Financial Crisis, the level is high and the slope becomes flatter or inverted. The currency depreciates due to a massive withdrawal of investments from the risky assets. Which is the main driver of re-weighting of investment from risky currencies to safe haven currencies: risk level or risk proximity? If investors perceive that a default is far from realization although the expected loss is large due to a concurrent global devaluation of assets, they would require even more compensation rather than unwind the position. However, if a specific credit event is very likely to be immediately triggered, investors would rebalance their position to avoid the losses, regardless of risk level. As a result, the link between risk proximity and currency returns becomes more pronounced for periods near major global crises.[28]

2.4.3 Risk-adjusted Uncovered Interest Rate Parity

This sub-section examines the role of sovereign credit risk in explaining the Uncovered Interest Rate Parity (UIP) puzzle. Since the reduced-form model in this chapter departs from the exact UIP relationship and is based on the net present value (NPV) framework, I briefly check whether the UIP puzzle is mitigated when incorporating credit risk factors as proxies for risk perceived at the current time t. Considering that one of the reasons for the UIP puzzle is ignoring risk premiums, I expect that inclusion of credit risk measures would help correct the abnormal UIP coefficients. I start with the original UIP regression, which is also known as the Fama Regression (Fama, 1984). This model is used as a benchmark to compare the results from the risk-adjusted UIP regression. I run the following regression

[28]Robustness checks are conducted over different horizons and with the relative credit risk measures. See Appendix.

and report the results in **Table 2.6**.

$$\Delta s_{t+3} = \beta_0 + \beta_1(i_t - i_t^*) + \epsilon_{t+3} \tag{2.13}$$

The null hypothesis that interest rate differentials have no information about subsequent exchange rate changes cannot be rejected and adjusted R^2s are close to zero. The UIP coefficients (β_1) are negative for seven countries, while positive for the other countries. The UIP puzzle is not as severe as in the existing empirical papers because the sample countries are mostly emerging economies and the sample period is relatively short including major global crises. Since the sample currencies are intrinsically risky, the deviation from the UIP relationship is smaller, as implied by Bansal and Dahlquist (2000) and Frankel and Poonawala (2010).[29] In addition, as shown from summary statistics, riskier currencies tend to depreciate more during the crisis. So, it is plausible to see positive coefficients when the sample period is exposed to severe financial turmoil. However, since the objective is to mitigate the UIP puzzle, I test whether the model augmented with risk premiums pushes the UIP coefficients toward positive for negative coefficient countries and improves the explanatory power.

Next, I augment the UIP with risk premiums, proxied by risk level (level factor) and proximity (slope factor). The regression equation is as follows.

$$\Delta s_{t+3} = \beta_0 + \beta_1(i_t - i_t^*) + \beta_2 L(D_t) + \beta_3 S(D_t) + \epsilon_{t+3} \tag{2.14}$$

The results are reported in **Table 2.7**. Compared to the results from the original UIP, I observe some improvement. The joint Wald test that regressors have no explanatory power cannot be rejected only for 5 countries at the 10% significance level and the goodness of fit measures increase up to 26%. I also notice improvement in the UIP coefficients. Out of seven countries which showed negative coefficients in the UIP regression, four countries now have positive coefficients and one country has less negative coefficients. Negative coefficients

[29] They find that the UIP coefficients are closer to 1 for emerging market currencies.

on risk measures are also consistent with theoretical predictions. Assuming risk aversion instead of risk neutrality, economic agents require a higher expected return on relatively riskier cross-border investments because of risk premiums. So, when the risk premium is ignored in the empirical regression of UIP relationship, it may cause the estimators to be biased due to the omitted variable problem (Obstfeld and Rogoff, 2000).[30] In this context, after incorporating risk premiums, measured by credit risk factors, the UIP puzzle could be mitigated.

2.5 Conclusions

This chapter relates sovereign credit risk level and proximity to currency returns. Theoretically, developing a CDS pricing model and constructing the level and slope factors from the term structure of sovereign credit default swap (CDS) spreads, I show that level represents how high the expected loss is expected – risk level – and slope implies how soon the actual credit event is likely to be realized – risk proximity. Combining with asset pricing models for defaultable bonds and exchange rate, I set up a model where the spot exchange rate is determined by both credit risk level and proximity. Empirically, I examine that the explanatory power of the model with both level and slope factors improves over the model with a single CDS spread, supporting my view that risk level and proximity capture different aspects of market participants' risk perception. Their relative role relies on the state of the economy as risk level matters more during normal times, while risk proximity increases in importance near crises. These findings suggest that market participants and policy authorities should closely monitor the movement of sovereign CDS spreads – not only the level but also the slope of CDS curves – to better understand the short-run currency behavior.

[30]Suppose the true model is $\Delta s_{t+3} = \beta_0 + \beta_1(i_t - i_t^*) - \rho_t + \epsilon_{t+3}$, where ρ_t is time-varying risk premium. The risk premium is positively correlated with the interest rate differential, as risk is generally higher in a higher interest rate country, and negatively correlated with exchange rate changes, as the risk premium compensates the investors. So, if the estimated model regresses $\Delta s_{t+3} = \beta_0 + \beta_1(i_t - i_t^*) + \epsilon_{t+3}$, then β_1 is negatively biased, resulting in the UIP puzzle.

Of course, there are many remaining tasks. Although I have restricted attention in this chapter to currency returns driven by credit risk due to clean measures of risk level and proximity from the term structure of CDS spreads, there is no reason why these measures cannot be obtained from the term structures of other asset classes such as bond yields and options. By investigating the pricing mechanism of different assets, I would be able to derive an interesting interpretation from the time-varying shape of term structures. Another next step would be to go beyond the reduced-form asset pricing model and to develop the DSGE model. By incorporating the term structure of assets into an otherwise general DSGE open economy model, I could gain insights on the connection between macroeconomic variables and asset prices.

Table 2.1: Summary Statistics for Exchange Rate and Interest Rates

	AU	BR	CL	CO	HU	IS	ID	IL	JP	KR
Panel A: Exchange Rate Change (Δs_{t+3})										
Mean										
Full	0.087	0.773	0.749	0.690	2.148	2.809	3.404	-1.784	0.370	-0.207
Pre-Crisis	-4.307	-13.291	-6.286	-8.658	-4.760	-2.443	2.137	-4.955	0.272	-5.525
Crisis	3.738	3.565	9.314	2.407	4.917	42.307	4.383	1.335	-9.483	16.465
Post-Crisis	1.561	7.311	2.554	5.064	5.073	-2.606	3.845	-0.819	2.433	-0.926
SD										
Full	27.644	33.594	23.711	28.927	30.373	33.081	20.228	16.901	21.064	22.155
Pre-Crisis	16.441	18.87	15.075	21.793	19.446	25.529	15.957	13.512	15.585	12.853
Crisis	59.086	60.240	48.206	49.812	59.310	67.001	41.599	30.490	28.362	46.174
Post-Crisis	21.895	30.072	18.802	25.486	25.931	17.587	15.349	14.467	21.420	16.458
AR(1)	0.727	0.715	0.674	0.681	0.689	0.723	0.724	0.721	0.710	0.684
Panel B: Interest Rate Differential ($i_t^3 - i_t^{3,*}$)										
Mean										
Full	2.746	11.081	3.111	4.462	4.377	7.146	6.403	1.211	-1.134	2.009
Pre-Crisis	2.235	12.021	0.486	2.972	4.571	7.072	5.753	0.618	-3.306	0.840
Crisis	3.803	10.860	4.821	7.731	8.037	14.251	9.392	2.125	-0.587	3.064
Post-Crisis	2.786	10.665	3.483	4.311	3.557	5.776	6.129	1.228	-0.179	2.373
SD										
Full	1.273	2.729	1.981	1.706	3.058	3.182	2.061	1.224	1.607	1.328
Pre-Crisis	1.188	3.737	0.801	1.301	3.200	1.959	2.230	0.808	1.303	1.202
Crisis	1.162	1.690	2.179	1.301	1.824	3.284	2.813	0.853	0.785	1.015
Post-Crisis	1.205	2.168	1.451	0.833	2.624	1.123	1.106	1.290	0.352	1.030
AR(1)	0.946	0.971	0.975	0.962	0.932	0.977	0.888	0.869	0.993	0.906

	MX	NO	PE	PH	PL	RO	ZA	SE	TH	TR
Panel A: Exchange Rate Change (Δs_{t+3})										
Mean										
Full	3.856	1.423	-0.508	-0.856	-0.234	1.632	5.126	1.123	-1.043	7.380
Pre-Cricis	-0.726	-6.906	-4.315	-8.176	-12.310	-7.396	2.268	-4.303	-4.622	-3.052
Crisis	13.133	7.042	0.010	10.401	11.710	11.073	3.016	7.521	2.042	14.563
Post-Crisis	4.276	4.485	1.309	0.544	3.429	4.266	7.002	2.558	0.135	11.185
SD										
Full	23.208	24.744	11.549	12.552	31.887	25.732	30.799	24.232	12.602	28.181
Pre-Crisis	10.312	18.470	7.998	11.808	21.866	21.315	27.972	19.032	12.886	26.331
Crisis	45.484	45.005	20.704	15.392	62.454	42.635	56.594	45.896	15.493	48.375
Post-Crisis	20.939	20.814	10.119	9.949	25.030	22.075	24.547	19.739	11.526	21.875
AR(1)	0.725	0.714	0.673	0.725	0.738	0.707	0.628	0.704	0.733	0.677
Panel B: Interest Rate Differential ($i_t^3 - i_t^{3,*}$)										
Mean										
Full	4.428	0.905	3.440	4.135	2.492	5.485	6.364	0.007	1.206	9.734
Pre-Crisis	4.439	-0.685	1.055	2.964	1.313	7.093	4.876	-1.131	-0.370	11.206
Crisis	6.432	3.110	5.226	9.562	4.552	11.186	10.155	1.706	1.605	14.406
Post-Crisis	4.025	1.246	4.254	3.634	2.661	3.569	6.342	0.228	1.899	8.318
SD										
Full	1.432	1.509	1.885	3.366	1.908	4.938	1.982	1.425	1.258	2.773
Pre-Crisis	1.876	1.106	0.706	2.814	2.351	6.496	1.664	1.342	0.554	1.471
Crisis	1.094	1.310	1.508	2.506	0.926	2.940	1.542	1.165	0.753	3.134
Post-Crisis	0.760	0.775	1.109	2.713	1.328	2.786	0.947	1.021	0.825	1.460
AR(1)	0.976	0.971	0.978	0.882	0.974	0.932	0.973	0.962	0.979	0.935

Note: 1. $\Delta s_{t+3} = s_{t+3} - s_t$ is the quarterly change in the exchange rate, where s_t is the logged home currency price per USD. 2. $i_t^3 - i_t^{3,*}$ is the difference in three-month zero coupon yields or interbank interest rates in the home and foreign countries (the US). 3. The sample period is from January 2004 to June 2017. All rates are reported in annualized percentage points. 5. The sample period is divided by two break dates: November 2007 and June 2009. Sub-periods are reported as Pre-Crisis, Crisis, and Post-Crisis, respectively.

Table 2.2: Principal Component Analysis for Term Structure of Sovereign CDSs

	AU	BR	CL	CO	HU	IS	ID	IL	JP	KR
Panel A: The first Principal Component (PC_{1t})										
Factor Loading										
1Y	0.326	0.176	0.278	0.202	0.366	0.502	0.402	0.297	0.151	0.394
2Y	0.337	0.314	0.351	0.291	0.414	0.451	0.421	0.359	0.223	0.415
3Y	0.405	0.398	0.386	0.397	0.422	0.419	0.439	0.405	0.297	0.422
5Y	0.458	0.473	0.453	0.482	0.421	0.384	0.410	0.457	0.435	0.414
7Y	0.454	0.490	0.458	0.488	0.415	0.355	0.396	0.457	0.541	0.406
10Y	0.448	0.498	0.485	0.497	0.409	0.309	0.379	0.447	0.598	0.399
Explained	88.277	95.662	95.952	92.826	97.111	98.397	96.249	94.779	95.068	97.156
$cor(PC_{2t}, L(D_t))$	0.936	0.986	0.979	0.975	0.979	0.978	0.963	0.967	0.986	0.974
Panel B: The second Principal Component (PC_{2t})										
Factor Loading										
1Y	0.428	0.604	0.471	0.635	0.648	0.562	0.556	0.546	0.515	0.512
2Y	0.493	0.524	0.486	0.525	0.365	0.326	0.368	0.449	0.508	0.401
3Y	0.330	0.312	0.357	0.262	0.134	0.035	0.178	0.300	0.449	0.243
5Y	-0.040	-0.102	-0.024	-0.099	-0.202	-0.314	-0.194	-0.102	0.148	-0.191
7Y	-0.357	-0.295	-0.412	-0.275	-0.379	-0.461	-0.424	-0.370	-0.217	-0.432
10Y	-0.579	-0.407	-0.494	-0.409	-0.495	-0.515	-0.553	-0.512	-0.454	-0.542
Explained	11.240	4.032	3.638	6.627	2.777	1.501	3.354	4.948	4.403	2.760
$cor(PC_{2t}, S(D_t))$	0.922	0.541	0.668	0.687	0.974	0.566	0.987	0.846	0.423	0.996

	MX	NO	PE	PH	PL	RO	ZA	SE	TH	TR
Panel A: The first Principal Component (PC_{1t})										
Factor Loading										
1Y	0.353	0.245	0.300	0.281	0.306	0.393	0.314	0.297	0.293	0.375
2Y	0.391	0.313	0.389	0.348	0.376	0.423	0.363	0.344	0.344	0.418
3Y	0.409	0.385	0.421	0.407	0.405	0.423	0.400	0.392	0.391	0.438
5Y	0.432	0.471	0.445	0.458	0.447	0.413	0.446	0.465	0.456	0.428
7Y	0.433	0.485	0.445	0.455	0.449	0.402	0.456	0.463	0.467	0.406
10Y	0.426	0.488	0.430	0.466	0.447	0.395	0.451	0.458	0.466	0.380
Explained	94.952	84.182	91.443	95.159	95.512	97.231	91.700	93.910	90.816	89.679
$cor(PC_{2t}, L(D_t))$	0.957	0.917	0.946	0.975	0.967	0.978	0.950	0.952	0.941	0.909
Panel B: The second Principal Component (PC_{2t})										
Factor Loading										
1Y	0.530	0.368	0.543	0.648	0.536	0.605	0.620	0.418	0.535	0.595
2Y	0.415	0.479	0.443	0.440	0.435	0.384	0.428	0.549	0.440	0.381
3Y	0.250	0.414	0.266	0.218	0.277	0.134	0.221	0.276	0.335	0.146
5Y	-0.094	0.078	-0.098	-0.131	-0.061	-0.232	-0.123	-0.018	-0.061	-0.185
7Y	-0.389	-0.279	-0.425	-0.315	-0.359	-0.410	-0.353	-0.302	-0.353	-0.416
10Y	-0.569	-0.616	-0.498	-0.472	-0.561	-0.496	-0.493	-0.597	-0.529	-0.521
Explained	4.734	14.808	7.432	4.674	4.143	2.649	7.960	5.415	8.987	9.516
$cor(PC_{2t}, S(D_t))$	0.954	0.856	0.900	0.801	0.846	0.997	0.921	0.830	0.887	0.994

Note: 1. Principal components are obtained as follows: 1)Let D_t denote the 6 x 1 vector of CDS spreads at each time t, Ω be the 6 x 6 covariance matrix of D_t; 2) The first principal component is $PC_{1t} = p_1^{'} D_t$, where p_1 is an eigenvector with the largest eigenvalue from Ω; 3) The second principal component is $PC_{2t} = p_2^{'} D_t$, where p_2 is an eigenvector with the second largest eigenvalue from Ω. 2. "Factor loadings" are p_1 and p_2, which represent the weight on each D_t^m of $m-$tenor. 3. "Explained" indicates how much of the variation in D_t is explained by each principal component. 4. In order to show the meaning of each principal component, I compute the correlation between the first principal component and the geometric level ($L(D_t) = D_t^{10}$), and the correlation between the second principal component and the geometric slope ($S(D_t) = D_t^1 - D_t^{10}$).

Table 2.3: Summary Statistics for Level and Slope factors of Sovereign CDSs

	AU	BR	CL	CO	HU	IS	ID	IL	JP	KR
Panel A: Level factor = The first Principal Component ($L(D_t) = PC_{1t}$)										
Mean										
Full	87.40	472.22	171.31	397.35	429.28	531.63	427.24	199.04	102.97	175.44
Pre-Crisis	30.33	567.71	49.01	495.38	55.76	26.90	374.38	70.73	14.52	69.63
Crisis	134.45	464.09	276.71	541.64	608.56	1375.72	882.01	286.31	83.05	453.56
Post-Crisis	100.68	429.07	193.76	322.85	572.77	474.98	355.95	236.56	150.22	172.19
SD										
Full	57.56	284.64	101.04	201.56	372.17	562.97	255.13	115.78	77.97	165.27
Pre-Crisis	17.16	409.56	8.75	271.46	20.13	27.58	114.88	21.28	6.61	28.46
Crisis	104.7	226.01	174.84	234.77	438.95	866.66	472.59	158.75	59.20	316.02
Post-Crisis	34.43	204.99	44.76	95.38	319.05	325.36	83.67	82.79	58.42	77.16
AR(1)	0.883	0.906	0.919	0.894	0.967	0.956	0.900	0.947	0.958	0.919
Panel B: Slope factor = The second Principal Component ($S(D_t) = PC_{2t}$)										
Mean										
Full	-24.84	-41.83	-30.85	-40.66	-75.21	-101.49	-170.32	-44.21	-5.87	-52.52
Pre-Crisis	-12.89	-80.60	-12.27	-90.22	-24.15	-16.61	-179.86	-22.81	-1.05	-26.18
Crisis	-11.59	17.61	-10.30	34.95	-28.94	-111.18	-151.12	-18.86	17.46	-31.20
Post-Crisis	-32.20	-35.42	-41.49	-32.40	-108.83	-118.14	-170.74	-58.37	-12.85	-69.64
SD										
Full	17.77	58.43	19.67	53.85	62.94	69.52	47.63	26.45	16.78	27.86
Pre-Crisis	8.11	48.34	2.62	35.38	8.61	7.57	38.18	6.51	1.80	10.65
Crisis	8.06	68.73	25.85	73.90	84.74	49.50	80.42	30.89	23.48	21.40
Post-Crisis	17.9	47.72	12.31	29.39	49.43	67.09	41.18	20.35	14.20	21.04
AR(1)	0.940	0.885	0.864	0.944	0.911	0.916	0.800	0.904	0.918	0.912

	MX	NO	PE	PH	PL	RO	ZA	SE	TH	TR
Panel A: Level factor = The first Principal Component ($L(D_t) = PC_{1t}$)										
Mean										
Full	265.88	43.09	315.21	426.51	195.97	403.92	350.44	66.65	216.73	486.88
Pre-Crisis	162.97	14.42	309.80	638.08	41.87	97.68	143.21	33.29	88.22	497.33
Crisis	462.37	50.90	480.71	655.31	315.65	777.71	567.41	109.36	354.31	709.23
Post-Crisis	275.23	48.11	284.49	277.64	247.72	463.92	408.96	66.19	252.41	438.08
SD										
Full	145.15	22.88	130.75	263.95	161.09	336.33	194.56	53.13	117.98	187.14
Pre-Crisis	64.67	11.03	109.77	312.17	17.46	47.41	68.28	41.96	16.57	229.68
Crisis	294.33	36.92	254.28	227.54	245.17	520.17	282.41	103.40	177.78	256.50
Post-Crisis	63.93	15.38	62.11	77.89	122.96	245.92	112.52	31.66	65.01	97.76
AR(1)	0.906	0.903	0.867	0.921	0.948	0.943	0.942	0.874	0.920	0.800
Panel B: Slope factor = The second Principal Component ($S(D_t) = PC_{2t}$)										
Mean										
Full	-79.91	-11.25	-86.60	-85.16	-42.50	-94.42	-76.18	-17.66	-45.75	-157.50
Pre-Crisis	-67.50	-3.92	-124.72	-129.65	-15.68	-43.37	-50.16	-17.17	-23.56	-204.11
Crisis	-54.73	1.53	-55.36	-2.86	-4.97	-75.80	-7.51	-1.37	0.38	-114.63
Post-Crisis	-90.71	-15.46	-78.49	-79.67	-63.05	-120.44	-102.50	-21.00	-65.75	-144.63
SD										
Full	32.41	8.51	37.27	58.50	33.55	55.52	57.32	12.76	37.11	60.96
Pre-Crisis	29.03	3.52	43.84	46.02	6.84	18.67	24.48	22.20	4.10	37.44
Crisis	26.39	5.78	27.57	92.87	40.06	71.63	84.10	4.74	39.66	100.41
Post-Crisis	30.30	5.62	23.01	27.86	22.72	45.09	44.28	7.20	30.90	45.77
AR(1)	0.939	0.919	0.902	0.948	0.927	0.908	0.939	0.756	0.939	0.914

Note: 1. I denote the first principal component as the “level”factor and the second principal component as the “slope”factor. 2. The sample period is from January 2004 to June 2017. The sample period is shorter for some countries due to data availability. All values are in annuity in basis points. 3. The sample period is divided by two break date: November 2007 and June 2009. Sub-periods are reported as Pre-Crisis, Crisis, and Post-Crisis, respectively.

Table 2.4: Explaining Currency Returns with Credit Risk: Structural Break Model

	AU	BR	CL	CO	HU	IS	ID	IL	JP	KR
Model 1: $\Delta s_{t+3} = \beta_0 + \beta_1 D_t^1 + \epsilon_{t+3}$										
$p-value$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.004	0.000
$ad.R^2$	0.338	0.123	0.238	0.151	0.308	0.351	0.190	0.209	0.076	0.259
Model 2: $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \epsilon_{t+3}$										
$p-value$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
$ad.R^2$	0.278	0.228	0.400	0.270	0.298	0.348	0.168	0.121	0.155	0.344
Model 3: $\Delta s_{t+3} = \beta_0 + \beta_1 S(D_t) + \epsilon_{t+3}$										
$p-value$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.148	0.000	0.000
$ad.R^2$	0.229	0.113	0.464	0.119	0.354	0.315	0.333	0.008	0.228	0.411
Model 4: $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \beta_2 S(D_t) + \epsilon_{t+3}$										
$p-value$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.162	0.000	0.000
$ad.R^2$	0.601	0.239	0.551	0.326	0.345	0.344	0.368	0.011	0.215	0.469

	MX	NO	PE	PH	PL	RO	ZA	SE	TH	TR
Model 1: $\Delta s_{t+3} = \beta_0 + \beta_1 D_t^1 + \epsilon_{t+3}$										
$p-value$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.001	0.000
$ad.R^2$	0.140	0.313	0.165	0.278	0.187	0.215	0.176	0.231	0.087	0.110
Model 2: $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \epsilon_{t+3}$										
$p-value$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.467	0.001
$ad.R^2$	0.128	0.259	0.241	0.266	0.209	0.172	0.257	0.238	-0.004	0.015
Model 3: $\Delta s_{t+3} = \beta_0 + \beta_1 S(D_t) + \epsilon_{t+3}$										
$p-value$	0.090	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.008	0.000
$ad.R^2$	-0.001	0.210	0.169	0.230	0.197	0.211	0.243	0.188	0.048	0.230
Model 4: $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \beta_2 S(D_t) + \epsilon_{t+3}$										
$p-value$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.017	0.000
$ad.R^2$	0.135	0.380	0.278	0.322	0.312	0.173	0.338	0.173	0.045	0.342

Note: 1. I first regress quarterly exchange rate changes on four different sets of risk factors and then apply the Bai and Perron (2003) test(with 15% trimming and 5 - 10% significance level) to detect the multiple structural breaks in the regression. Zero to two breaks are detected depending on sample countries. 2. After identifying the break dates, structural break dummy variables for each sub-period are incorporated into the regression. 3. P-value is for the Wald test that factors jointly have no explanatory power 4. Adjusted R^2 is reported.

Table 2.5: Explaining Currency Returns with Credit Risk: Markov Switching Model

$$\Delta s_{t+3} = \beta_{0,\xi_t} + \beta_{1,\xi_t} L(D_t) + \beta_{2,\xi_t} S(D_t) + \epsilon_{t+3},\ \ where\ \epsilon_{t+3}\ i.i.d.N(0, \sigma^2_{\xi_t})$$

	AU	BR	CL	CO	HU	IS	ID	IL	JP	KR
$\beta_{0,0}$	-4.198	1.968	1.815	11.851	-12.176**	-14.430***	5.753	-16.928***	7.955**	-4.482
	(3.994)	(5.061)	(9.686)	(7.674)	(5.927)	(4.369)	(8.933)	(6.219)	(3.108)	(5.086)
$\beta_{0,1}$	97.863***	68.338***	-51.302***	130.680***	35.107	69.635*	59.529	10.136	-22.892***	6.124
	(16.376)	(13.859)	(6.179)	(23.621)	(32.727)	(36.905)	(49.895)	(7.418)	(7.459)	(12.501)
$\beta_{1,0}$	-0.163**	-0.052***	-0.061**	-0.062***	-0.019***	0.011***	-0.040***	-0.006	0.080**	-0.046***
	(0.072)	(0.014)	(0.026)	(0.017)	(0.006)	(0.003)	(0.006)	(0.017)	(0.031)	(0.012)
$\beta_{1,1}$	-0.467***	-0.056***	-0.112***	-0.222***	-0.005	-0.083**	0.005	0.006	0.084*	0.005
	(0.054)	(0.015)	(0.038)	(0.052)	(0.039)	(0.040)	(0.020)	(0.026)	(0.048)	(0.028)
$\beta_{2,0}$	-0.342***	-0.124**	-0.271**	-0.072	-0.174***	-0.055*	-0.058	-0.163***	-0.433**	-0.103
	(0.103)	(0.059)	(0.124)	(0.056)	(0.033)	(0.030)	(0.047)	(0.050)	(0.191)	(0.072)
$\beta_{2,1}$	0.863**	-0.007	-4.391***	0.476***	-0.130	-0.517	0.224	-0.119	-0.186	-0.456
	(0.428)	(0.062)	(0.664)	(0.131)	(0.142)	(0.329)	(0.299)	(0.140)	(0.198)	(0.746)
σ_0	12.654***	16.919**	15.783**	19.116**	17.858**	16.723**	9.541**	10.986**	12.376*	13.034
	(4.537)	(7.188)	(6.849)	(8.454)	(7.399)	(6.637)	(4.208)	(4.700)	(6.779)	(10.308)
σ_1	21.778*	26.913*	21.035	25.918**	29.949*	60.536	31.679	13.126**	15.273**	24.929
	(12.498)	(15.029)	(17.477)	(13.073)	(16.853)	(37.525)	(20.187)	(5.229)	(7.189)	(22.049)
P_{00}	0.942***	0.943***	0.982***	0.965***	0.952***	0.991***	0.965***	0.936***	0.833***	0.954
	(0.225)	(0.238)	(0.280)	(0.260)	(0.250)	(0.359)	(0.251)	(0.219)	(0.196)	(0.832)
P_{11}	0.883**	0.843***	0.893	0.893***	0.788***	0.897*	0.864***	0.844***	0.884***	0.780
	(0.356)	(0.215)	(0.572)	(0.322)	(0.241)	(0.524)	(0.285)	(0.260)	(0.290)	(0.522)
$ad.R^2$	0.687	0.716	0.526	0.559	0.625	0.457	0.296	0.650	0.651	0.557
AIC	8.685	9.236	8.695	9.166	9.171	8.953	8.219	8.207	8.659	8.573

Note: 1. The model is estimated by a two-state Markov-Switching Model. After estimating the parameters by Maximum Likelihood Estimation (MLE), the filtered probabilities are computed by Hamilton filter. See the text for model specification. 2. Coefficient estimates are reported with the Newey-West standard errors in parentheses. Asterisks indicate significance levels at 1% (***), 5% (**), and 10% (*), respectively. 3. By using the filtered probabilities, I obtain the estimated $\Delta \hat{s}_{t+3} = [\hat{\beta}_{0,0} + \hat{\beta}_{1,0} L(D_t) + \hat{\beta}_{2,0} S(D_t)]\hat{P}_0 + [\hat{\beta}_{0,0} + \hat{\beta}_{1,0} L(D_t) + \hat{\beta}_{2,0} S(D_t)]\hat{P}_1$ and compute adjusted R^2 by using the residual sum of squares(RSS). 4. AIC is provided.

(Continued)

$$\Delta s_{t+3} = \beta_{0,\xi_t} + \beta_{1,\xi_t} L(D_t) + \beta_{2,\xi_t} S(D_t) + \epsilon_{t+3},\ where\ \epsilon_{t+3}\ i.i.d.N(0, \sigma^2_{\xi_t})$$

	MX	NO	PE	PH	PL	RO	ZA	SE	TH	TR
$\beta_{0,0}$	4.237	82.263***	7.944***	2.576	-17.702***	-18.356***	-4.807	-5.645	16.052***	5.885
	(4.371)	(9.919)	(1.870)	(3.615)	(5.140)	(5.751)	(4.214)	(3.631)	(4.222)	(16.161)
$\beta_{0,1}$	79.158**	-3.358	7.460	-5.527	50.751**	4.417	49.700***	62.594***	-11.750***	-1.196
	(39.635)	(5.957)	(7.970)	(12.627)	(23.429)	(9.072)	(9.636)	(16.367)	(2.291)	(46.904)
$\beta_{1,0}$	-0.043***	0.081	-0.039***	-0.007	-0.020	-0.005	-0.049***	-0.152***	-0.017	-0.026
	(0.008)	(0.238)	(0.003)	(0.005)	(0.013)	(0.007)	(0.009)	(0.033)	(0.011)	(0.020)
$\beta_{1,1}$	-0.046	-0.183**	0.000	0.084***	0.034	0.028***	-0.043**	-0.195***	0.005	0.022
	(0.042)	(0.081)	(0.005)	(0.026)	(0.047)	(0.009)	(0.020)	(0.075)	(0.008)	(0.056)
$\beta_{2,0}$	-0.025	3.514***	0.003	-0.006	-0.236***	-0.088***	-0.147*	-0.239*	-0.059	-0.007
	(0.054)	(0.887)	(0.012)	(0.028)	(0.071)	(0.032)	(0.080)	(0.128)	(0.048)	(0.051)
$\beta_{2,1}$	0.427**	-0.669***	0.029	0.317***	0.318**	-0.112	0.045	1.356*	-0.086***	-0.215
	(0.205)	(0.208)	(0.058)	(0.116)	(0.133)	(0.095)	(0.138)	(0.721)	(0.027)	(0.240)
σ_0	11.083**	16.066*	4.192*	9.294**	17.906**	15.737**	16.695*	13.880**	6.596**	15.089*
	(5.520)	(9.131)	(2.178)	(3.816)	(7.371)	(7.626)	(8.620)	(5.945)	(2.947)	(7.748)
σ_1	26.617*	16.546**	13.795**	15.840*	27.974	19.090*	24.670	18.352*	7.900**	24.225
	(15.600)	(6.624)	(6.676)	(9.030)	(18.762)	(10.800)	(16.364)	(9.594)	(3.404)	(15.114)
P_{00}	0.961***	0.773***	0.933***	0.991***	0.941***	0.896***	0.905***	0.932***	0.773***	0.916***
	(0.287)	(0.274)	(0.264)	(0.314)	(0.211)	(0.177)	(0.213)	(0.208)	(0.196)	(0.181)
P_{11}	0.923*	0.957***	0.923***	0.972**	0.762***	0.811***	0.855	0.876***	0.924***	0.847**
	(0.484)	(0.298)	(0.238)	(0.489)	(0.220)	(0.183)	(0.776)	(0.247)	(0.190)	(0.397)
$ad.R^2$	0.516	0.620	0.262	0.198	0.665	0.667	0.656	0.672	0.708	0.682
AIC	8.527	8.804	7.281	7.729	9.220	9.020	9.261	8.739	7.461	9.105

Note: 1. The model is estimated by a two-state Markov-Switching Model. After estimating the parameters by Maximum Likelihood Estimation (MLE), the filtered probabilities are computed by Hamilton filter. See the text for model specification. 2. Coefficient estimates are reported with the Newey-West standard errors in parentheses. Asterisks indicate significance levels at 1% (***), 5% (**), and 10% (*), respectively. 3. By using the filtered probabilities, I obtain the estimated $\Delta\hat{s}_{t+3} = [\hat{\beta}_{0,0} + \hat{\beta}_{1,0}L(D_t) + \hat{\beta}_{2,0}S(D_t)]\hat{P}_0 + [\hat{\beta}_{0,0} + \hat{\beta}_{1,0}L(D_t) + \hat{\beta}_{2,0}S(D_t)]\hat{P}_1$ and compute adjusted R^2 by using the residual sum of squares(RSS). 4. AIC is provided.

Table 2.6: Uncovered Interest Rate Parity: OLS

$$\Delta s_{t+3} = \beta_0 + \beta_1(i_t^3 - i_t^{3,*}) + \epsilon_{t+3}$$

	AU	BR	CL	CO	HU	IS	ID	IL	JP	KR
β_0	-9.299	2.702	-7.894*	-6.550	3.670	-21.789*	10.691	-5.717**	1.217	-7.400**
	(7.665)	(15.105)	(4.589)	(10.875)	(4.843)	(12.962)	(8.573)	(2.752)	(3.427)	(3.561)
β_1	3.482	-0.189	3.182*	1.915	-0.330	3.313	-1.214	3.049	0.747	3.524
	(3.442)	(1.429)	(1.759)	(2.691)	(1.384)	(2.030)	(1.595)	(1.877)	(1.305)	(2.178)
$p-value$	0.058	0.850	0.002	0.187	0.690	0.000	0.126	0.012	0.472	0.008
$ad.R^2$	0.018	-0.006	0.061	0.005	-0.005	0.102	0.009	0.036	-0.003	0.038

	MX	NO	PE	PH	PL	RO	ZA	SE	TH	TR
β_0	6.387	-3.964	-4.563*	-1.368	-7.671*	3.135	11.634	0.249	-3.698	16.852
	(12.260)	(3.621)	(2.557)	(2.508)	(4.139)	(4.919)	(15.333)	(3.204)	(2.386)	(17.499)
β_1	-0.600	5.273*	1.277	0.122	2.940	-0.330	-1.021	5.472**	2.175*	-0.916
	(3.180)	(2.876)	(0.783)	(0.489)	(2.470)	(1.121)	(2.632)	(2.512)	(1.208)	(1.966)
$p-value$	0.646	0.000	0.011	0.682	0.027	0.503	0.407	0.000	0.006	0.276
$ad.R^2$	-0.005	0.082	0.037	-0.005	0.025	-0.004	-0.002	0.092	0.042	0.001

Note: 1. The model is estimated by OLS. 2. Coefficient estimates are reported with the Newey-West standard errors in parentheses. Asterisks indicate significance levels at 1% (***), 5% (**), and 10% (*), respectively. 3. P-value is for the Wald test that factors jointly have no explanatory power 4. Adjusted R^2 is reported.

Table 2.7: Risk-adjusted Uncovered Interest Rate Parity: OLS

$$\Delta s_{t+3} = \beta_0 + \beta_1(i_t^3 - i_t^{3,*}) + \beta_2 L(D_t) + \beta_3 S(D_t) + \epsilon_{t+3}$$

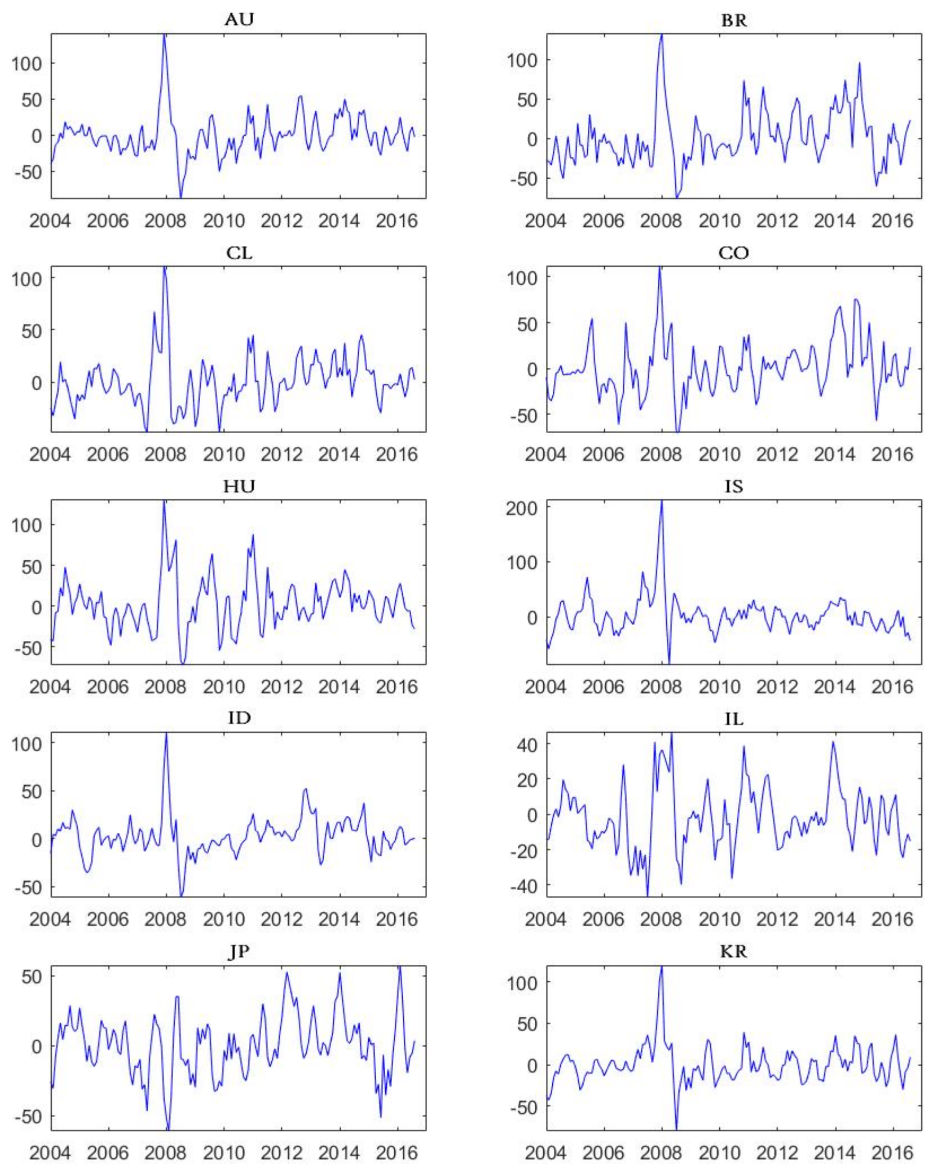
	AU	BR	CL	CO	HU	IS	ID	IL	JP	KR
β_0	-18.366***	-26.103	-2.481	-0.760	-16.526**	-27.705*	-10.089	-9.667*	-5.617	-2.557
	(6.135)	(17.977)	(6.724)	(11.786)	(7.581)	(15.897)	(11.084)	(5.629)	(7.618)	(4.305)
β_1	10.401***	5.758**	5.551***	6.381*	2.294	4.081	-1.355	3.711	-0.873	6.506*
	(3.814)	(2.616)	(2.142)	(3.334)	(1.744)	(2.482)	(1.809)	(2.276)	(1.945)	(3.351)
β_2	-0.258***	-0.066**	-0.109***	-0.078***	-0.014	-0.007	-0.008	-0.007	0.037	-0.037**
	(0.059)	(0.026)	(0.038)	(0.025)	(0.011)	(0.011)	(0.011)	(0.023)	(0.045)	(0.018)
β_3	-0.530***	0.149**	-0.215*	-0.068	-0.199***	-0.041	-0.147***	-0.101	-0.212	0.083
	(0.139)	(0.066)	(0.121)	(0.058)	(0.062)	(0.046)	(0.044)	(0.078)	(0.153)	(0.094)
$p-value$	0.000	0.000	0.000	0.000	0.000	0.001	0.000	0.020	0.132	0.001
$ad.R^2$	0.256	0.133	0.258	0.131	0.120	0.105	0.127	0.048	0.017	0.084

	MX	NO	PE	PH	PL	RO	ZA	SE	TH	TR
β_0	3.154	4.372	-4.751	0.487	-12.829**	-6.535	-2.767	9.930	-3.611	20.200
	(11.723)	(7.720)	(4.604)	(4.600)	(5.308)	(7.660)	(16.049)	(7.600)	(4.017)	(18.569)
β_1	0.061	7.558**	2.286***	0.325	3.647	-0.370	2.827	7.024**	2.364	-0.179
	(3.514)	(3.215)	(0.822)	(0.795)	(3.619)	(1.296)	(3.273)	(2.878)	(1.663)	(2.611)
β_2	-0.015	-0.433***	-0.029***	-0.006	-0.016	0.007	-0.056**	-0.136**	-0.011	-0.024
	(0.021)	(0.151)	(0.011)	(0.005)	(0.035)	(0.012)	(0.027)	(0.063)	(0.015)	(0.029)
β_3	-0.054	-0.682***	-0.066**	0.001	-0.158	-0.073	-0.126**	0.044	-0.048	-0.004
	(0.083)	(0.256)	(0.031)	(0.029)	(0.117)	(0.059)	(0.061)	(0.225)	(0.037)	(0.068)
$p-value$	0.545	0.000	0.000	0.515	0.017	0.131	0.000	0.000	0.005	0.361
$ad.R^2$	-0.006	0.227	0.143	-0.004	0.045	0.017	0.106	0.160	0.061	0.002

Note: 1. The model is estimated by OLS. 2. Coefficient estimates are reported with the Newey-West standard errors in parentheses. Asterisks indicate significance levels at 1% (***), 5% (**), and 10% (*), respectively. 3. P-value is for the Wald test that factors jointly have no explanatory power 4. Adjusted R^2 is reported.

Figure 2.1: Exchange Rate Change

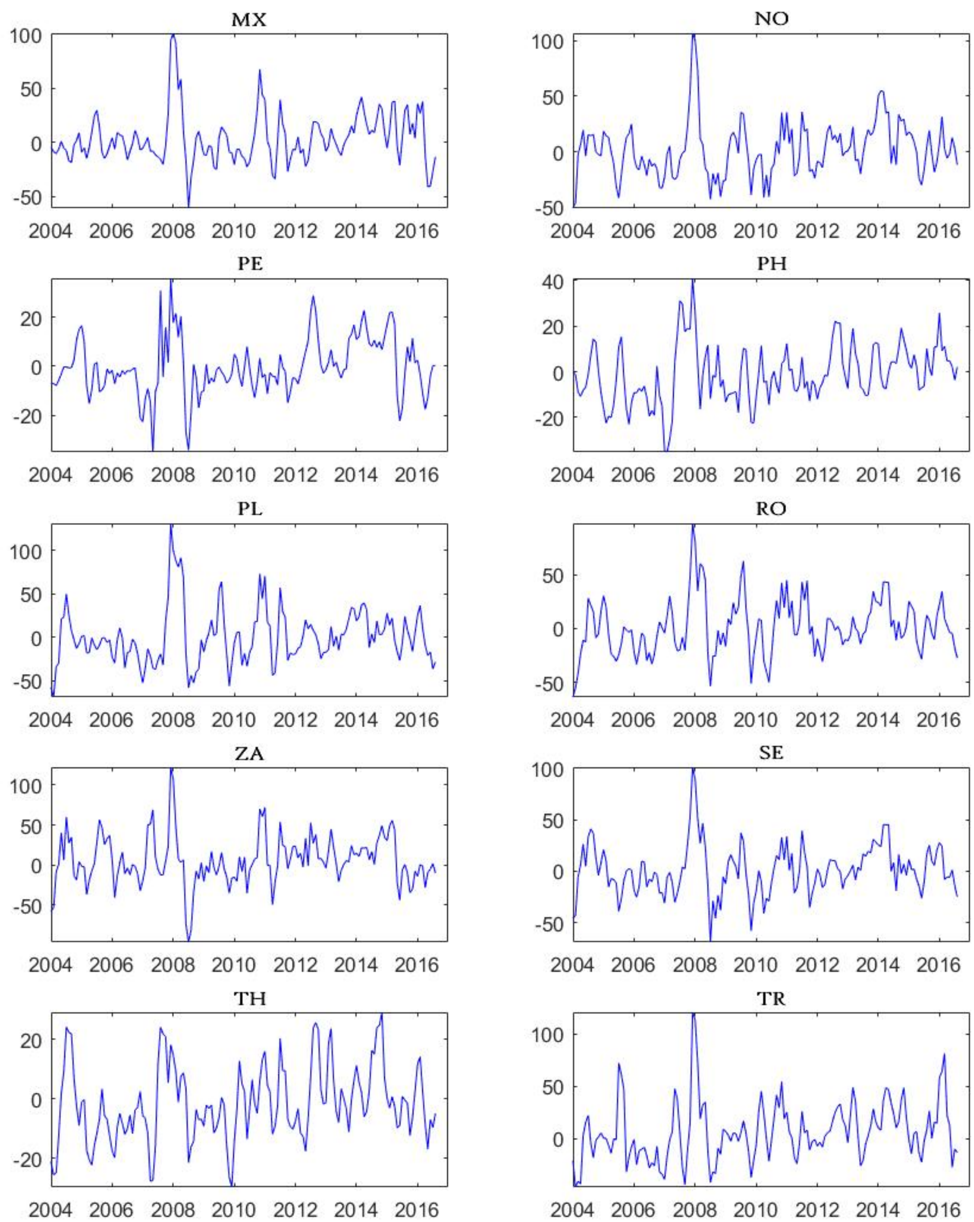
(Annualized %; Home Currency/USD)



Note: 1. This figure shows the quarterly change of the exchange rate, $\Delta s_{t+3} = s_{t+3} - s_t$, where s_t is the logged home currency price per USD. 2. The sample period is from January 2004 to June 2017. All rates are in annualized percentage points.

(Continued)

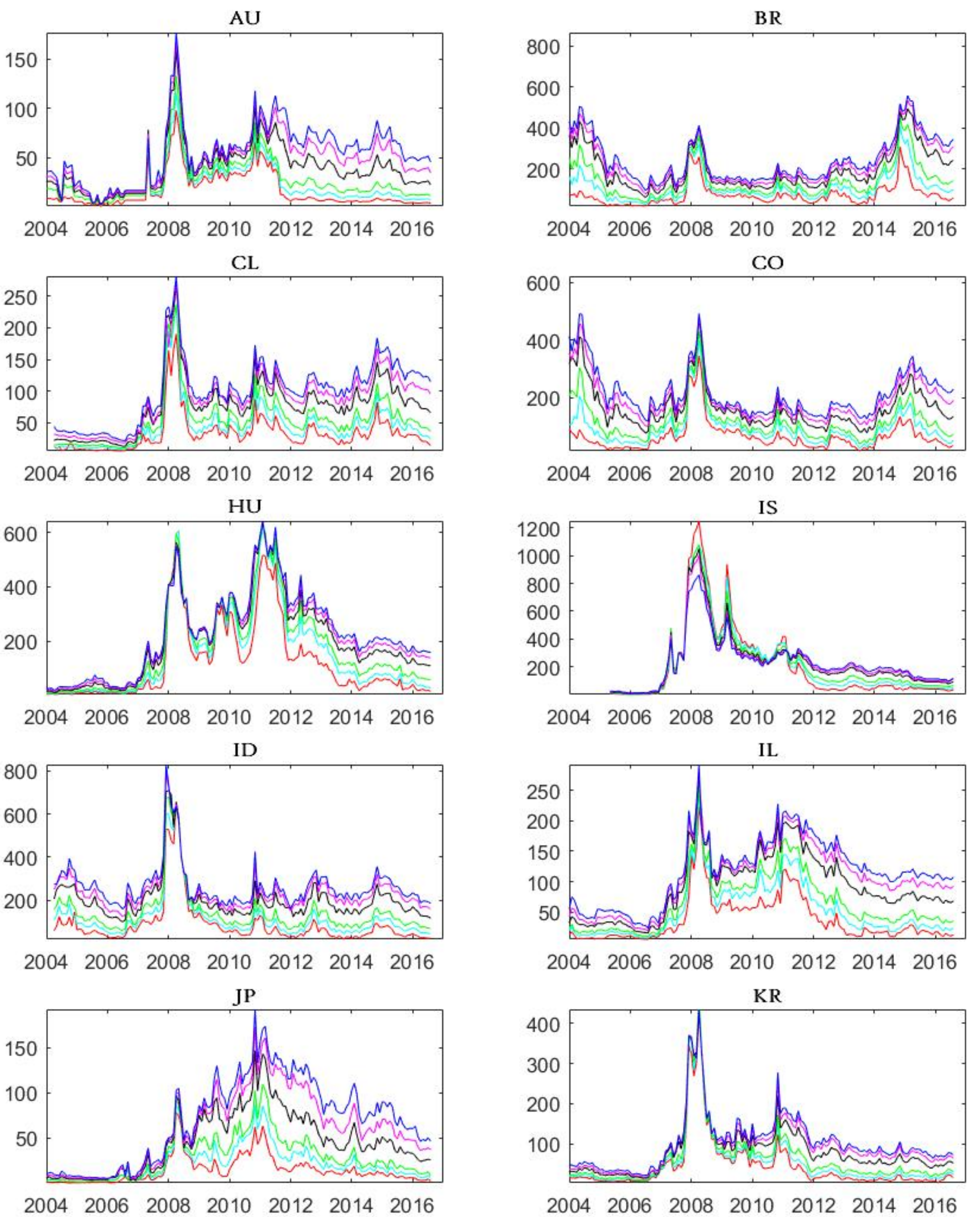
(Annualized %; Home Currency/USD)



Note: 1. This figure shows the quarterly change of the exchange rate, $\Delta s_{t+3} = s_{t+3} - s_t$, where s_t is the logged home currency price per USD. 2. The sample period is from January 2004 to June 2017. All rates are in annualized percentage points.

Figure 2.2: Sovereign Credit Default Swap Spreads

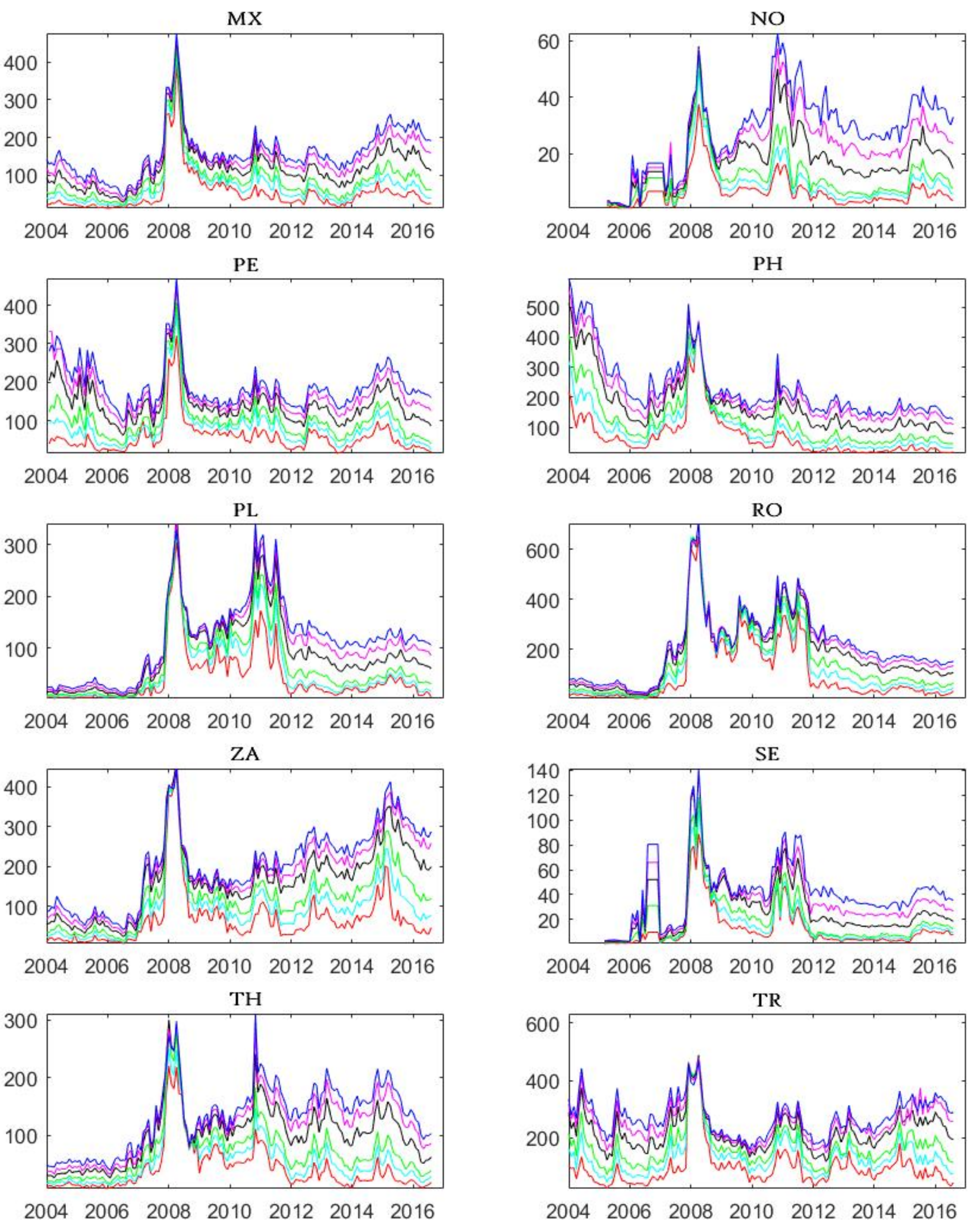
(Basis Point)



Note: 1. This figure shows sovereign credit default swap (CDS) spreads of 1, 2, 3, 5, 7, and 10 year tenors: Red for 1 year, Cyan for 2 year, Green for 3 year, Black for 5 year, Magenta for 7 year and Blue for 10 year tenor CDS spread. 2. The sample period is from January 2004 to June 2017, but is shorter for some countries. All values are in annuity basis points.

(Continued)

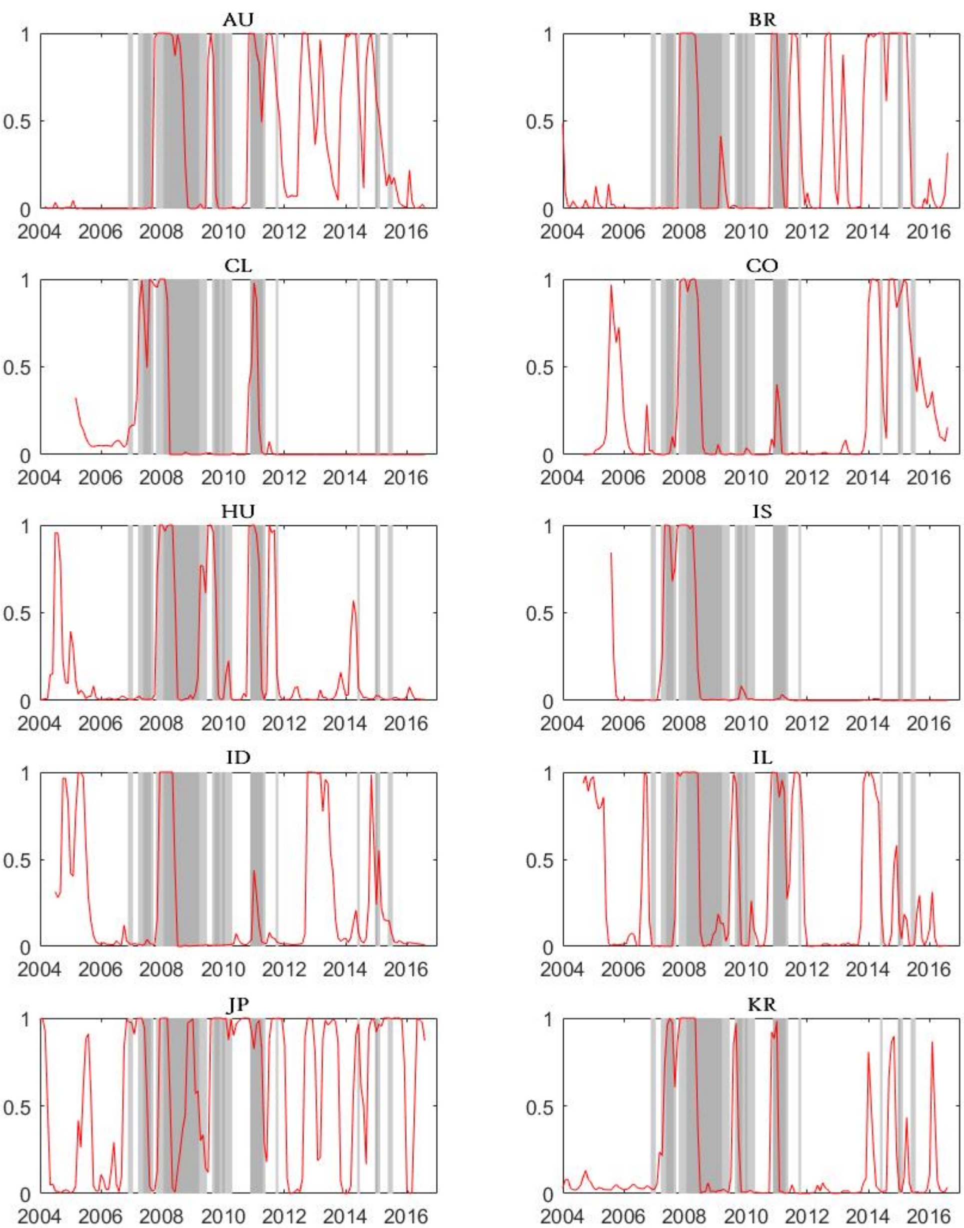
(Basis Point)



Note: 1. This figure shows sovereign credit default swap (CDS) spreads of 1, 2, 3, 5, 7, and 10 year tenors: Red for 1 year, Cyan for 2 year, Green for 3 year, Black for 5 year, Magenta for 7 year and Blue for 10 year tenor CDS spread. 2. The sample period is from January 2004 to June 2017, but is shorter for some countries. All values are in annuity basis points.

Figure 2.3: Probability in High-Volatility State from Markov-Switching Model

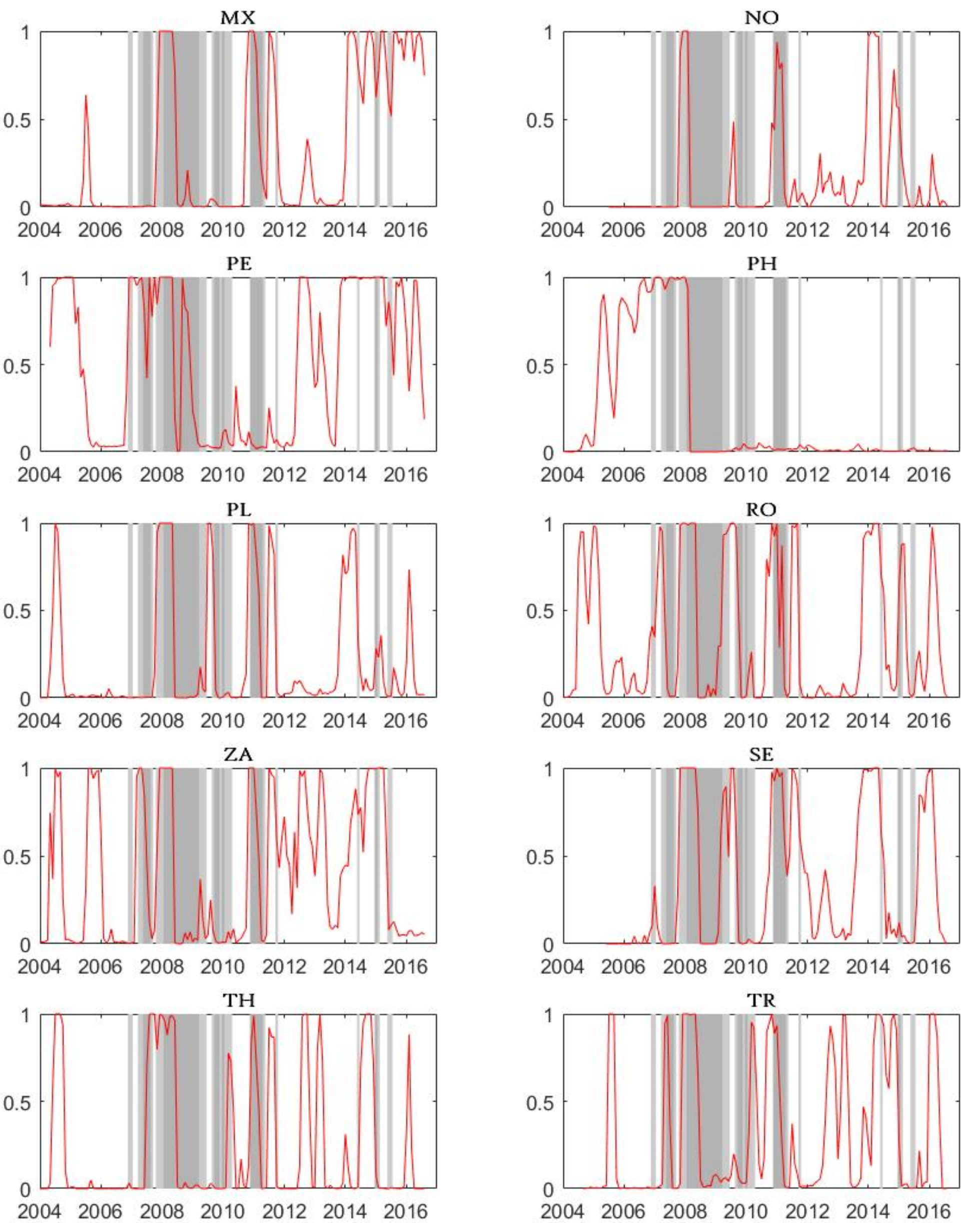
(%)



Note: 1. This figure shows the filtered probability in state 1 from Markov-Switching model estimation: $\Delta s_{t+3} = \beta_{0,\xi_t} + \beta_{1,\xi_t} L(D_t) + \beta_{2,\xi_t} S(D_t) + \epsilon_{t+3},\ where\ \epsilon_{t+3}\ i.i.d.N(0, \sigma^2_{\xi_t})$. State 1 represents a "high volatility"state. 2. Dark shades show the months when the VIX index is greater than 25 and light shades show the months when the VIX index is between 20 and 25.

(Continued)

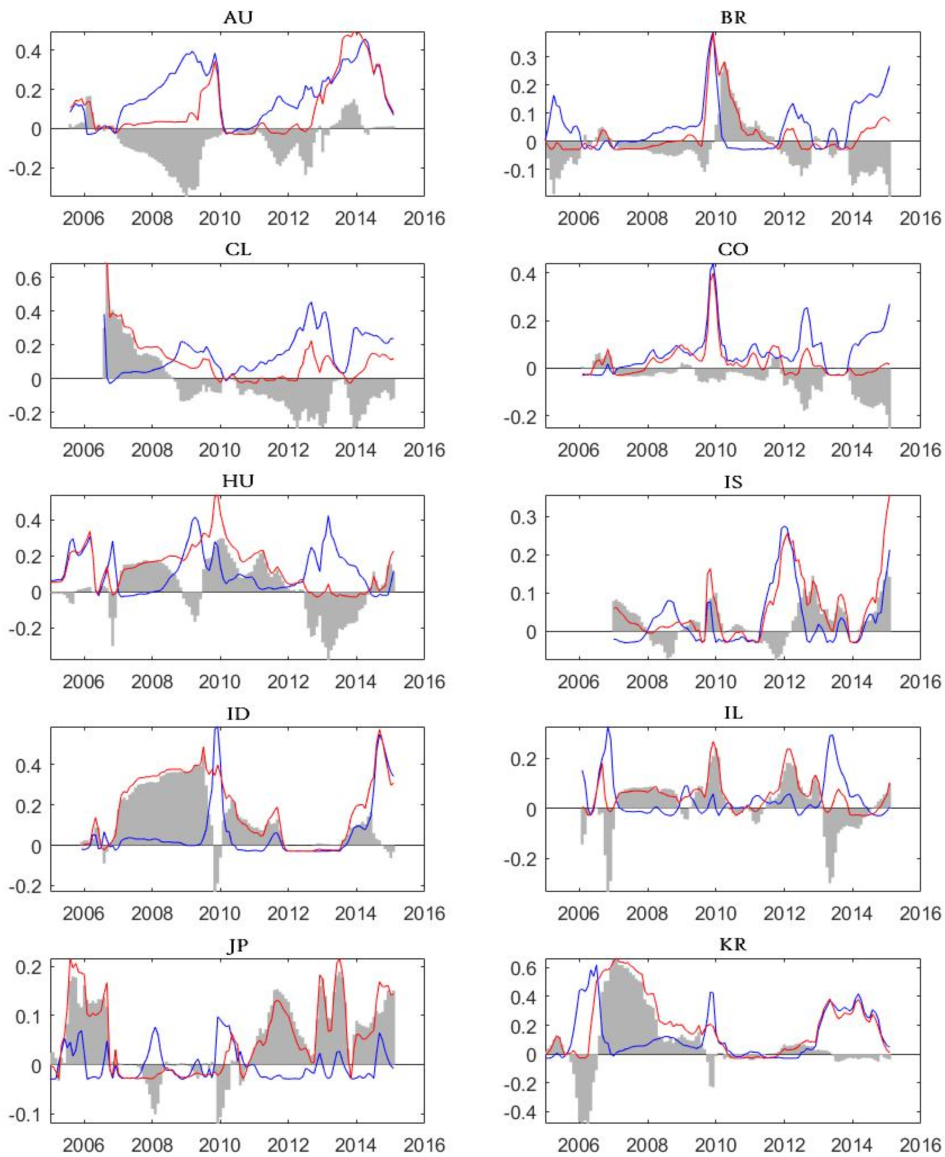
(%)



Note: 1. This figure shows the filtered probability in state 1 from Markov-Switching model estimation: $\Delta s_{t+3} = \beta_{0,\xi_t} + \beta_{1,\xi_t} L(D_t) + \beta_{2,\xi_t} S(D_t) + \epsilon_{t+3},\ where\ \epsilon_{t+3}\ i.i.d.N(0, \sigma^2_{\xi_t})$. State 1 represents a "high volatility"state. 2. Dark shades show the months when the VIX index is greater than 25 and light shades show the months when the VIX index is between 20 and 25.

Figure 2.4: Comparing the Ad.R^2 over Rolling Windows: Risk Level or Proximity?

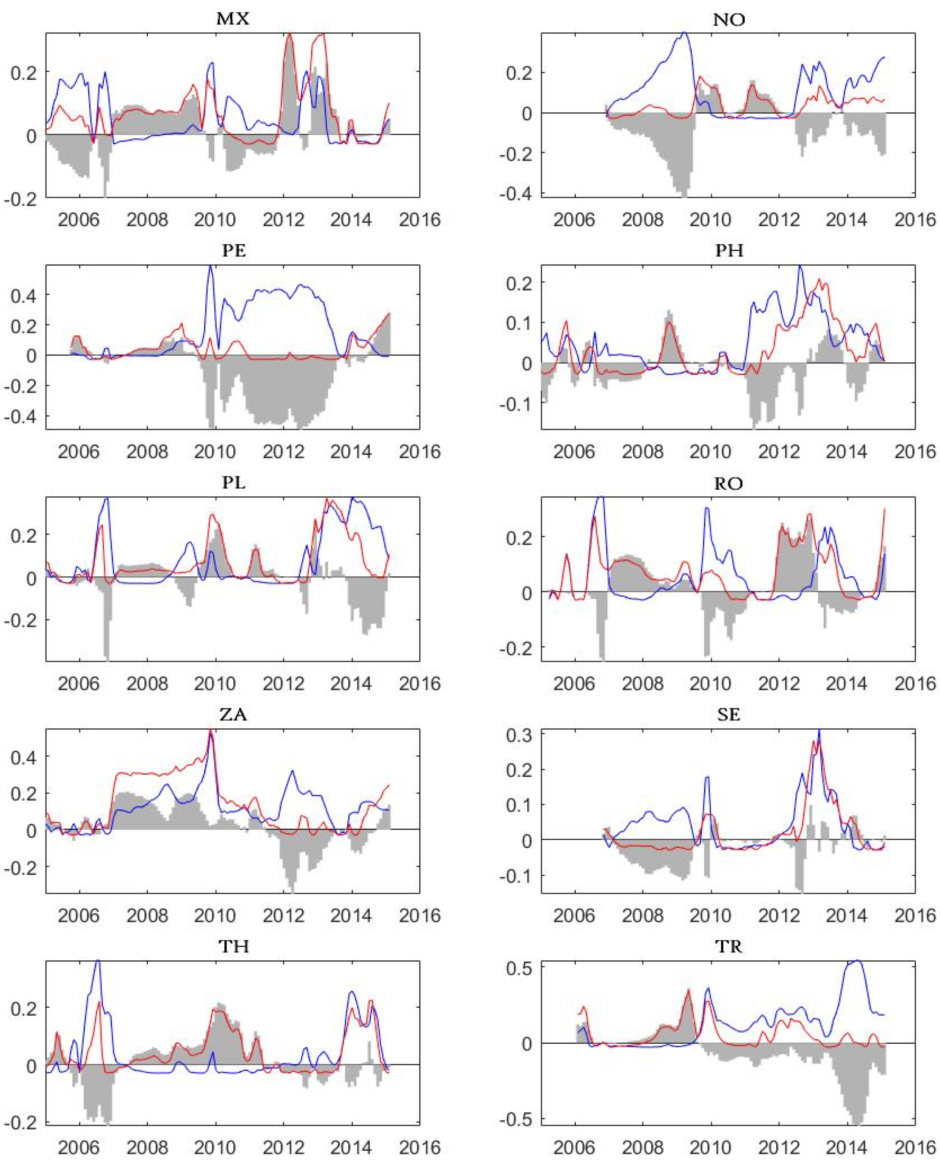
(Adjusted R^2)



Note: 1. This figure compares the goodness of fit measures (*adjusted* R^2) over 36-month rolling windows from two models: 1) Level-only (Risk-level only) model: $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \epsilon_{t+3}$; 2) Slope-only (Risk-proximity only) model: $\Delta s_{t+3} = \beta_0 + \beta_1 S(D_t) + \epsilon_{t+3}$. 2. Blue line represents adjusted $R^2 s$ from the level-only model and Red line shows adjusted $R^2 s$ from the slope-only model. Gray bar shows the difference in adjusted $R^2 s$ from the two models defined as "adjusted R^2 from the slope-only model − adjusted R^2 from the level-only model". If the difference is positive, the slope-model explains relatively more than level-only model for that regression window. 3. X-axis represents the midpoint of each window.

(Continued)

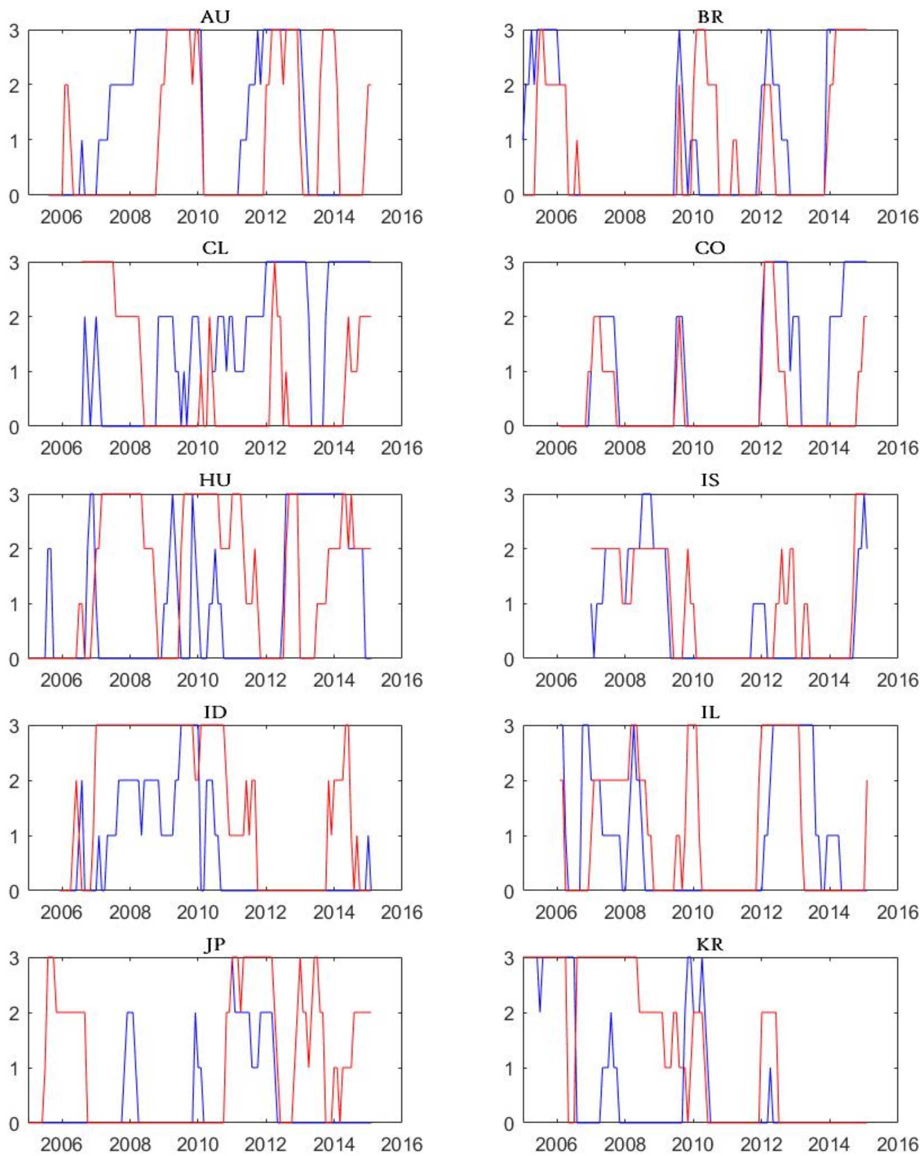
(Adjusted R^2)



Note: 1. This figure compares the goodness of fit measures (*adjusted* R^2) over 36-month rolling windows from two models: 1) Level-only (Risk-level only) model: $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \epsilon_{t+3}$; 2) Slope-only (Risk-proximity only) model: $\Delta s_{t+3} = \beta_0 + \beta_1 S(D_t) + \epsilon_{t+3}$. 2. Blue line represents adjusted $R^2 s$ from the level-only model and Red line shows adjusted $R^2 s$ from the slope-only model. Gray bar shows the difference in adjusted $R^2 s$ from the two models defined as "adjusted R^2 from the slope-only model − adjusted R^2 from the level-only model". If the difference is positive, the slope-model explains relatively more than level-only model for that regression window. 3. X-axis represents the midpoint of each window.

Figure 2.5: Comparing the $P - values$ over Rolling Windows: Risk Level or Proximity?

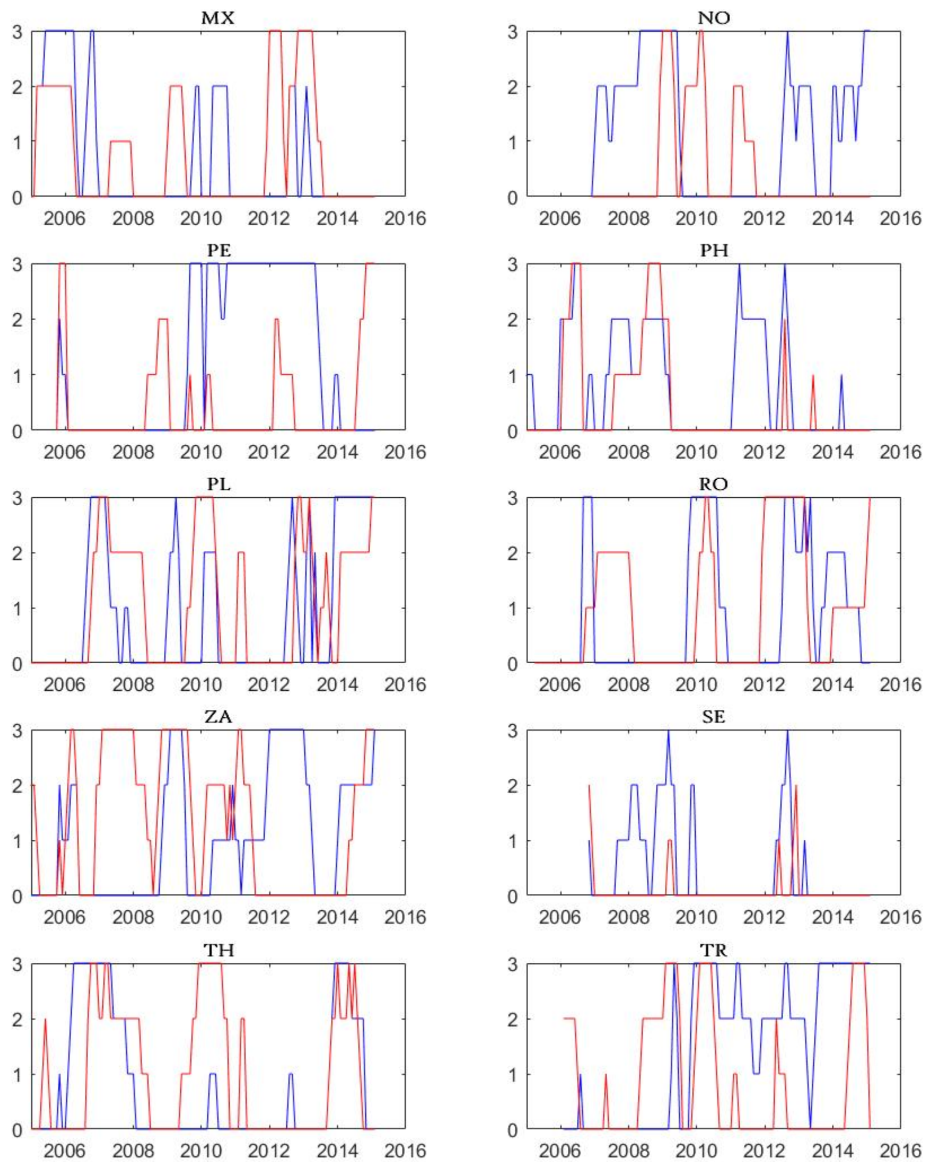
($P - values$)



Note: 1. This figure compares the statistical significance ($P - value$) over 36-month rolling windows for the level (risk-level) and slope (risk-proximity) factors. From the regression of $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \beta_1 S(D_t) + \epsilon_{t+3}$, I obtain $P - values$ for the level and slope factors. If $P - value \leq 0.01$, I give "3". If $0.01 < P - value \leq 0.05$, I give "2". If $0.05 < P - value \leq 0.1$, I give "1". Otherwise, "0". 2. Blue line for the level factor and Red line for the slope factor. 3. X-axis represents the midpoint of each window.

(Continued)

$(P - values)$



Note: 1. This figure compares the statistical significance $(P - value)$ over 36-month rolling windows for the level (risk-level) and slope (risk-proximity) factors. From the regression of $\Delta s_{t+3} = \beta_0 + \beta_1 L(D_t) + \beta_1 S(D_t) + \epsilon_{t+3}$, I obtain $P - values$ for the level and slope factors. If $P - value \leq 0.01$, I give "3". If $0.01 < P - value \leq 0.05$, I give "2". If $0.05 < P - value \leq 0.1$, I give "1". Otherwise, "0". 2. Blue line for the level factor and Red line for the slope factor. 3. X-axis represents the midpoint of each window.

Chapter 3

GLOBAL FINANCIAL CRISIS AND THE EXCHANGE RATE-YIELD CURVE CONNECTION[1]

3.1 Introduction

From the theoretical perspective, the literature has long emphasized how market expectations over future macroeconomic conditions or policy stance can affect exchange rate behavior; the same is true for the perceived riskiness over future horizons. Empirically, however, assessments of the relative importance of risk versus expectations in determining actual exchange rates have been much less conclusive. This chapter aims to evaluate the roles of these two channels in explaining recent currency movements, especially in light of the changing global economic conditions over the 2007-2008 Global Financial Crisis period. Since neither market expectations nor perceived future risk can be observed directly, we utilize insights from the exchange-yield curves literature and propose theoretically-motivated measures for them. (See Bekaert et al. (2007) and Chen and Tsang (2013), among others, for the connection between exchange rates and yield curves.) Specifically, using well-established models for the term structure of interest rates, we decompose information in the yield curves into expected future rates (based on expected future macroeconomic conditions) and term risk premiums. We then explore how their respective influence on exchange rate movements has changed with the onset of the crisis and the unconventional macro policy responses, such as Zero Lower Bound (ZLB), Quantitative Easing (QE) and Operation Twist.

Conceptually, we view this chapter as bridging two mostly separate strands of literature. First, on the international macroeconomics side: against decades of negative findings in testing exchange rate models, work by Engel et al. (2007), Molodtsova and Papell (2009),

[1] This chapter is based on co-authored work with Yu-chin Chen and Kwok Ping Tsang.

and Wang and Wu (2012) show that models in which monetary policy follows an explicit Taylor (1993) interest rate rule deliver improved empirical performance, both in in-sample fits and in out-of-sample forecasts. These studies emphasize the importance of expectations about future macroeconomic dynamics, and argue that the nominal exchange rate should be viewed as an asset price embodying the net present value of its expected future fundamentals.[2] While generally recognizing the presence of risk, this literature largely ignores risk in empirical testing and renders it an unobservable.[3] On the finance side, research shows that systematic sources of financial risk, as captured by latent factors, drive excess currency returns both across currency portfolios and over time. (See Lustig et al. (2011), Menkhoff et al. (2012), Farhi et al. (2009) and references therein for the connection between risk factors and currency portfolio returns.) These studies firmly establish the role of risk but are silent on the role of macroeconomic conditions, including monetary policy actions, in determining the exchange rate. They thus fall short on capturing the potential feedback between macroeconomic forces, expectations formation, and perceived risk in exchange rate dynamics. By incorporating information in the yield curves, we can examine these relationships and see how exchange rates price in and react to Taylor-rule macro fundamentals as well as expectations and risk that are also priced in the yield curves across countries.

We first present the NPV representation of exchange rate derived in **Chapter 1** where central banks follow a Taylor-rule interest rate rule, while explicitly considering the ZLB constraint. Since cross-country yield differences of all maturities can proxy both the expected future relative macro fundamentals and the difference in the underlying risks across countries as demonstrated in the previous chapter, they provide a theoretically-sensible and easy way to decompose and test separately the importance of expectations and systematic risk in

[2]Since the Taylor-rule fundamentals - measures of inflation and output gap - affect expectations about future monetary policy actions, changes in these variables induce the nominal exchange rate responses.

[3]Engel et al. (2007), for example, establish a link between exchange rates and fundamentals in a present value framework. After explicitly recognizing the possibility that risk premiums may be important in explaining exchange rates, they "do not explore that avenue in this paper, but treat it as an unobserved fundamental."

driving currency behavior. The usefulness of the yield curves in capturing expectations has become even greater after 2008 when the central banks in most of the advanced economies aggressively cut their policy rates and hit the ZLB to deal with the Great Recession. As the monetary policy tools have been switched from conventional policy-rate-cuts to unconventional policies which directly affect the medium to long yields, more attention has been paid to the yields with longer maturities when trying to read the market expectations about future macroeconomic conditions. Recently there has been burgeoning empirical literature on the effect of unconventional monetary policies on the bond yields. Cook and Devereux (2016), Gertler and Karadi (2015) find that the "forward guidance" is an important policy instrument as well as short rates which can affect the medium to long yields. Large Scale Asset Purchases like QE and Operation Twist have also been proven to effectively lower the yields with medium to longer maturities (Wright, 2012; Swanson and Williams, 2014). Given the observation that the short rates react to and thus reflect the macro fundamentals in normal times, longer yields under the ZLB take over the role of describing central banks' assessment of macroeconomic developments. A change in the informational contents of yields at different maturities, therefore, motivates us to use the entire yield curve, instead of a single short rate, to proxy the expectations about future macroeconomic conditions.

For our empirical analyses, we look at monthly exchange rate changes for eight country pairs - Australia, Canada, Denmark, Japan, New Zealand, Sweden, Switzerland, and the UK relative to the US - over the period from January 1995 to March 2016.[4] For each country pair, we use the zero-coupon yield data with maturities ranging from three months to ten years. To summarize the cross-country yields in a parsimonious way, we extract three Nelson-Siegel (Nelson and Siegel, 1987) factors. These three latent risk factors, which we refer to as the *relative level*, *relative slope*, and *relative curvature*, capture movements at the long, short, and medium part of the relative yield curves between the two countries. The Nelson-Siegel factors are well known to provide excellent empirical fit for the yield curves, providing a

[4]We mainly present results based on the dollar cross rates, though the qualitative conclusions extend to other pair-wise combinations of currencies as presented in Appendix.

succinct summary of both expectations about future macroeconomic dynamics as well as the systematic sources of risk that may underlie the pricing of different financial assets. Taking into account the possibility of structural breaks, we first confirm results established in Chen and Tsang (2013) that these yield curve factors indeed have robust explanatory power for subsequent exchange rate behavior. We then proceed to examine the specific role of risk versus expectations in these results.

In order to decompose the yield curves into expectations and risk, we employ four alternative methods based on different concepts of terms structure modeling that are well-known in the literature. These include the Nelson-Siegel latent factor model, the Nelson-Siegel latent factor model which allows interaction with macroeconomic fundamentals as discussed in Diebold and Li (2006), Ang and Piazzesi (2003)'s discrete-time affine Gaussian term structure model and also the Cochrane and Piazzesi (2005) approach.[5] Based on these alternative and admittedly all incomplete measures of expectations and risk, we demonstrate that both expectations and risk contained in the yield curves act as important determinants for quarterly exchange rate changes, providing empirical support for the present value models of exchange rate determination. This also provides support for the view that the same set of country-specific time-varying latent risks is priced into both the bond and the currency markets. We view this result as a clear indication that neither the macro nor the finance (risk) side of exchange rate determination should be ignored. Given the above findings, we investigate which of expectations and risk play a more important role in determining exchange rate changes. To illustrate the changing relevance of various exchange rate determinants, we use a variety of testing procedures including the Wald test, rolling-regressions, as well as Hodrick (1992)'s partial R^2 methods. Overall, the evidence points to a time-varying relationship among the yield curves, macro fundamentals, and subsequent exchange rate dynamics that

[5]As an example for the Nelson-Siegel model augmented with macro variables, we use an estimated VAR that allows for dynamic interactions between macro fundamentals and the yield curve factors, to construct measures of *expected yields*, which is average of expected short yields, for different maturities for each country. We then take the difference between the fitted yields from the model and the *expected yields* to separate out the time-varying *bond term premiums*. The *relative expected yields* and the *relative term premiums* are defined as the difference in *expected yields* and *term premiums* between each country-pair.

is broadly consistent with unconventional monetary policies and altered risk tolerance in investors during the crisis period. Finally, we propose the joint macro-finance model which can incorporate both Taylor-rule type and unconventional monetary policies.

To summarize, our main results are as follows: 1) empirical exchange rate equations based on only macro-fundamentals or only latent risk factors can miss out on the two crucial elements that drive currency dynamics: risk and expectations; 2) decomposing the yield curves into expectations for future macro dynamics versus term premiums, we show that both are important and can explain a considerable amount of the variations in subsequent quarterly exchange rate changes; 3) expectations play a stronger and more consistent role over the full sample period, while risk measures pick up their significance Post-Break; 4) macro variables offer little marginal explanatory power beyond what's in the yield-curve-based expectations and risk measures Pre-Break, but as the yield curves can no longer encompassing the same information about future macro dynamics under the ZLB, macro fundamentals themselves become important in explaining exchange rate Post-Break. These findings suggest that exchange rate behavior should be jointly modeled with the yield curve and macro fundamentals, supporting an open-economy extension of the joint macro-finance framework used to model the yield curve.

3.2 Theoretical Framework

3.2.1 Exchange Rate Determination Pre- and Post-ZLB

Assuming a standard Taylor-rule type monetary policy rule, the NPV equation of exchange rate presented in **Chapter 1** dictates that the nominal exchange rate depends on the expected future path of macro-fundamentals and time-varying risks. In this chapter, we modify the monetary policy rule by explicitly considering the Zero Lower Bound (ZLB) constraint. If the Taylor-rule implied rate is higher than the lower bound, they set the policy rates reacting to inflation and output (or unemployment) deviations from their target levels (and the home central bank reacts to the real exchange rate in addition). On the other hand, if the Taylor-rule implied rate is lower than the lower bound, the ZLB constraint binds. The monetary policy rules can be expressed as

$$\begin{aligned} i_t &= \max\{\underline{i},\ \mu_t + \beta_y \widetilde{y}_t + \beta_\pi(\pi_t - \bar{\pi}_t) + \delta q_t + u_t\} \\ i_t^* &= \max\{\underline{i}^*,\ \mu_t^* + \beta_y^* \widetilde{y}_t^* + \beta_\pi^*(\pi_t^* - \bar{\pi}_t^*) + u_t^*\} \end{aligned} \tag{3.1}$$

where $\underline{i}$ is the lower bound rate.

We consider symmetric monetary policy rules in home and foreign countries: both countries implement the Taylor-rule based rates, or both countries set the rates at or near zero.[6] Then, approximating the monetary policy rules, eq.(3.1), with $m = 1$, we can specify the home relative to foreign monetary policy rules as:

$$i_t^1 - i_t^{1,*} = \max\{\underline{i} - \underline{i}^*,\ \beta f_t^R + \delta s_t\} \tag{3.2}$$

Next, applying the NPV equation from **Chapter 1** delivers the following equations: If home and foreign central banks follow the Taylor rule (ZLB constraint does not bind),

$$s_t = \underbrace{-\sum_{j=0}^{\infty}\left(\frac{1}{1+\delta}\right)^{j+1}\beta E_t f_{t+j}^R}_{\text{Expectations}} + \underbrace{\sum_{j=0}^{\infty}\left(\frac{1}{1+\delta}\right)^{j+1} E_t \rho_{t+j}^1}_{\text{Risk}} \tag{3.3}$$

If the central banks cut the policy rates to the lower bound and commit to keeping the rates for n-periods of time (ZLB constraint binds from t to $t+n$)[7],

$$s_t = \underbrace{-n(\underline{i} - \underline{i}^*) - \sum_{j=n+1}^{\infty}\left(\frac{1}{1+\delta}\right)^{j-n}\beta E_t f_{t+j}^R}_{\text{Expectations}} + \underbrace{\sum_{j=0}^{n} E_t \rho_{t+j}^1 + \sum_{j=n+1}^{\infty}\left(\frac{1}{1+\delta}\right)^{j-n} E_t \rho_{t+j}^1}_{\text{Risk}} \tag{3.4}$$

[6] We admit that there exists the case when one country implements the Taylor-rule rate and the other country hits the ZLB. But, the asymmetric cases do not make big difference for our analysis.

[7] e.g., The FOMC in late 2013 and early 2014 said that it would continue to keep the federal funds rate at the lower bound at least until the unemployment rate fell to 6.5% and inflation increased to 2%.

These formulations show that the exchange rate depends on both expected macro fundamentals (expectations) and the perceived risks over time (risk). Compared to standard NPV equation (3.3), the ZLB-NPV equation (3.4) does not incorporate the current and short-run expectations and puts more weights on risk in the exchange rate determination.

3.2.2 The Yield Curve: Proxy for Expectations and Risk Pre- and Post-ZLB

In **Chapter 1**, I show that the cross-country yield curves can be used as proxies for both expectations about future macro fundamentals and the time-varying risks. Recall that according to the expectation hypothesis (EH), the *relative yields* are the sum of *relative expected yields* and *relative term premiums.*

$$i_t^m - i_t^{m,*} \equiv \frac{1}{m}\sum_{j=0}^{m-1} E_t\left[i_{t+j}^1 - i_{t+j}^{1,*}\right] + (\theta_t^m - \theta_t^{m,*}) \tag{3.5}$$

Suppose that the lower bound is announced and expected to persist for n-periods. Then, the equation above should be modified to

$$i_t^m - i_t^{m,*} \equiv (\underline{i} - \underline{i}^*) + \frac{1}{m-n}\sum_{j=n+1}^{m-1} [E_t i_{t+j}^1 - E_t i_{t+j}^{1,*}] + (\theta_t^m - \theta_t^{m,*}) \tag{3.6}$$

Under the standard Taylor rule where the ZLB does not bind, the *relative expected yields* proxy the first summation on the right-hand side of eq.(3.3) since

$$E_t[i_{t+j}^1 - i_{t+j}^{1,*}] = \beta E_t f_{t+j}^R + \delta E_t s_{t+j}, \forall j \tag{3.7}$$

In the other case when the ZLB binds n-periods, the different formulation of *relative expected yields* exactly correspond to the first two summations of eq.(3.4) because

$$E_t[i_{t+j}^1 - i_{t+j}^{1,*}] = \begin{cases} \underline{i} - \underline{i}^* & \text{for } j \leq n \\ \beta E_t f_{t+j}^R + \delta E_t s_{t+j} & \text{for } j \geq n+1 \end{cases} \tag{3.8}$$

The *relative term premiums* also link to the currency risk premiums with and without the ZLB constraint by

$$\rho_t^m - \frac{1}{m}\sum_{j=0}^{m-1} E_t \rho_{t+j}^1 = \theta_t^m - \theta_t^{m,*} \tag{3.9}$$

For our empirical studies, it is essential to decompose the yield curves into the *expected yields* and *term premiums*, which are used as proxies for expectations and risk, respectively. Since there are alternative methods based on different concepts of term structure modeling and one is not necessarily considered as superior to the others, we adopt four major models in the literature to decompose the yield curves: 1) Nelson-Siegel latent factor model (hereafter, NS model); 2) Nelson-Siegel latent factor model which allows interaction with macro fundamentals as discussed in Diebold, Rudebusch, and Aruoba (2006) (hereafter, NSM model); 3) Ang and Piazzesi (2003)'s affine Gaussian term structure model (hereafter, Affine model); 4) Cochrane and Piazzesi (2005) approach (hereafter, CP model). We first estimate the models to construct measures of *expected yields*, which are averages of expected short yields, for different maturities for each country. We then take the difference between the fitted yields from the model and the *expected yields* to separate out the time-varying *bond term premiums* at each maturity. The *relative expected yields* and the *relative term premiums* are defined as the difference in *expected yields* and *term premiums* between each country-pair.[8]

3.2.3 Theoretical Predictions

Based on our theoretical framework presented in the previous subsection, we make three main predictions to be empirically tested as follows:

[8]For details, please refer to Appendix

Prediction 1: Both expectations and risk drive the currency movements.
The NPV representations, eq.(3.3) and (3.4), directly predict this. In Section 3.4.1, we explore whether expectations and risk, measured by *relative expected yields* and *relative term premiums* extracted from the cross-country yield curves, individually explains the subsequent quarterly exchange rate changes. We also compare the explanatory powers of the expectation-only, the risk-only models with the both-expectation-risk model.

Prediction 2: Relative role of expectations and risk differs pre- and post-ZLB.
From two different NPV equations pre- and post-ZLB, eq.(3.3) and (3.4), we infer that the expectations play a less role while risk plays a more role in the ZLB period, compared to pre-ZLB period. As our theoretical model links the expected macro-fundamentals to the exchange rate via the monetary-policy rule which sets the short rates based on the assessment of macroeconomic conditions, this mechanism successfully works in normal times pre-ZLB. However, this linkage loosens in the ZLB period because the lower-bounded short rates deviate from what the underlying economic conditions require them to be, even though unconventional monetary policies effectively lower the medium- to long-term interest rates under the ZLB and as such the yield curve partially delivers the information about how the central banks view the macroeconomic developments.[9] On the other hand, risk matters more post-ZLB. The lack of conventional rate-cut policy tool limits the effectiveness of policy responses, and thus raises the concerns of investors on persistent risk. Consequently, risk becomes more important in driving the cross-country investments and the currency movements. In **Section 3.4.2**, we test this prediction by using various methods including the Wald test, Hodrick (1992)'s partial R^2, and rolling regressions.

[9]Wright (2012), Swanson and Williams (2014) show that Large Scale Asset Purchases like QE and Operation Twist effectively lower the yields with medium to longer maturities under the ZLB. On the other hand, Wu and Xia (2016) compute the negative valued shadow rates which are found to reflect the macroeconomic condition well. Combining these findings, it can be said that the actual yield curves under the ZLB do still have information about expectations, but not completely.

Prediction 3: Relative role of yields and macro differs pre- and post-ZLB. Given that the link between macro-fundamentals and the yield curves weakens under the ZLB, we posit that the contemporary macroeconomic indicators become relevant determinants of the exchange rates post-ZLB, while they deliver little information in addition to the yield curves pre-ZLB. **Section 3.4.3** empirically investigates this changing relevance of the macroeconomic variables as determinants of the exchange rate.

3.3 Preliminary Empirics

This section takes a preliminary look at the data. Given that our theoretical model predicts different relationships between the exchange rate and its determinants pre-ZLB and post-ZLB, and the beginning of ZLB period broadly coincides with the onset of the Global Financial Crisis in 2008, we first describe the data with a potential break around 2008. We then regress the relative yield curve factors on the exchange rate changes with a structural break to confirm the usefulness of yield curves in explaining the currency movements (Chen and Tsang, 2013), and to check a potential change in the exchange rate - yield curve connection as the theory predicts.

3.3.1 Data Description

The main data we examine consists of monthly observations from January 1995 to March 2016 for Australia (AU), Canada (CA), Denmark (DK), Japan (JP), New Zealand (NZ), Sweden (SE), Switzerland (CH), the United Kingdom (UK) and the United States (US) of the following series: 1) exchange rate data: End-of-period monthly exchange rates are obtained from the FRED database. We use the logged exchange rate, measured as the per-dollar rates. Exchange rate change from t to $t+m$ is expressed as $\Delta s_{t+m} = s_{t+m} - s_t$ and annualized; 2) macroeconomic data: We obtain headline CPI and unemployment rate from the OECD main economic indicators.[10] Inflation rate is defined as 12-month percentage

[10]Unemployment rate for Switzerland is from Swiss Federal Statistical Office.

change of the CPI. Unemployment gap is obtained by detrending the unemployment rate using the Hodrick-Prescott filter; 3)yield data: zero-coupon bond yields include maturities of 3, 6, 12, 24, 36, 48, 60, 72, 84, 96, 108 and 120 months, where the yields are computed using the curved stripping method and obtained from Bloomberg.[11] The yields are from the last trading day of each month.

Table 3.1 report the summary statistics of exchange rates and macroeconomic variables. Considering potential structural breaks due to the Global Financial Crisis and the ZLB, the sample period is divided by the preliminary break date, May 2008.[12] For three-month exchange rate change Δs_{t+3}, we see that all currencies except the Japanese Yen have appreciated before the break, and all currencies except the Swiss Franc have depreciated after the break. This would be consistent with the idea that the US dollar (along the Swiss Franc) is commonly considered safe haven currencies. The volatility of exchange rate has been increased after the break except for Japan and Switzerland. Two macro variables we use are the relative unemployment gap and inflation rate of each of eight countries to those of the US.[13] Relative unemployment gaps have been higher before the break but lower after the break, implying more job loss in the US labor market. Australia has higher inflation rate, while Japan, Sweden and Switzerland have lower inflation rate, compared to the US. For Canada, Denmark, New Zealand and the UK, the relative inflation rate is lower before the break but higher after the break.

Figure 3.1 describes the cross-country yields by showing the sample average of *relative yields*, $i_t^{R,m} = i_t^m - i_t^{m,*}$, before and after the break, May 2008. The interpretation of the *relative yield curves* extends readily from their single-country counterparts. We see that the level of the *relative yields*, which is the gap between the entire home yield curve and the

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[11] Please refer to Kushnir (2009) for details on the construction of the data.

[12] The break date in this section is chosen considering that the structural break test in the following sections reports May 2008 as a most common break date across countries.

[13] We use unemployment gap to proxy the output gap since monthly data is easily obtained. Monthly industrial production (IP) index is also available, but the choice of variables for the output gap does not make a big difference for our empirical analysis.

US one, is positive (except Japan and Switzerland) before the break. After the break, we observe a dramatic change in the slope of the *relative yield curves*. For most of the countries (except New Zealand and the UK), the slope changes from negative to positive. Noting that the slope represents the short minus long yield, the positive slope means that the yield curve of home country is relatively flat compared to that of the US.

We construct the *relative expected yields* and *relative term premiums* by decomposing the cross-country yield curves by four major methods. **Figure 3.2** and **3.3** visualize the sample averages of *relative expected yields* and *relative term premiums* estimated by four models.[14] We note that the sum of *relative expected yields* and *relative term premiums* of each maturity is equal to *relative yields* of the same maturity. The *relative term premiums* are very small close to zero in the shortest maturity and departs from zero over longer maturities, as it can be interpreted as the extra compensation for holding longer maturity bond. Similar to the *relative yield curves*, we observe that the expectation and perceived riskiness of various sovereign bonds at different horizons shifted significantly in May 2008.

3.3.2 Structural Break in Exchange Rate-Yield Curve Connection

We confirm findings in Chen and Tsang (2013) that relative Nelson-Siegel yield curve factors have predictive power for subsequent quarterly exchange rate changes.[15] Compared to the previous work, we cover a larger set of country-pairs, and the data sample covers the recent financial crisis. As such, we put an emphasis on possible structural breaks in the yield curve-exchange rate relation. For each of the eight-country pairs, we run the following

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[14]These figures selectively show the decomposed results from four models: one model for two country-pairs. Full data descriptions are provided in Appendix along with the model estimation methods.

[15]The Nelson and Siegel (1987) exponential components framework are used to distill the entire relative yield curves, period-by-period, into a three relative factors. Assuming symmetry and exploiting the linearity in the factor-loadings, we extract three factors of *relative level, slope, curvature* (L_t^R, S_t^R, C_t^R) as follows: $i_t^{R,m} = i_t^m - i_t^{m,*} = L_t^R + S_t^R \left(\frac{1-\exp(-\lambda m)}{\lambda m}\right) + C_t^R \left(\frac{1-\exp(-\lambda m)}{\lambda m} - \exp(-\lambda m)\right) + \epsilon_t^m$. The parameter λ is set to 0.0609, in accordance with the literature such as Diebold and Li (2006).

regressions and report the results in **Table 3.2**:

$$\Delta s_{t+3} = \beta_0 + \beta_1 L_t^R + \beta_2 S_t^R + \beta_3 C_t^R + \epsilon_{t+3} \tag{3.10}$$

To address possible parameter instabilities, we test for endogenous structural breaks in the regression. These tables report results based on the Quants-Andrews unknown breakpoint test (Andrews, 1993; Hansen, 1997) with 15% trimming and 5% significance level. The break dates are mostly around 2008 when the Global Financial Crisis was triggered and the policy rates hit the ZLB.[16] We first find that the predictive power of the relative yield curve is apparent. Contrary to results typical in the empirical exchange rate literature which tend to find essentially no explanatory power, especially at the monthly or quarterly frequency, we see that the regressions here can produce adjusted R^2 at least 11% up to 31%. Testing the joint significance of relative factors in explaining currency behavior, we also see that the p-values from the Wald test are all below 1%, indicating strongly rejections of the hypothesis that yield curves contain no information about subsequent currency behavior. These results establish the predictive power of the relative factors, and justify the use of yield curves as candidates for capturing both expectations and risk in explaining the exchange rate dynamics.[17] The evidence on the structural break around 2008 is clear as the coefficient estimates change their signs and magnitudes depending on the state of the economy, potentially lending support on our theoretical predictions about changing relations of the exchange rate to expectations, risk and macro-fundamentals.

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[16] We also tested for the structural breaks using different methods such as Bai and Perron (2003) multiple break test. In all cases, our break dates are detected as one of the multiple breaks.

[17] Admitting that high adjusted R^2 may be due to multi-month ahead prediction with overlapping regressors of high persistence, here and in the following sections, we check the robustness with non-overlapping regressions (i.e., quarterly prediction with quarterly data). We still find fairly high goodness of fit measures. All the results are provided in Appendix.

3.4 The Relative Role of Expectations, Risk, and Macro

We empirically test our three theoretical predictions: 1) Both expectations and risk drive the currency movements; 2) The relative role of expectations and risk differs pre- and post-ZLB. The expectations play a less role, while risk plays a more role under the ZLB, compared to without the ZLB; 3) The relative role of yields and macro differs pre- and post-ZLB. The contemporary macroeconomic indicators become relevant determinants of the exchange rates post-ZLB, while they deliver little information in addition to the yield curves pre-ZLB.

3.4.1 Expectations and Risk in Explaining the Currency Movements

We examine if either expectations (measured by the *relative expected yields*) or risk (measured by *relative term premiums*) can individually explain subsequent quarterly exchange rate changes. Following the idea of three Nelson-Siegel framework, we summarize the entire variables at all maturities with three factors : *level, slope* and *curvature.*[18] We also compare the explanatory powers of two individual models to that of the model with both expectations and risk. In particular, we run the following regressions with a structural break for each currency:[19]

$$\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \epsilon_{t+3} \quad (3.11)$$

$$\Delta s_{t+3} = \beta_0 + \beta_1 S(\theta_t^R) + \beta_2 C(\theta_t^R) + \epsilon_{t+3} \quad (3.12)$$

$$\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \beta_4 S(\theta_t^R) + \beta_5 C(\theta_t^R) + \epsilon_{t+3} \quad (3.13)$$

[18]The *level*, *slope* and *curvature* factor of the *relative expected yields* are constructed as follows: $L(E_t i_t^R) = E_t i_t^{R,120}$, $S(E_t i_t^R) = E_t i_t^{R,3} - E_t i_t^{R,120}$, $C(E_t i_t^R) = 2E_t i_t^{R,24} - (E_t i_t^{R,3} + E_t i_t^{R,120})$. The *level*, *slope* and *curvature* factor of the *relative term premiums* are defined as follows: $L(\theta_t^R) = \theta_t^{R,120}$, $S(\theta_t^R) = \theta_t^{R,3} - \theta_t^{R,120}$, $C(\theta_t^R) = 2\theta_t^{R,24} - (\theta_t^{R,3} + \theta_t^{R,120})$. The correlation between *level* and *slope* factor of the *relative term premium* is close to negative unity since the *term premiums* at the shortest maturity is almost zero. So, the *level* factor is excluded from regressors to avoid the multi-collinearity problem.

[19]We test for endogenous structural breaks in the regression, based on Quants-Andrews unknown breakpoint test (with 15% trimming and 5% significance level). After identifying the break, structural break dummy variables for each sub-period are incorporated into the regression.

Table 3.3 presents the p-values from the joint Wald test and adjusted R^2 for the expectation-only model (3.11), the risk-only model (3.12), and the both-expectation-risk model (3.13). As shown in the table, the p-values from the joint Wald test are all below 1% for the expectation-only model and mostly below 10% for the risk-only model. The hypothesis that the *relative expected yields*, equivalently the expectations, have no information about exchange rate changes is strongly rejected. The *relative term premiums*, equivalently the risk, are strong and robust determinants of currency movements, supporting the view that differential risks in the relative bond markets are priced into the corresponding currency values. The goodness of fit measures (adjusted R^2) say that the expectations can explain up to 34%, while the risk can explain up to 30% of the variations in exchange rate changes even though adjusted R^2's are low for some countries depending on models. Comparing the explanatory power of *relative term premium* factors to *relative expected yield* factors, the latter explains more over the full sample period. We observe that the model with both expectations and risk can jointly explain up to 34%.[20] The explanatory power improves considerably over only Nelson-Siegel factor model discussed in **Section 3.3.2**. The adjusted R^2's increase by 3%p to 13%p. The goodness of fit measures also improves over the expectation-only and the risk-only models. These results strongly support our idea that both expectations and risk are important determinants of the currency movements and neither should not be ignored.[21]

Table 3.4 shows the coefficients from the model with both expectations and risk.[22] The signs and magnitudes of coefficients change considerably before and after the breaks. Broadly speaking, higher factors tend to lead to subsequent appreciation of the currency or relatively small depreciation in the "Pre-Break" period, but result in depreciation of the currency or

[20]Concerned about multi-collinearity problem due to use of the same underlying factors, we perform the Belsley collinearity test and find no severe collinearity.

[21]Note that this chapter does not evaluate which term structure model is better in decomposing the yield curves because each model has its advantages and caveats. Rather, the focus of our research is to decompose the yield curves into expectations and risk to show both contribute to explaining currency behavior, no matter which term structure model is employed.

[22]We do not report the coefficients on *relative expected yield* factors or on *relative term premium* factors. Since we argue that both expectation and risk matter for currency movements, regression with only one group of factors may suffer from the omitted variable problem and thus the interpretation may be misleading.

relatively small appreciation during the "Post-Break". Although we do not explicitly test for any specific macroeconomic models, our results nevertheless have intuitive economic interpretations. As implied by our NPV representations, eq.(3.3) and (3.4), there exist two channels that determines the exchange rate dynamics: expectations and risk channel. Expectation channel is well-explained by the monetary-policy rule. Suppose the home country is already in an inflationary gap and its expected inflation rate in the long-term is higher than that in the US. Its central bank will raise the policy rate, resulting in appreciation of its currency. [23] On the other hand, suppose the home country's economy is in the recession and even more downturn is expected (difference in the expected output gaps is negative). Then, the central bank will lower its policy rate and exchange rate rises.[24] Since the *relative level* of expected yields reflects the expected inflation, higher level before the crisis results in its currency appreciation (or less depreciation). During the crisis, as higher *relative slope* of expected yields or equivalently flatter yield curve implies deeper recession or slower recovery, the home country experiences depreciation of its currency.

Turning to risk channel, we refer to recent empirical evidence about the carry-trade strategy.[25] This line of research says that when the volatility of exchange rate is low, the currency with higher risk appreciates as investors require compensation for holding risky currency. However, under high volatility, abrupt withdrawal of investment in favor of safe haven currencies causes loss to the currency with higher risk. The yield curve factors contain information about two types of risk. One type of risk is associated with the expected path of macro fundamentals. If the home country is expected to develop higher inflation and lower

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[23]If $E_t[(\pi_{t+j} - \bar{\pi}_{t+j}) - (\pi^*_{t+j} - \bar{\pi}^*_{t+j})] > 0$ for large j, $s_t < 0$ according the NPV equations since $\beta_\pi > 0$. Similarly, $s_{t+3} < 0$ as well, but $s_{t+3} < s_t$ due to smaller discount factors($(\frac{1}{1+\delta})^{j+1} > (\frac{1}{1+\delta})^{j-2}$). As a results, the exchange rate appreciates over subsequent quarter.

[24]The NPV equations with $\beta_y > 0$ imply that if $E_t[\widetilde{y}_{t+j} - \widetilde{y}^*_{t+j}] < 0$ for large j, $s_{t+3} > s_t > 0$.

[25]The carry-trade is a strategy under which investors take long positions on high-yield currency and short positions on low-interest rate currency. Lustig and Verdelhan (2007) find that the portfolios constructed by the carry-trade strategy yield high returns. Clarida et al. (2009) show that returns to the carry trade depend on the volatility of exchange rate. That is, when the volatility is low, the carry-trade gives gain. And when the volatility is high, it causes loss.

economic growth than the US, its currency is less desirable and thus risky.[26] The expectation risk can be captured by higher *relative level* and *slope* of expected yields. The other type of risk is about uncertainty in the financial markets. Considering that the relatively higher location and humped-shape of yield curve reflect the uncertainty in the long-term and the near-term respectively, higher *relative slope* and *curvature* of term premiums delivers the information that the financial risk in the home country is relatively high, compared to that in the US. As both types of risk are embedded in the factors, the regression result shows that the currency with higher risk appreciates before the crisis when the volatility is low, and depreciates during the crisis when the volatility is high. Therefore, combining the expectation and risk channels together, the sign-switching property of coefficients from our regression can be intuitively established.

3.4.2 The Relative Role of Expectations vs. Risk

Our theoretical model predicts that the expectations play a less role while risk plays a more role in the ZLB period. The change in the monetary policy from conventional policy-rate-cuts to unconventional ones has altered the transmitting channels between yield curves and exchange rate changes. First, the expectation channel has become weak. As implied by eq.(3.4), the ZLB constraint makes short-term expected macro fundamentals less connected to and longer-term fundamentals more related to the exchange rate. Although unconventional policies are designed to affect the medium- to long-term interest rates and expectations, zero bound on short-term interest rate prevents the yield curve from fully reflecting the market expectations in the economy. Intuitively, even though home country is expected to undergo severe economic downturn, the central bank cannot physically cut the rate further. As a result, currency depreciation from the expected fall in short rate in the

[26]As implied by eq.(3.3) and (3.4), the currency appreciates over time when the expected inflation and output gap is relatively high (when marginal utility of consumption is relatively low), while it depreciates when the economy is expected to grow slowly and fall into recession (when marginal utility is high). As the covariance between the currency return and marginal utility of consumption is negative, the home currency is a bad hedge for the associated risk.

aftermath of the crisis can hardly be predicted.

On the other hand, risk channel has become strong. Again from the same NPV representation during the ZLB period, we see the high weights on the short-term risk in determining the exchange rate. The lack of expectation channel in the short-run is replaced by risk channel as the investors are suspicious about the effectiveness of the policy responses. Another explanation based on the change in the market participants' attitude toward underlying risk can support the enlarged role of risk factors. Investors have become more aware of and sensitive to risk after they experience the rare but bad events such as the collapse of Lehman Brothers. They excessively react to risk-related news after mid-2008 and for a while. This explanation is compatible with recent finance literature. Lettau et al. (2014) and Dobrynskaya (2014) highlight that downside market risk explains currency returns more than upside market risk. Risk becomes more relevant in explaining the currency behavior during the market downturn. There are different reasons why investors may be more reactive to losses than to gains.[27] Among them, "loss aversion" by Barberis et al. (2001) can lend support on our findings. While investors are less loss averse after prior gains, they become more loss averse and more sensitive to additional setbacks after a prior loss. In our sample period, after observing huge loss during the Global Financial Crisis, investors care more about risk.

Empirically, we investigate which explains more of variation in currency movement. We first test for the joint significance of each group in sub-periods identified by the structural break from the regression (3.13), using the Wald statistics. We report the results from NSM model in **Table 3.5**.[28] Comparing the relative role of expectations and risk in explaining the subsequent quarterly exchange rate changes, the expectations seem to be more significant than risk in the full sample period. The null hypothesis that the expectations do not explain

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[27]See Dobrynskaya (2014) and references therein.

[28]Hereafter, we focus only on NSM model because NS, NSM and CP model share the similar decomposed expected yields and term premiums. NSM model, compared to Affine model, is conceptually better in expressing the expected yields as our theoretical model argues that current and future expected macro fundamentals are reflected in the future path of short-rates through monetary policy. Results from other models can be provided upon request.

exchange rate changes ("No expectations?") is rejected at 1% significance level for all the currencies, and the null that the risks have no explanatory power ("No risk?") is rejected for four countries. Turing to the sub-sample periods divided by the break dates, we can reject both hypotheses of "No expectations?" and "No risk?" for more countries in "Post-Break" period than in "Pre-Break" period. Both the expectation and risk channel become stronger after the break, but the difference in statistical significance is more apparent for the risk channel. Risk factors matter for only two countries before the break at 10% significance level, while they become relevant except for Canada, Denmark, and New Zealand after the break.[29] These exceptions are broadly consistent with our theoretical explanation above because the policy rate has not fallen below 2.5% in New Zealand, and has been 0.25% in Canada for relatively short period of time from April 2009 to May 2010.[30] And since Denmark, like other European advanced countries, cut the policy rates near zero only after 2012, the role of risk might be averaged out during the Post-Break period. The results from the regression with multiple structural breaks clearly show that risk factors have become relevant after the second break around 2011 - 2012 in these countries.[31]

As a next step, we try to show quantitatively the portion explained by each group: expectation and risk. We first construct a six-variable VAR(1), which consists of one-month exchange rate change, three factors from *relative expected yields* and two factors from *relative term premiums*. Following Hodrick (1992), we calculate the partial R^2 for each variable for explaining exchange change at various horizons.[32] Though the variable that enters the VAR system is the one-month exchange rate change, we can iterate forward the VAR to calculate the explanatory power of each variable for exchange rate change of longer horizons. In order

[29]We also check the robustness with different denominating currencies including Japanese Yen and Swiss Franc and find the similar patterns. See Appendix for the results.

[30]Australia also did not hit the ZLB, but shows the significant role of risk after the break. However, we argue that the change in monetary policy is one of the most plausible explanations and admit other reasons such as the market's changing risk attitude.

[31]See Appendix for the results.

[32]See Appendix for a detailed discussion on the method.

to compare the explained portion by each group in sub-periods, we divide the sample period into two by the break date around 2008. **Figure 3.4** shows the results for explaining exchange rate change at 1, 3, 6 and 12-month in the future. The portion explained by expectations in “Post-Break” is similar to or lower than in “Pre-Break”. However, the portion explained by risk in “Post-Break” is higher than in “Pre-Break” for five out of eight countries, and even higher than the portion explained by expectations for two countries. Similar to the Wald test results, exceptions are Australia, Canada, and New Zealand which have not or little experienced the ZLB. The bottom line, as seen from the Wald test and the partial R^2 analysis, is that the expectation channel becomes weak while risk channel pulls its weight during the ZLB period.

We further investigate the relative role of expectations and risk from a different angle. **Figure 3.5** shows the adjusted R^2’s with 24-month rolling window from the expectation-only and the risk-only models.[33] Again, the goodness of fit measures (adjusted R^2) of the expectation-only model is higher than that of the risk-only model in most of the time, while the risk-only model catch up during the “Post-Break” period except Japan and Switzerland.

3.4.3 The Relative Role of Macro-Fundamentals

We argue that the yield curves fall short of fully representing the expectations about macroeconomic conditions during the ZLB period. We further investigate this structural change in the informational content embodied in the yield curves and its relation to exchange rate dynamics by examining the role of macro fundamentals as additional explanatory variables. In order to test our idea, we regress both yield factors and Taylor-rule fundamentals on the currency movements as the following:

$$\Delta s_{t+3} = \beta_0 + \beta_1 L_t^R + \beta_2 S_t^R + \beta_3 C_t^R + \beta_4 u_t^R + \beta_5 \pi_t^R + \epsilon_{t+3} \tag{3.14}$$

[33] The factors are estimated by NSM model. The window size of 24 months is chosen to show the clear picture. The choice of window size does not affect the result.

Table 3.6 reveals that we can strongly reject the hypothesis that neither macro fundamentals nor yield factors can predict exchange rate movement next quarter for all the countries. Note that the explanatory power of these variables is pretty high: the adjusted R^2 can be up to 39%. The goodness of fit measures increase by up to 10%p compared to the yield-only model, eq.(3.10), and by up to 6%p compared to the both-expectation-risk model, eq.(3.13). Contemporaneous macroeconomic conditions explain a part of subsequent currency behavior which is not explained by the expectation and risk extracted from the cross-country yield curves.

In order to investigate the role of macroeconomic variables in explaining the currency behavior, we conduct the Wald test. For the full sample period, the null hypothesis that the latent yield factors do not explain exchange rate changes ("No Yields?") is strongly rejected. The null hypothesis that the macro variables have no contribution ("No Macro?") cannot be rejected only for Japan. Considering that Japan has experienced the prolonged deflationary gap, no explanatory power of macro variables is reasonable. The statistical significance depends on the sub-sample periods. Yield factors are consistently significant in explaining the exchange rate changes as the null "No Yields?"is rejected except one country before and after the break. On the other hand, the null "No Macro?"cannot be rejected for five countries before the break and for one country after the break. Under traditional Taylor-rule type monetary policy before the break, as the yield curve incorporate the information about macro fundamentals, the use of macro variables as additional regressors is redundant. However, during the Post-ZLB period, contemporaneous macro variables obtain the significance as the yield factors do not successfully represent the underlying macro condition.[34]

In addition to the Wald test, we further take several testing procedures including Hodrick's partial R^2 and rolling regression. We first quantitatively show the portion explained by yield factors and macro variables in **Figure 3.6**. The results are consistent with the Wald test. Macro fundamentals explain more variation in the 1- to 12-month exchange rate changes

[34]We check the robustness with different denominating currencies. See Appendix for the results.

after 2008 than before 2008 in seven countries. Surprisingly, the portion explained by macro variables even outweigh the portion explained by yield factors after 2008 in three countries. Rolling regression results in **Figure 3.7** graphically confirm these findings. Comparing the adjusted R^2's with 24-month rolling windows from the yield-only and the macro-only model, we see that the macro-only model gains its explanatory power after 2008.[35]

Our findings above confirm our theoretical predictions that there is a structural change in the yield-macro-exchange rate connection due to the ZLB, and suggest that the exchange rate dynamics should be jointly modeled with the yield curve and macro fundamentals to compromise both conventional and unconventional monetary policies. This is a natural extension of a macro-finance approach of Ang and Piazzesi (2003), Diebold, Rudebusch, and Aruoba (2006) to an international setting. **Figure 3.8** compares the fitted quarterly exchange rate changes from the joint model and from the yield-only model against the actual quarterly exchange rate changes. Consistently, the joint model outperforms the yield-only model in fitting the currency movement, especially after 2008, supporting our proposal of the joint model.

3.5 Conclusions

This chapter incorporates both macroeconomic and financial elements into exchange rate modeling. Separating out the *expected yields* and the *term premiums* from the yields, we show that investors' expectation about the future path of monetary policy and their perceived risk both drive exchange rate dynamics. Their relative role depends on the state of the economy as expectations have a consistent role while risk presents itself as a significant determinant during the crisis. We then examine why risk matters more and propose the recent changes in monetary policies like ZLB and QE as one explanation. In the same context, we argue that the macro variables should be jointly modeled with yield curve factors in order to capture both Taylor-rule type and unconventional monetary policies.

[35]The yield-only model regresses the quarterly exchange rate on relative Nelson-Siegel factors, while the macro-only model regress on *relative inflation rate* and *relative unemployment gap*.

Table 3.1: Summary Statistics for Exchange Rate and Macro Variables

		AU	CA	DK	JP	NZ	SE	CH	UK
Panel A: Exchange Rate Change (Δs_{t+3})									
Mean	Full	0.085	-0.171	0.775	0.969	-0.174	0.603	-0.967	0.565
	Pre-Break	-1.855	-2.433	-1.397	0.931	-1.255	-1.554	-1.188	-1.636
	Post-Break	3.457	3.763	4.551	1.037	1.705	4.353	-0.581	4.394
SD	Full	25.355	16.533	20.226	22.940	25.246	22.584	20.226	17.058
	Pre-Break	19.257	13.441	18.850	23.412	21.480	19.304	20.042	12.535
	Post-Break	33.275	20.340	22.020	22.223	30.762	27.080	20.647	22.464
$AR(1)$	Full	0.733	0.634	0.705	0.701	0.748	0.724	0.664	0.729
Panel B: Relative Unemployment Gap (u_t^R)									
Mean	Full	-0.010	-0.009	-0.006	0.005	-0.023	-0.017	-0.005	-0.005
	Pre-Break	0.013	0.013	0.027	0.042	-0.004	-0.005	0.038	0.022
	Post-Break	-0.050	-0.049	-0.063	-0.060	-0.055	-0.037	-0.080	-0.053
SD	Full	0.322	0.236	0.351	0.302	0.374	0.384	0.404	0.302
	Pre-Break	0.319	0.239	0.342	0.273	0.374	0.419	0.388	0.277
	Post-Break	0.325	0.228	0.360	0.337	0.373	0.313	0.421	0.339
$AR(1)$	Full	0.793	0.663	0.860	0.812	0.815	0.644	0.835	0.866
Panel C: Relative Inflation Rate (π_t^R)									
Mean	Full	0.394	-0.390	-0.312	-2.152	-0.111	-1.156	-1.679	-0.220
	Pre-Break	0.104	-0.618	-0.546	-2.666	-0.380	-1.399	-1.727	-0.900
	Post-Break	0.898	0.004	0.095	-1.257	0.356	-0.733	-1.596	0.962
SD	Full	1.283	0.854	0.985	1.554	1.188	1.169	0.789	1.323
	Pre-Break	1.321	0.757	0.832	1.139	0.937	1.213	0.529	0.880
	Post-Break	1.042	0.871	1.098	1.765	1.418	0.955	1.103	1.117
$AR(1)$	Full	0.905	0.910	0.925	0.948	0.900	0.934	0.904	0.956

Note: 1.$\Delta s_{t+3} = s_{t+3} - s_t$ is the quarterly change of the exchange rate, where s_t is the logged home currency price per USD. If Δs_{t+3} is positive, the home currency is depreciated relative to USD. If Δs_{t+3} is negative, the home currency is appreciated relative to USD. 2. $u_t^R = u_t - u_t^*$ is the relative unemployment gap, constructed as the difference between the detrended unemployment rates in the home country and in the US. Time series of the unemployment rate is detrended by Hodrick-Prescott filter. 3. $\pi_t^R = \pi_t - \pi_t^*$ is the relative inflation rate, defined as the 12-month change of the CPI in each country relative to that in the US. 3. Sample period is from January 1995 to March 2016. All rates are reported in annualized percentage. 5. Sample period is divided by the breakdate, May 2008. Sub-periods before and after the breakdate are reported as “Pre-Break” and “Post-Break”, respectively. The breakdate is chosen, considering that the most common breakdate by the structural break tests in the later sections is May 2008.

Table 3.2: Explaining Exchange Rate Change with Yield Factors

$$\Delta s_{t+3} = \beta_0 + \beta_1 L_t^R + \beta_2 S_t^R + \beta_3 C_t^R + \epsilon_{t+3}$$

	AU	CA	DK	JP	NZ	SE	CH	UK
β_0								
Pre-Break	4.536	-1.315	-3.834**	11.916	8.603	-1.912	4.034	0.506
	(3.428)	(1.798)	(1.820)	(15.707)	(8.449)	(1.849)	(9.474)	(2.737)
Post-Break	-22.716	-16.827***	-10.721	31.812***	-45.982***	-13.713**	-20.748	-6.986**
	(20.430)	(4.527)	(6.541)	(11.256)	(12.180)	(6.921)	(18.651)	(3.178)
β_1								
Pre-Break	-4.216**	-3.790*	-1.469	3.855	-4.232	-2.327**	3.577	-1.741
	(1.978)	(2.027)	(1.925)	(5.193)	(6.962)	(1.080)	(4.663)	(2.333)
Post-Break	-4.558	4.553	-2.611	57.891***	10.223*	4.449	-17.778	9.860*
	(10.943)	(4.235)	(4.326)	(8.638)	(5.636)	(6.493)	(13.290)	(5.969)
β_2								
Pre-Break	-2.194	-2.831***	-1.538	-1.142	-1.324	-1.546	-0.774	-0.438
	(1.892)	(0.802)	(1.758)	(2.269)	(1.322)	(1.264)	(1.534)	(1.317)
Post-Break	19.050**	19.229***	7.061*	41.258***	18.868***	19.191***	-12.316	19.766***
	(8.785)	(2.646)	(3.620)	(8.555)	(2.472)	(4.944)	(11.020)	(2.439)
β_3								
Pre-Break	-2.535***	-0.740	-2.428**	-0.337	-3.573***	-3.159***	-2.294***	-1.237
	(0.977)	(0.974)	(1.023)	(1.258)	(1.253)	(0.916)	(0.881)	(0.978)
Post-Break	-4.727**	-0.049	2.203	5.763***	3.187*	-3.368*	4.538	2.509
	(2.189)	(1.173)	(2.680)	(1.844)	(1.749)	(1.758)	(3.575)	(2.114)
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.019	0.000
Adj. R^2	0.175	0.215	0.144	0.135	0.291	0.282	0.111	0.305
Breakdate	May,08	Jan,07	May,08	Jul,08	May,08	May,08	Jun,11	May,08

Note: 1. We first regress 3-month exchange rate changes on the relative NS factors and then apply the Quants-Andrews unknown breakpoint test (Andrews, 1993; Hansen, 1997) with 15% trimming and 5% significance level to detect the structural break in the regression. Break dates are mostly around 2008, as reported in the last row. Each breakdate indicates the first month of subsequent sub-period. 2. After identifying the break date, structural break dummy variables for each sub-period are incorporated into the regression. We categorize two phases identified by the break as "Pre-Break" and "Post-Break". 3. Coefficient estimates are reported with the Newey-West standard errors in the parentheses. Asterisks indicate significance levels at 1% (***), 5% (**), and 10% (*) respectively. 4. P-value is for the Wald test that factors jointly have no explanatory power ($H_0 : \beta_1 = \beta_2 = \beta_3 = 0, \ \forall \ sub - periods$).

Table 3.3: Explaining Exchange Rate Change with Expectations or Risk or Both

	AU	CA	DK	JP	NZ	SE	CH	UK
Panel A: Expectation-Only Model								
$\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \epsilon_{t+3}$								
1) NS model								
p-value	0.001	0.000	0.000	0.000	0.000	0.000	0.001	0.000
Adj. R^2	0.219	0.194	0.148	0.140	0.307	0.248	0.117	0.308
Breakdate	May,08	Jan,07	May,08	Aug,12	May,08	May,08	Feb,08	May,08
2) NSM model								
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.012	0.000
Adj. R^2	0.196	0.242	0.207	0.140	0.319	0.281	0.110	0.297
Breakdate	Dec,03	Jan,07	May,08	May,98	May,08	May,08	Jun,11	Jun,08
3) Affine model								
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.010	0.000
Adj. R^2	0.241	0.213	0.137	0.202	0.335	0.243	0.103	0.293
Breakdate	May,08	Jan,07	May,08	Sep,08	May,08	May,08	Feb,08	May,08
4) CP model								
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj. R^2	0.259	0.180	0.154	0.158	0.338	0.261	0.134	0.292
Breakdate	Jan,04	Jan,07	May,08	Sep,12	Apr,08	Oct,05	Jun,11	May,08
Panel B: Risk-Only Model								
$\Delta s_{t+3} = \beta_0 + \beta_1 S(\theta_t^R) + \beta_2 C(\theta_t^R) + \epsilon_{t+3}$								
1) NS model								
p-value	0.068	0.026	0.001	0.004	0.000	0.009	0.001	0.000
Adj. R^2	0.134	0.096	0.093	0.152	0.141	0.122	0.115	0.295
Breakdate	May,08	Oct,07	Oct,05	Oct,98	Jan,09	May,08	Jun,11	May,08
2) NSM model								
p-value	0.043	0.002	0.189	0.001	0.001	0.581	0.000	0.014
Adj. R^2	0.154	0.119	0.086	0.134	0.159	0.072	0.124	0.060
Breakdate	May,08	Oct,07	May,08	Sep,02	Mar,08	May,08	Jun,11	Oct,08
3) Affine model								
p-value	0.001	0.000	0.000	0.041	0.000	0.001	0.030	0.050
Adj. R^2	0.167	0.182	0.111	0.109	0.243	0.092	0.099	0.190
Breakdate	May,08	May,08	Mar,08	May,98	Apr,08	May,01	Jun,11	May,08
4) CP model								
p-value	0.067	0.056	0.000	0.023	0.004	0.033	0.013	0.559
Adj. R^2	0.148	0.087	0.121	0.090	0.151	0.077	0.096	0.076
Breakdate	May,08	Oct,07	Sep,05	Jun,03	Feb,08	Mar,08	Jun,11	Sep,07

(Continued)

	AU	CA	DK	JP	NZ	SE	CH	UK
Panel C: Both-Expectations-Risk Model								
$\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \beta_4 S(\theta_t^R) + \beta_5 C(\theta_t^R) + \epsilon_{t+3}$								
1) NS model								
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj. R^2	0.288	0.264	0.163	0.260	0.329	0.303	0.137	0.335
Breakdate	May,08	Jan,07	May,08	Sep,08	May,08	May,08	Jun,11	Jun,08
2) NSM model								
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj. R^2	0.237	0.248	0.198	0.264	0.326	0.313	0.148	0.339
Breakdate	May,08	Jan,07	May,08	Sep,08	May,08	May,08	Jun,11	Jun,08
3) Affine model								
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj. R^2	0.290	0.256	0.184	0.260	0.338	0.321	0.112	0.308
Breakdate	May,08	Jul,08	May,08	Sep,08	May,08	May,08	Jan,09	May,08
4) CP model								
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj. R^2	0.262	0.176	0.197	0.192	0.334	0.262	0.135	0.321
Breakdate	May,08	Jan,07	Oct,05	Nov,07	Apr,08	May,08	Jun,11	May,08

Note: 1. The *relative expected yield* and *relative term premium* factors are estimated by four models: 1) NS model, 2) NSM model, 3) Affine model, 4) CP model. 2. The unknown break point is tested by Quants-Andrews test with 15% trimming and 5% significance level (Andrews, 1993; Hansen, 1997). Break dates are mostly around 2008 with some exceptions. Each breakdate indicates the first month of subsequent sub-period. 3. After identifying the break date, structural break dummy variables for each sub-period are incorporated into the regression. 4. P-value is for the Wald test that factors jointly have no explanatory power ($H_0 : \beta_i = 0,\ \forall\ i,\ \forall\ sub-periods$).

Table 3.4: Explaining Exchange Rate Change with Both Expectations and Risk

$$\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \beta_4 S(\theta_t^R) + \beta_5 C(\theta_t^R) + \epsilon_{t+3}$$

	AU	CA	DK	JP	NZ	SE	CH	UK
β_0								
Pre-	12.665***	-2.538	-3.964	67.769***	-1.407	-3.121	3.372	0.061
Break	(4.541)	(2.616)	(3.926)	(26.236)	(12.162)	(3.040)	(14.448)	(4.942)
Post-	7.643	-23.573***	-13.846	-17.433	-65.486***	18.517	-2.172	-14.925**
Break	(17.662)	(6.390)	(9.825)	(20.945)	(16.321)	(16.375)	(19.560)	(7.161)
β_1								
Pre-	-9.867***	-4.358	-2.957	29.241***	4.574	4.227	2.342	-2.952
Break	(3.350)	(4.515)	(2.032)	(8.940)	(9.077)	(8.435)	(7.372)	(3.451)
Post-	-12.502	7.354	-2.345	72.169***	27.840***	-12.688	2.482	25.492***
Break	(8.790)	(6.816)	(5.554)	(8.870)	(8.958)	(10.464)	(8.819)	(8.733)
β_2								
Pre-	-2.238	-3.162***	-4.884*	-7.576***	-3.524	-2.900	-1.306	0.874
Break	(2.164)	(0.915)	(2.504)	(2.821)	(2.356)	(2.115)	(3.202)	(1.492)
Post-	5.872	20.391***	13.429***	79.160***	12.335**	22.648***	-23.539***	28.347***
Break	(8.948)	(4.402)	(4.462)	(12.919)	(5.244)	(4.806)	(8.058)	(4.799)
β_3								
Pre-	-0.419	1.013	-2.178	-21.446***	-14.394***	-13.789	-7.775	2.098
Break	(3.525)	(3.758)	(4.080)	(7.460)	(4.174)	(8.376)	(6.352)	(3.600)
Post-	-1.816	10.347	6.396	36.345***	2.165	-4.222	1.084	1.350
Break	(7.605)	(7.137)	(10.751)	(9.162)	(4.457)	(8.229)	(9.818)	(9.280)
β_4								
Pre-	-3.184	0.212	2.481	1.015	-3.500	-0.777	-7.297	-3.472
Break	(2.963)	(2.902)	(4.702)	(4.858)	(9.230)	(2.408)	(6.863)	(7.613)
Post-	36.771***	-0.364	3.973	-44.652***	-8.557	14.884*	42.109***	1.760
Break	(14.101)	(10.414)	(8.269)	(7.313)	(10.826)	(8.194)	(8.684)	(9.846)
β_5								
Pre-	8.678*	1.054	1.143	-34.353***	-16.377	-15.322	-5.247	-7.087
Break	(5.201)	(5.532)	(7.307)	(8.225)	(10.166)	(10.806)	(6.107)	(4.722)
Post-	21.133	20.679*	-1.984	71.734***	17.962	-7.456	20.950	14.799**
Break	(13.992)	(11.853)	(15.560)	(15.858)	(21.505)	(11.185)	(13.013)	(6.257)
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj. R^2	0.237	0.248	0.198	0.264	0.326	0.313	0.148	0.339
Breakdate	May,08	Jan,07	May,08	Sep,08	May,08	May,08	Jun,11	Jun,08

Note: 1. The *relative expected yield* and *relative term premium* factors are estimated by NSM model. The results using factors from other models can be provided upon request. 2. After identifying the unknown breakpoint around 2008 by the Quants-Andrews test (Andrews, 1993; Hansen, 1997) with 15% trimming and 5% significance level, structural break dummy variables for each sub-period are incorporated into the regression. 3. Coefficient estimates are reported with the Newey-West standard errors in the parentheses. Asterisks indicate significance levels at 1% (***), 5% (**), and 10% (*) respectively.

Table 3.5: The Relative Role of Expectations vs. Risk – Wald Test

$$\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \beta_4 S(\theta_t^R) + \beta_5 C(\theta_t^R) + \epsilon_{t+3}$$

	AU	CA	DK	JP	NZ	SE	CH	UK
Wald test								
No Expectation?								
Full	0.000	0.000	0.000	0.000	0.000	0.000	0.001	0.000
Pre-Break	0.000	0.003	0.001	0.008	0.001	0.000	0.363	0.671
Post-Break	0.050	0.000	0.014	0.000	0.000	0.000	0.000	0.000
No Risk?								
Full	0.003	0.461	0.957	0.000	0.455	0.103	0.000	0.003
Pre-Break	0.248	0.982	0.805	0.000	0.256	0.366	0.292	0.025
Post-Break	0.001	0.168	0.891	0.000	0.648	0.059	0.000	0.014
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj. R^2	0.237	0.248	0.198	0.264	0.326	0.313	0.148	0.339
Breakdate	May,08	Jan,07	May,08	Sep,08	May,08	May,08	Jun,11	Jun,08

Note: 1. The *relative expected yield* and *relative term premium* factors are estimated by NSM model. The results using factors from other models can be provided upon request. 2. After identifying the unknown breakpoint around 2008 by the Quants-Andrews test (Andrews, 1993; Hansen, 1997) with 15% trimming and 5% significance level, structural break dummy variables for each sub-period are incorporated into the regression. 3. The row labeled "No Expectation" reports the p-values of the Wald tests for the null hypothesis that the *relative expected yield* factors have no explanatory power ($\beta_1 = \beta_2 = \beta_3 = 0$, for each sub-period), and the "No Risk" row tests the null hypothesis that the *relative term premium* factors do not matter ($\beta_4 = \beta_5 = 0$, for each sub-period). For example, "No Expectation" in "Pre-Break" period tests the null hypothesis that the *relative expected yield* factors have no explanatory power during the sub-period before the breakdate. 4. P-value is for the Wald test that factors jointly have no explanatory power ($H_0 : \beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = 0, \; \forall \; sub-periods$).

Table 3.6: The Relative Role of Yields vs. Macro – Wald Test

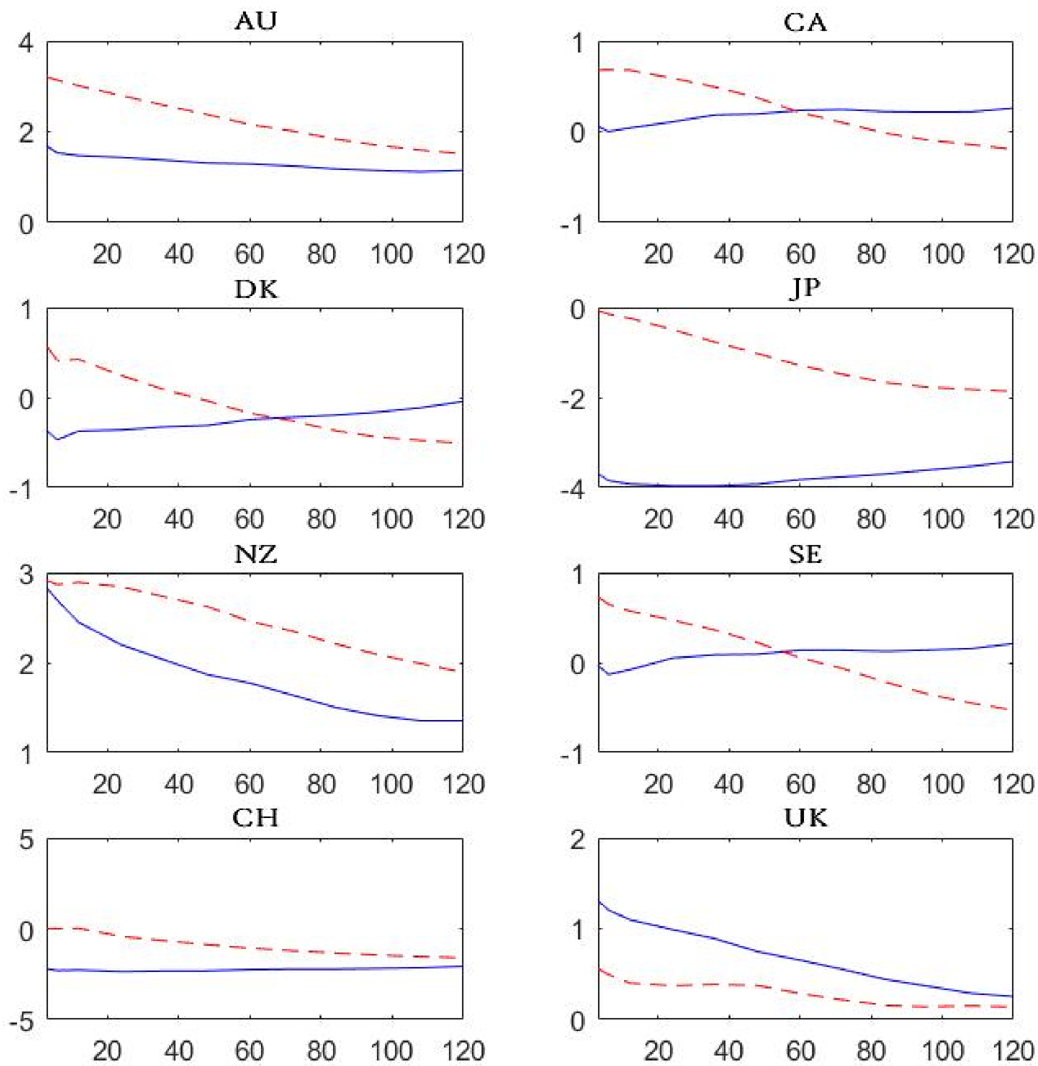
$$\Delta s_{t+3} = \beta_0 + \beta_1 L_t^R + \beta_2 S_t^R + \beta_3 C_t^R + \beta_4 u_t^R + \beta_5 \pi_t^R + \epsilon_{t+3}$$

	AU	CA	DK	JP	NZ	SE	CH	UK
Wald test								
No Yields?								
Full	0.003	0.000	0.000	0.000	0.000	0.000	0.000	0.001
Pre-Break	0.001	0.014	0.000	0.047	0.002	0.099	0.000	0.625
Post-Break	0.352	0.000	0.061	0.000	0.000	0.000	0.031	0.000
No Macro?								
Full	0.004	0.029	0.000	0.313	0.000	0.000	0.000	0.000
Pre-Break	0.645	0.843	0.111	0.711	0.004	0.305	0.001	0.02
Post-Break	0.001	0.005	0.000	0.126	0.006	0.000	0.000	0.002
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj. R^2	0.241	0.276	0.248	0.149	0.388	0.340	0.185	0.383
Breakdate	May,08	Jan,07	May,08	Jul,08	May,08	May,08	Sep,03	May,08

Note: 1. After identifying the unknown breakpoint around 2008 by the Quants-Andrews test (Andrews 1993, Hansen 1997) with 15% trimming and 5% significance level, structural break dummy variables for each sub-period are incorporated into the regression. 2. The row labeled "No Yields" reports the p-values of the Wald tests for the null hypothesis that the *relative yield* factors have no explanatory power ($\beta_1 = \beta_2 = \beta_3 = 0$, for each sub-period), and the "No Macro"row tests the null hypothesis that macroeconomic fundamentals do not matter ($\beta_4 = \beta_5 = 0$, for each sub-period). For example, "No Yields" in "Pre-Break" period tests test the null hypothesis that *relative yield* factors have no explanatory power during the sub-period before the breakdate. 3. P-value is for the Wald test that factors jointly have no explanatory power ($H_0 : \beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = 0,\ \forall\ sub-periods$).

Figure 3.1: Relative Yield Curves Before and After the Break

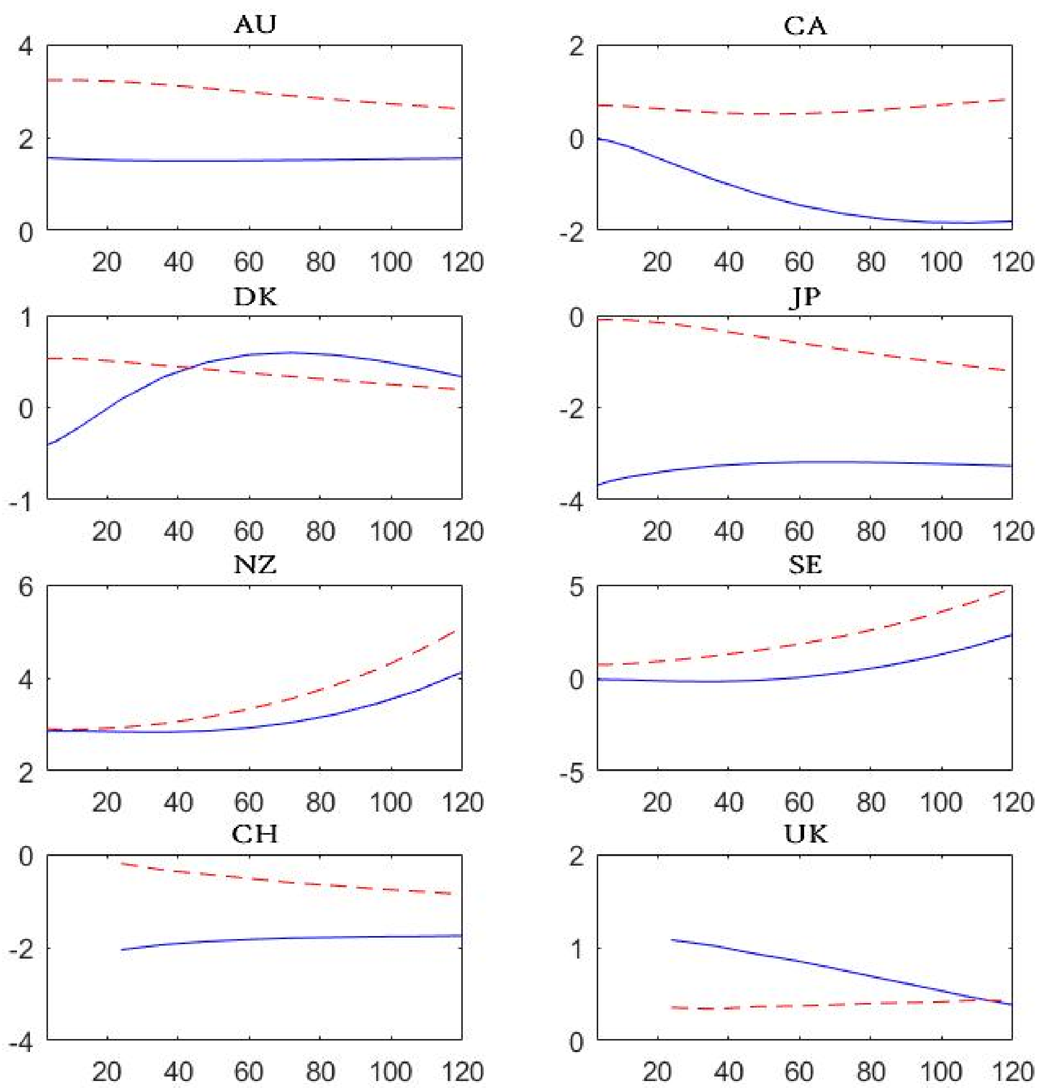
(Annualized %)



Note: 1. The sample averages of *relative yields*, $i_t^{R,m} = i_t^m - i_t^{*,m}$, over January 1995 - April 2008 (Blue solid line) and over May 2008 - December 2015 (Red dashed line) are shown in the figure. 2. The break date is chosen, considering that the most common breakdate by the structural break tests in the related regressions is May 2008.

Figure 3.2: Relative Expected Yield Curves Before and After the Break

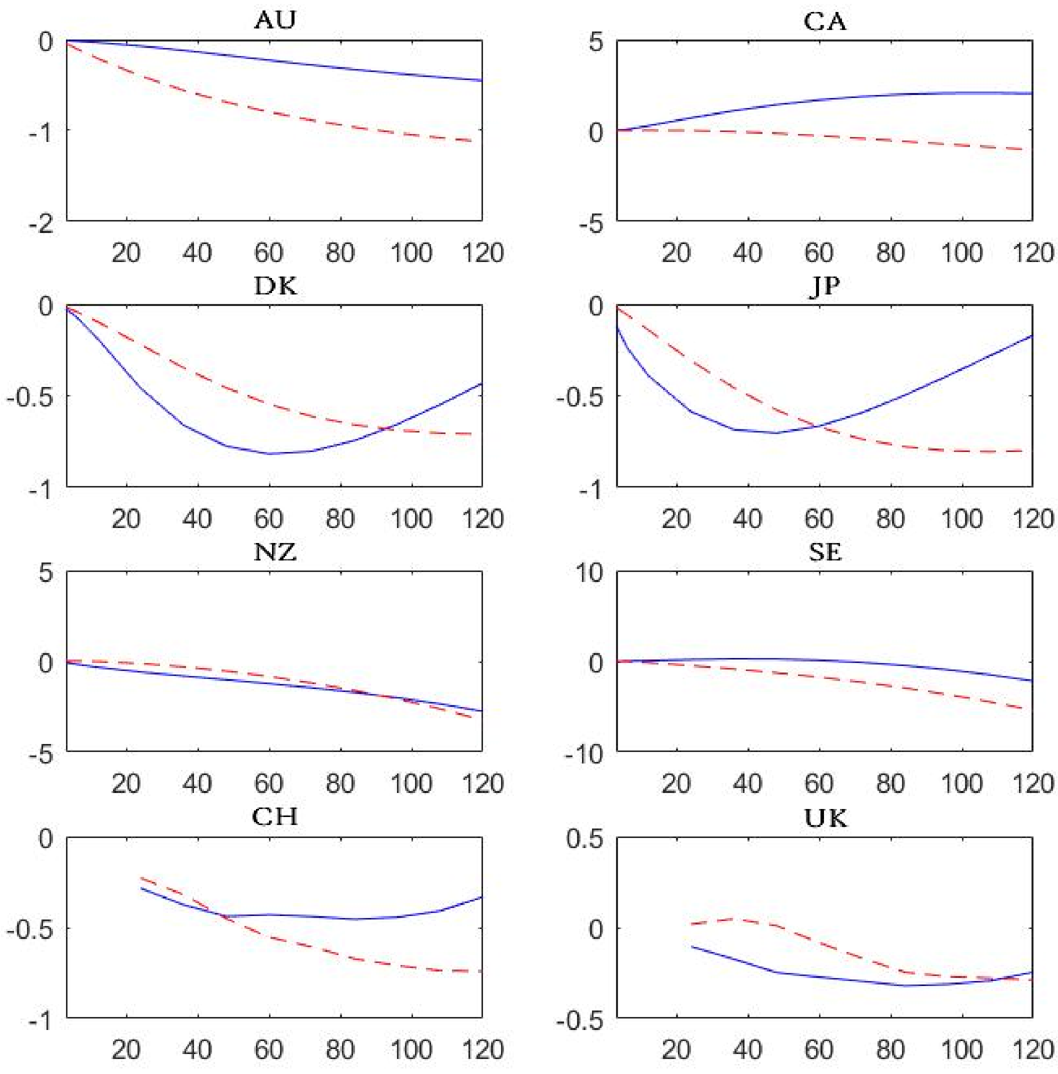
(Annualized %)



Note: 1. The sample averages of *relative expected yields*, $E_t i_t^{R,m} = E_t i_t^m - E_t i_t^{*,m}$, over January 1995 - April 2008 (Blue solid line) and over May 2008 - December 2015 (Red dashed line) are shown in the figure. 2. We selectively show the *relative expected yields* from four models: NS model for AU and CA, NSM model for DK and JP, Affine model for NZ and SE, CP model for CH and UK. Figures from each model with all eight countries are available upon request.

Figure 3.3: Relative Term Premium Curves Before and After the Break

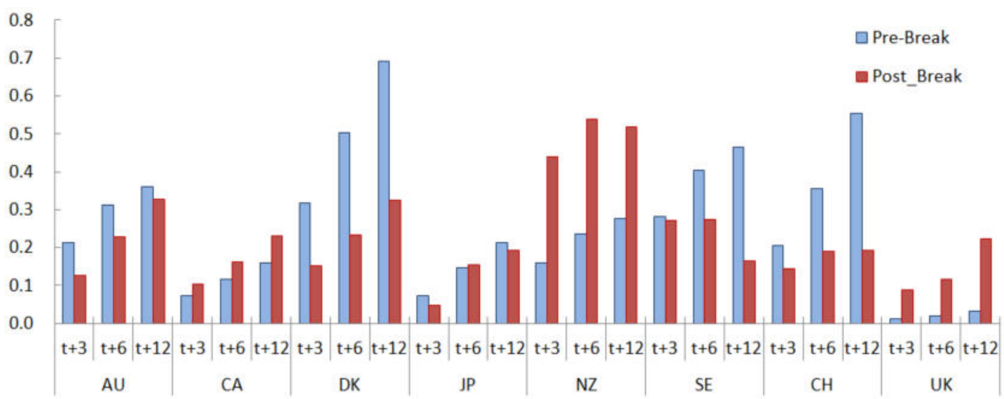
(Annualized %)



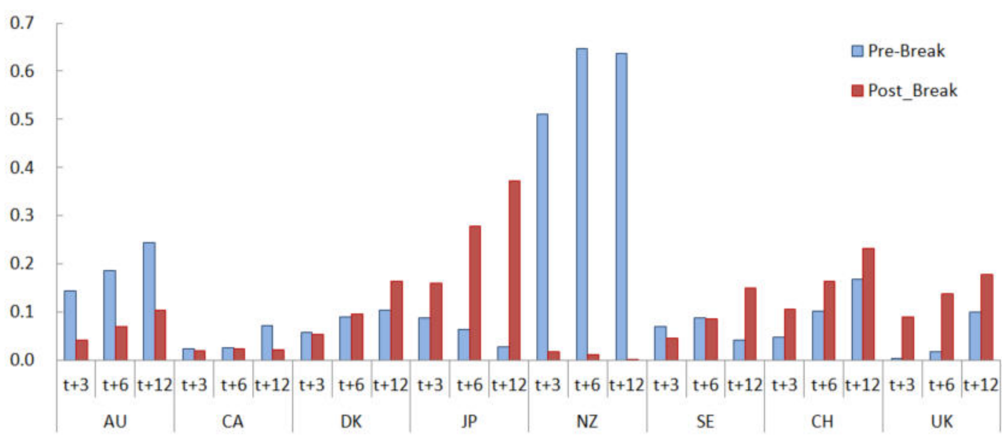
Note: The sample averages of *relative term premiums*, $\theta_t^{R,m} = \theta_t^m - \theta_t^{*,m}$, over January 1995 - April 2008 (Blue solid line) and over May 2008 - December 2015 (Red dashed line) are shown in the figure. 2. We selectively show the *relative term premiums* from four models: NS model for AU and CA, NSM model for DK and JP, Affine model for NZ and SE, CP model for CH and UK. Figures from each model with all eight countries are available upon request.

Figure 3.4: Explained Portion by Expectations and Risk: Partial R^2

1. Explained Portion by Expectations



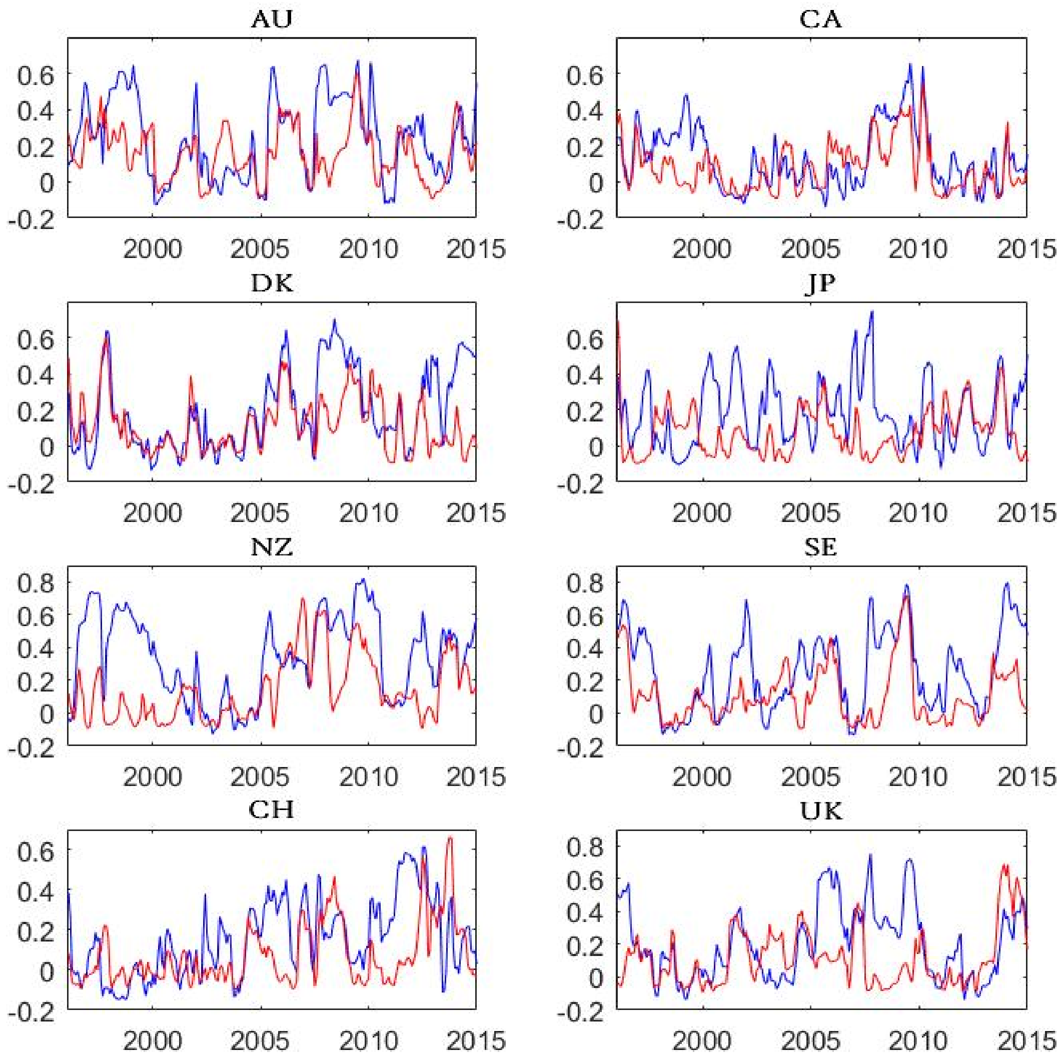
2. Explained Portion by Risk



Note: 1. The partial R^2 reports the contribution of each variable in explaining Δs_{t+k} for $k = 1, 3, 6, 12$. With $f_t = \left(\Delta s_t, L(E_t i_t^R), S(E_t i_t^R), C(E_t i_t^R), S(\theta_t^R), C(\theta_t^R)\right)$, it is constructed by first estimating $f_t - \mu = A(f_{t-1} - \mu) + v_t$ and then using $\widehat{A}$ and the estimated covariance matrix of the $VAR(1)$, as in Hodrick (1992). See Appendix for details. 2. Factors from the NSM model are used. Results from other models can be provided upon request. 3. We report the partial R^2 for two different sample periods: 1) Pre-Break: from January 1995 to the break date, which is identified in **Table 3.5**, 2) Post-Break: from the break date to December 2015. 4. "Expectation" and "Risk" are the sum of the partial R^2's of each group. As the variables are correlated, the explanatory power of each group is different from arithmetical sum. They are reported only for comparison purpose.

Figure 3.5: Comparing the Ad.R^2 over Rolling Windows: Expectations or Risk?

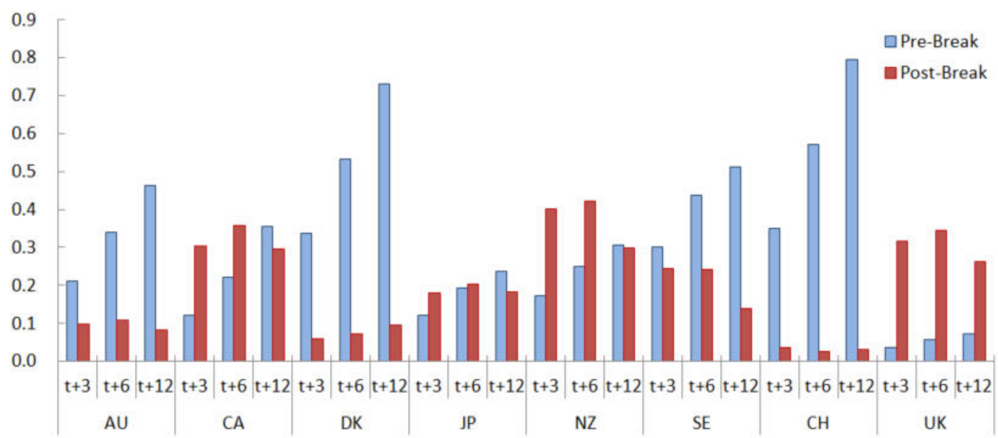
(Adjusted R^2)



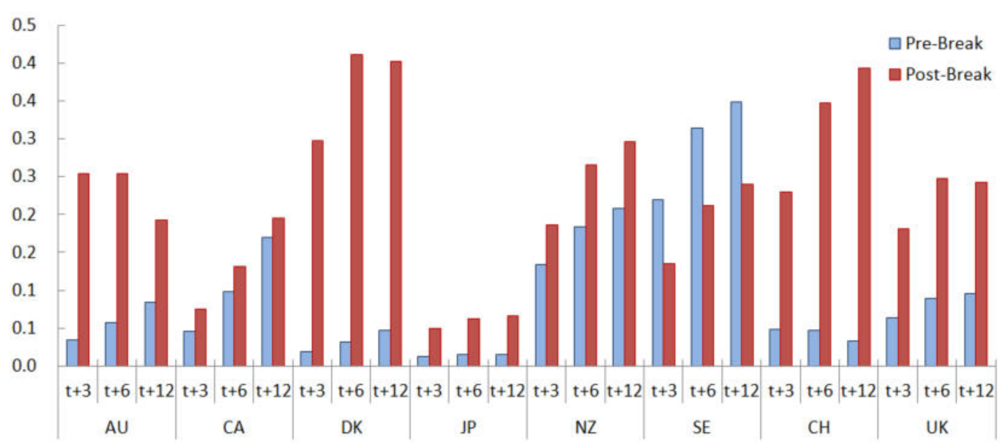
Note: 1. This figure shows the adjusted R-squared from the following OLS regressions with 24-month rolling window: (i) 3-month exchange rate change on the *relative expected yield* factors (Blue solid line), (ii) 3-month exchange rate change on the *relative term premium* factors (Red solid line). 2. Factors are obtained from NSM model. 3. The date on the x-axis represents the midpoint month of each window.

Figure 3.6: Explained Portion by Yields and Macro: Partial R^2

1. Explained Portion by Yield Factors



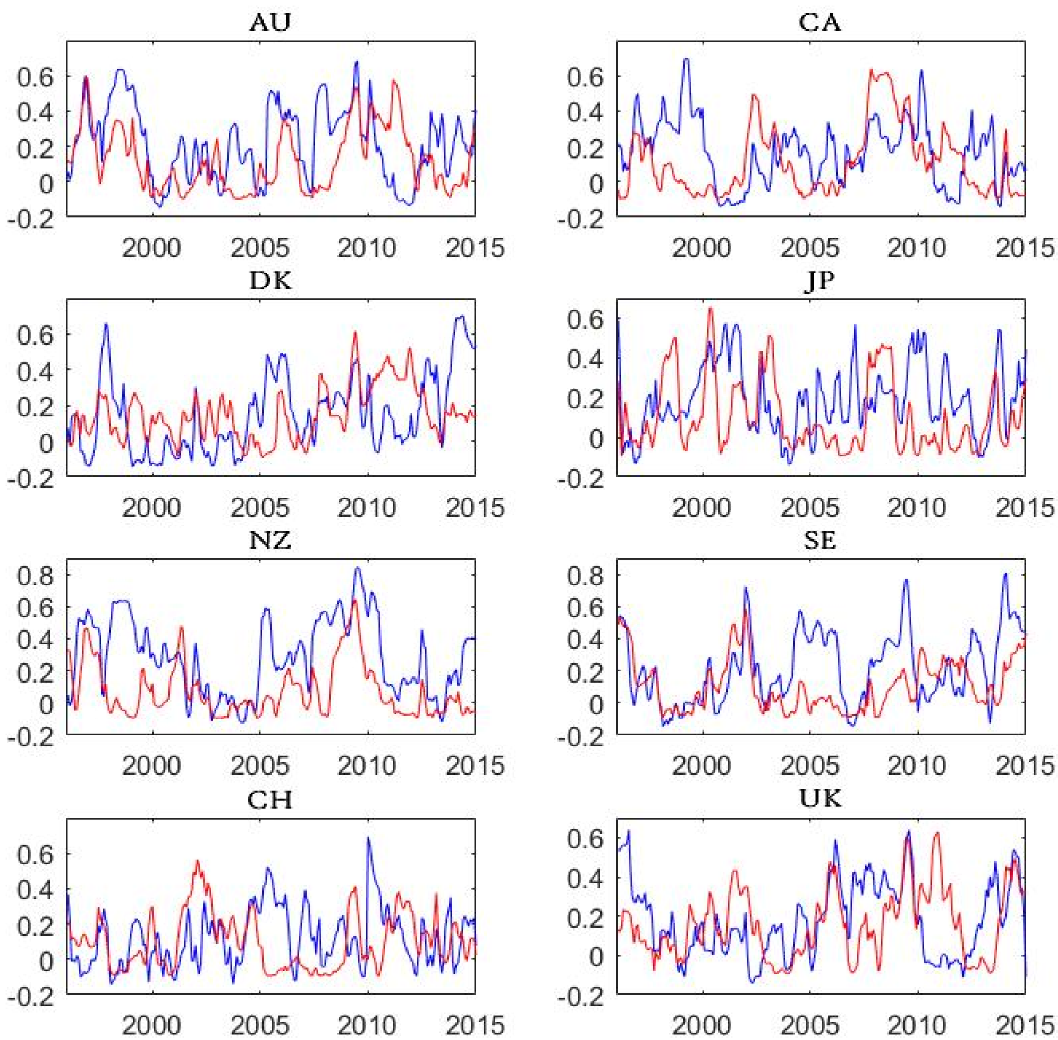
2. Explained Portion by Macro-Fundamentals



Note: 1. The partial R^2 reports the contribution of each variable in explaining Δs_{t+k} for $k = 1, 3, 6, 12$. It is constructed by first estimating $f_t - \mu = A(f_{t-1} - \mu) + v_t$, where $f_t = \left(u_t^R, \pi_t^R, \Delta s_t, L_t^R, S_t^R, C_t^R\right)$, and then using $\widehat{A}$ and the estimated covariance matrix of the $VAR(1)$, as in Hodrick (1992). See Appendix for details. 2. We report the partial R^2 for two different sample periods: 1) Pre-Break: from January, 1995 to the break date, which is identified in **Table 3.6** 2) Post-Break: from the break date to December, 2015. 3. "Yields" and "Macro" are the sum of the partial R^2's of each group. As the variables are correlated, the explanatory power of each group is different from arithmetical sum. They are reported only for comparison purpose.

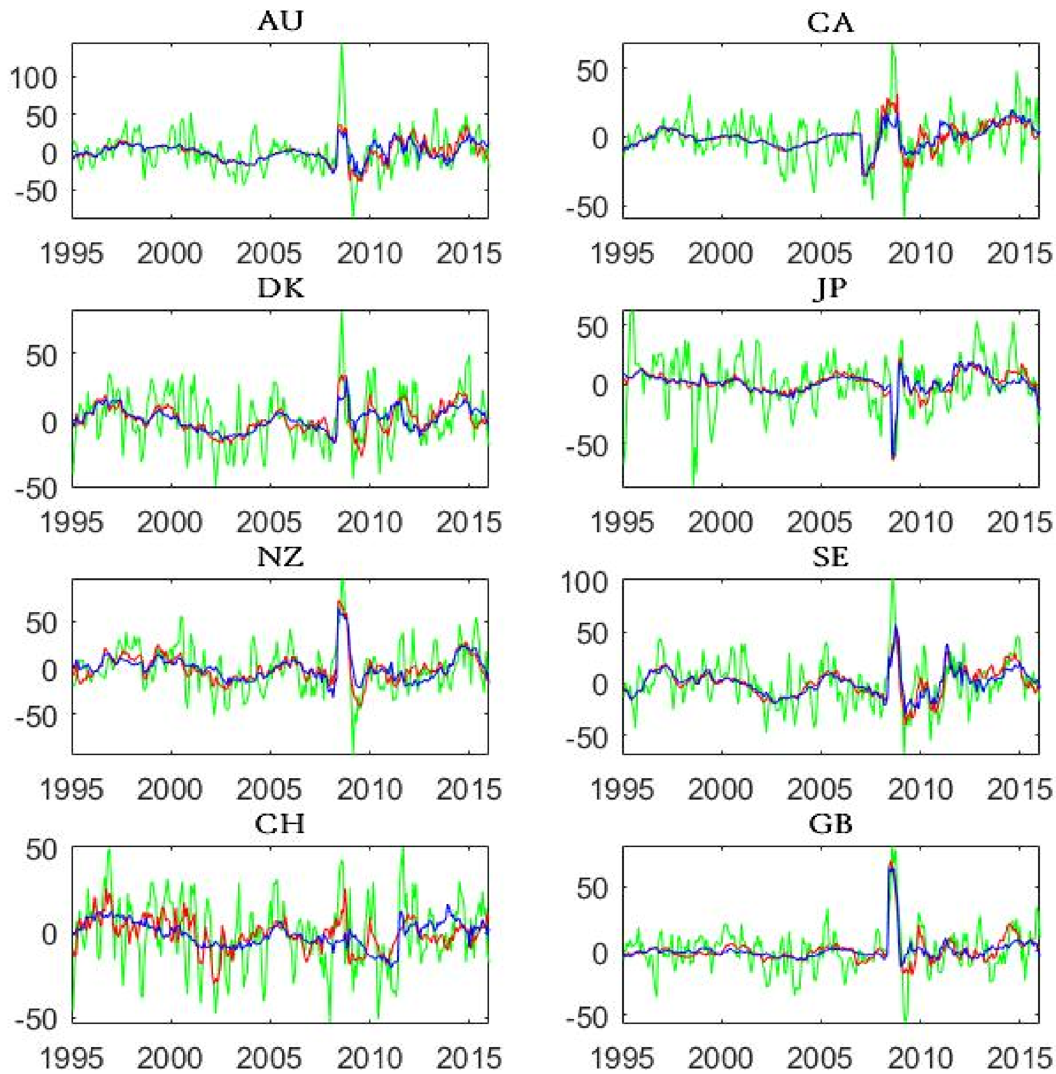
Figure 3.7: Comparing the Ad.R^2 over Rolling Windows: Yield or Macro?

(Adjusted R^2)



Note: 1. This figure shows the adjusted R-squared from the following OLS regressions with 24-month rolling window: (i) 3-month exchange rate change on the *relative yield* factors (Blue solid line), (ii) 3-month exchange rate change on the macro variables (Red solid line). 2. The date on the x-axis represents the midpoint month of each window.

Figure 3.8: Fitted vs. Actual Quarterly Exchange Rate Changes



Note: 1. This figure shows the fitted quarterly exchange rate changes from the joint model (Red solid line), the fitted quarterly exchange rate changes from the yield-only model (Blue solid line) and the actual quarterly exchange rate changes (Green solid line). 2. Sample period is from January 1995 to March 2016. All rates are in annualized percentage points.

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Appendix A

APPENDIX TO CHAPTER 2

A.1 What is Sovereign Credit Default Swap?

A sovereign credit default swap (CDS) is a bilateral Over-the-Counter (OTC) agreement in which protection buyer purchases insurance against contingent credit events by paying an annuity premium quarterly or bi-annually and protection seller compensates buyer for the loss given credit events. Buyers are usually banks, security firms, and hedge funds, while sellers are mostly insurance companies and banks. In practice, there is no default in government bonds. Instead, International Swap and Derivative Association (ISDA) references four types of credit events: acceleration, failure to pay, restructuring, and repudiation.

There are three types of cash flows associated with a CDS contract. 1) If there is no credit event until the end of the contract, only protection buyer pays an annuity premium for the contract tenor. The protection seller pays nothing. 2) If one of the credit events is triggered prior to the tenor, the buyer pays an annuity premium until the credit event happens and the contract terminates. The seller compensates the buyer for its loss given a credit event. There are two types of settlement. In a physical settlement, the buyer provides the seller the deliverable obligations of the reference entity and the seller pays the buyer a cash payment amounting to its full aggregate notional amount. In a cash settlement, the seller pays the buyer a cash payment amounting to a full in the market value of a debt obligation of the reference entity. 3) Each counter party of the CDS contract can unwind the contract prior to the tenor. In order to unwind the contract, they need to agree on early termination or assignment to another counter party or they can offset the transaction in the market. In any case, they need to pay or receive the Mark-to-Market value, which is the present value of the difference in annuity payment over the remaining contract period.

Sovereign CDSs have been increasing in use from the early 2000s. According to the Bank for International Settlement (BIS) statistics, sovereign CDS contracts increased dramatically from \$6.4 trillion in 2004 to a peak of \$58.2 trillion in 2007. The amount has since come down because the credit derivatives were central to the 2007-2009 financial crisis.

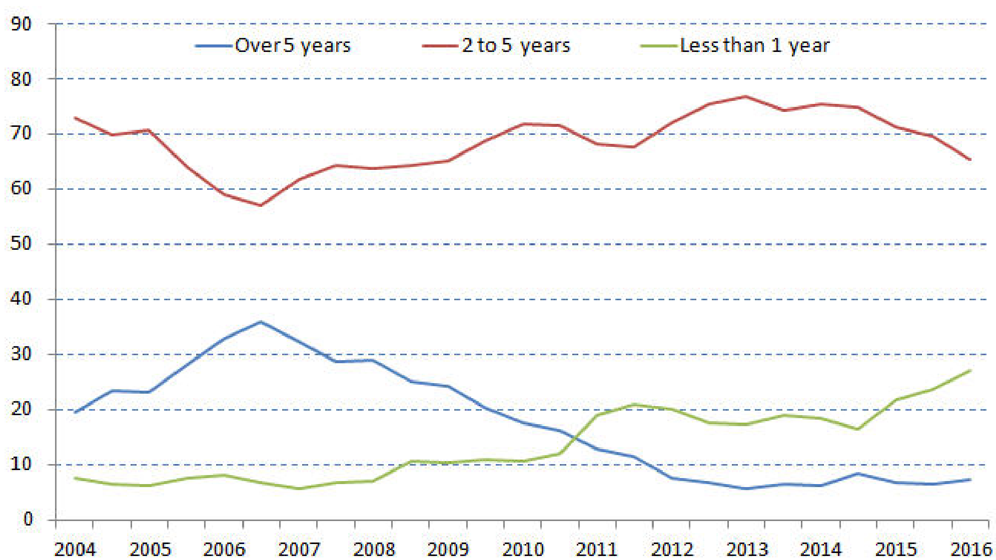
Sovereign CDS: Outstanding notional amount (billion USD)

2004	2005	2006	2007	2008	2009	2010	2011	2012	2013	2014	2015	2016
6396	13908	28650	58244	41883	32693	29898	28626	25068	21020	16399	12294	9857

Note: 1. Each number represents the gross notional amount outstanding in the second half of each year. 2. source: www.bis.org

While corporate CDSs are mostly concentrated around five-year contracts, sovereign CDSs are well diversified over different tenors. Based on the BIS statistics, I compute the market shares of each tenor contract from 2004 to 2016. The total volume of the notional amount outstanding with a tenor of less than one year is 12.01%, with a tenor of between two to five years 67.10%, and with a tenor of over five years 20.88%. The interesting feature is that after 2011, short tenors have gradually been trading more relative to long tenors.

Market of sovereign CDS over different tenors (% in volume)



Note: 1. Market share is computed as the total volume of the notional amount outstanding of each tenor group compared to that of all tenors. 2. source: www.bis.org

A.2 Robustness Checks

I conduct extensive robustness checks to corroborate the key findings: 1) sovereign credit risk measures can explain a considerable amount of the variation in currency returns; 2) risk level and slope factors together explain the currency movements better than one-year CDS – both risk level and proximity play a role in forecasting currency returns; 3) the relationship between credit risk and currency returns is state-dependent; and 4) risk level matters more during normal times, while risk proximity is more important near crises.

Over different horizons ($m = 1-, 3-, 6-, 12 - months$), I repeat the regressions of the exchange rate changes on four sets of credit risk measures (eq.(2.8), (2.9)(2.10) and (2.11)) with structural breaks. Adjusted R^2 results in **Table A.1** are comparable to those in **Table 2.4**. All four sets of credit risk measures can explain a sizable amount of the variation in currency movements from monthly to yearly. Adjusted R^2s are very high, up to 79% in the yearly prediction, while relatively low in the monthly prediction. Comparing the explanatory powers across models, the model with both level and slope factors consistently produces the highest goodness of fit. The Markov-Switching model is also applied to different prediction horizons to check the state-dependent property of estimates. The results for monthly prediction reported in **Table A.2** replicate all the findings from quarterly prediction: two regimes represent high and low volatility states, the persistence of the high volatility state is country-dependent, and coefficient estimates switch depending on the state of the economy.

In the main results, I rely on one country's CDS spreads rather than the cross-country difference in CDS spreads when constructing credit risk measures, assuming no default risk in the US.[1] However, if default risk exists, however small it is, the relative credit risk of one country to the other should be taken into account in examining bilateral currency movements. Here, I relax the assumption of no default in the denominating country and explore whether main results are preserved. Specifically, I define the exchange rate against the Japanese Yen.

[1] I admit that this assumption was partly due to the lack of CDS data for the US.

Cross-country differences in CDS spreads between the home country and Japan are used to construct relative one-year CDSs ($D_t^{1,R} = D_t^1 - D_t^{1,*}$), relative level factors ($L(D_t^R) = L(D_t) - L(D_t^*)$), and relative slope factors($S(D_t^R) = S(D_t) - S(D_t^*)$).[2] Using these relative measures, I check the robustness of findings from the structural break model and the Markov-Switching model. The model specifications are the same except for replacing one-country risk measures with cross-country differences. **Table A.3** and **A.4** present the results consistent with key findings. Whether I assume default risk in the country of safe haven currency or not does not qualitatively affect the conclusions of the study.

I compare the explanatory power and the statistical significance of the level and slope factors with 36-months rolling windows. Numerous regressions are performed for combinations of four horizons ($m = 1-, 3-, 6-, 12-months$), two denominating currencies (USD and JPY) and corresponding credit measures (only own country's credit risk or relative credit risk). **Figure A.1** and **A.2** shows one example out of many results - the result from the regression of the six-month exchange rate against JPY on the relative level and slope factors. All the other results from differently combined models are very similar. Overall, it seems safe to say that risk level counts more, in general, while risk proximity takes a pivotal role during the crisis.

[2]Historical CDS data for Japan is available. The Japanese yen is perceived as a safe haven currency, while the accumulated sovereign debt is the highest in the world, implying potential credit risk. In this regard, Japan is a good sample country for this robustness check.

Table A.1: Robustness Check over Different Horizons: Structural Break Model

$$\begin{aligned}
\textbf{Model 1:} \quad & \Delta s_{t+m} = \beta_0 + \beta_1 D_t^1 + \epsilon_{t+m} \\
\textbf{Model 2:} \quad & \Delta s_{t+m} = \beta_0 + \beta_1 L(D_t) + \epsilon_{t+m} \\
\textbf{Model 3:} \quad & \Delta s_{t+m} = \beta_0 + \beta_1 S(D_t) + \epsilon_{t+m} \\
\textbf{Model 4:} \quad & \Delta s_{t+m} = \beta_0 + \beta_1 L(D_t) + \beta_2 S(D_t) + \epsilon_{t+m}
\end{aligned}$$

	AU	BR	CL	CO	HU	IS	ID	IL	JP	KR
$m = 1$										
Model 1	0.075	0.025	0.151	0.054	0.043	0.245	0.158	0.024	0.031	0.180
Model 2	0.088	0.039	0.184	0.067	0.036	0.250	0.160	0.027	0.022	0.194
Model 3	0.009	0.045	0.158	0.030	0.035	0.262	0.127	0.027	0.051	0.071
Model 4	0.080	0.085	0.187	0.083	0.047	0.257	0.159	0.029	0.050	0.187
$m = 3$										
Model 1	0.338	0.123	0.238	0.151	0.308	0.351	0.190	0.209	0.076	0.259
Model 2	0.278	0.228	0.400	0.270	0.298	0.348	0.168	0.121	0.155	0.344
Model 3	0.229	0.113	0.464	0.119	0.354	0.315	0.333	0.008	0.228	0.411
Model 4	0.601	0.239	0.551	0.326	0.345	0.344	0.368	0.011	0.215	0.469
$m = 6$										
Model 1	0.501	0.331	0.481	0.354	0.547	0.560	0.400	0.285	0.315	0.534
Model 2	0.483	0.485	0.604	0.401	0.528	0.559	0.390	0.207	0.346	0.581
Model 3	0.297	0.317	0.540	0.320	0.519	0.420	0.308	0.279	0.364	0.536
Model 4	0.595	0.545	0.643	0.477	0.553	0.506	0.466	0.323	0.391	0.634
$m = 12$										
Model 1	0.493	0.625	0.671	0.744	0.493	0.673	0.666	0.433	0.484	0.730
Model 2	0.506	0.705	0.649	0.663	0.471	0.659	0.568	0.452	0.535	0.712
Model 3	0.448	0.524	0.380	0.418	0.467	0.514	0.477	0.449	0.478	0.495
Model 4	0.554	0.769	0.672	0.781	0.500	0.670	0.671	0.546	0.555	0.787

Note: 1. For robustness checks, I run the same regression equations in **Table 2.4** over different horizons ($m = 1-, 3-, 6-, 12-$month) and report the adjusted R^2. 2. Four sets of risk factors are considered as regressors and the Bai and Perron (2003) test(with 15% trimming and 5 - 10% significance level) is applied to detect the multiple structural breaks in the regression. After identifying from zero to two breaks, structural break dummy variables for each sub-period are incorporated into the regression.

(Continued)

$$\textbf{Model 1:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 D_t^1 + \epsilon_{t+m}$$
$$\textbf{Model 2:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 L(D_t) + \epsilon_{t+m}$$
$$\textbf{Model 3:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 S(D_t) + \epsilon_{t+m}$$
$$\textbf{Model 4:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 L(D_t) + \beta_2 S(D_t) + \epsilon_{t+m}$$

	MX	NO	PE	PH	PL	RO	ZA	SE	TH	TR
$m = 1$										
Model 1	0.113	0.055	0.095	0.054	0.087	0.053	0.037	0.079	0.047	0.044
Model 2	0.114	0.072	0.080	0.062	0.086	0.073	0.037	0.067	0.033	0.031
Model 3	0.021	0.023	0.056	0.055	0.021	0.040	0.039	0.019	0.020	0.038
Model 4	0.108	0.059	0.082	0.059	0.083	0.075	0.054	0.087	0.048	0.036
$m = 3$										
Model 1	0.140	0.313	0.165	0.278	0.187	0.215	0.176	0.231	0.087	0.110
Model 2	0.128	0.259	0.241	0.266	0.209	0.172	0.257	0.238	-0.004	0.015
Model 3	-0.001	0.210	0.169	0.230	0.197	0.211	0.243	0.188	0.048	0.230
Model 4	0.135	0.380	0.278	0.322	0.312	0.173	0.338	0.173	0.045	0.342
$m = 6$										
Model 1	0.314	0.509	0.352	0.473	0.455	0.398	0.353	0.503	0.210	0.348
Model 2	0.258	0.525	0.385	0.468	0.472	0.368	0.404	0.506	0.247	0.306
Model 3	0.330	0.392	0.375	0.369	0.391	0.365	0.423	0.334	0.291	0.463
Model 4	0.366	0.624	0.444	0.545	0.471	0.417	0.497	0.511	0.270	0.578
$m = 12$										
Model 1	0.532	0.502	0.686	0.711	0.612	0.508	0.500	0.533	0.542	0.548
Model 2	0.483	0.527	0.698	0.699	0.634	0.491	0.591	0.557	0.557	0.472
Model 3	0.437	0.409	0.617	0.535	0.564	0.424	0.460	0.288	0.390	0.453
Model 4	0.568	0.555	0.725	0.732	0.639	0.505	0.681	0.564	0.559	0.582

Note: 1. For robustness checks, I run the same regression equations in **Table 2.4** over different horizons ($m = 1-, 3-, 6-, 12-$month) and report the adjusted R^2. 2. Four sets of risk factors are considered as regressors and the Bai and Perron (2003) test(with 15% trimming and 5 - 10% significance level) is applied to detect the multiple structural breaks in the regression. After identifying from zero to two breaks, structural break dummy variables for each sub-period are incorporated into the regression.

Table A.2: Robustness Check over Different Horizons: Markov Switching Model

$$\Delta s_{t+m} = \beta_{0,\xi_t} + \beta_{1,\xi_t} L(D_t) + \beta_{2,\xi_t} S(D_t) + \epsilon_{t+m},\ where\ \epsilon_{t+m}\ i.i.d.N(0, \sigma^2_{\xi_t})$$

	AU	BR	CL	CO	HU	IS	ID	IL	JP	KR
$\beta_{0,0}$	1.384	-12.973	-54.970***	2.615	-8.234	-14.315*	26.023**	-1.144	1.470	6.369
	(6.184)	(7.984)	(7.098)	(6.651)	(5.882)	(7.424)	(10.161)	(4.974)	(4.902)	(6.101)
$\beta_{0,1}$	123.417**	82.096**	-6.212	53.420	67.195	2.201	-97.073***	-20.680	-9.244	-19.878
	(51.484)	(33.780)	(6.742)	(35.313)	(61.063)	(32.953)	(24.842)	(12.619)	(8.154)	(18.778)
$\beta_{1,0}$	-0.260***	0.008	0.847***	-0.023	-0.029**	0.013*	-0.077***	-0.112	-0.148	-0.126***
	(0.058)	(0.012)	(0.028)	(0.018)	(0.012)	(0.007)	(0.011)	(0.069)	(0.120)	(0.016)
$\beta_{1,1}$	-0.008	-0.115**	-0.047*	-0.107	-0.056	-0.088	0.025*	0.040	0.030	0.173***
	(0.104)	(0.055)	(0.026)	(0.074)	(0.058)	(0.069)	(0.015)	(0.035)	(0.056)	(0.036)
$\beta_{2,0}$	-0.524***	0.024	-1.487***	-0.035	-0.167***	-0.053	-0.004	-0.242	-1.694*	-0.101
	(0.191)	(0.077)	(0.156)	(0.090)	(0.056)	(0.048)	(0.052)	(0.208)	(0.864)	(0.092)
$\beta_{2,1}$	2.243*	0.226	-0.383***	0.137	-0.257	-1.257*	-0.650***	-0.360*	0.064	-0.105
	(1.356)	(0.145)	(0.136)	(0.217)	(0.269)	(0.757)	(0.136)	(0.200)	(0.220)	(0.312)
σ_0	33.541**	31.895*	21.085	26.239**	33.993***	35.297***	20.814**	17.581**	24.115**	22.853**
	(13.541)	(17.139)	(19.498)	(11.974)	(12.661)	(13.349)	(8.359)	(7.599)	(11.221)	(11.408)
σ_1	54.078	67.545**	31.998***	63.728**	89.880*	103.548	27.827*	37.067**	37.081**	40.233*
	(39.964)	(30.734)	(11.279)	(27.917)	(47.810)	(66.021)	(16.336)	(15.633)	(15.058)	(23.643)
P_{00}	0.912***	0.863***	0.309	0.936***	0.966***	0.989**	0.853**	0.967***	0.961***	0.690
	(0.286)	(0.329)	(0.790)	(0.277)	(0.290)	(0.389)	(0.338)	(0.343)	(0.351)	(0.706)
P_{11}	0.301	0.812**	0.977***	0.926***	0.842**	0.868*	0.397	0.964***	0.970***	0.375
	(0.541)	(0.369)	(0.257)	(0.331)	(0.384)	(0.487)	(0.364)	(0.348)	(0.353)	(0.400)
$ad.R^2$	0.493	0.215	0.381	0.042	0.245	0.219	0.626	0.100	0.115	0.528
AIC	10.335	10.737	9.989	10.444	10.510	10.380	9.418	9.502	9.853	9.820

Note: 1. For robustness checks, I estimate the same model in **Table 2.5** over different horizons. Here, estimation results for $m = 1-$month are reported. Results for $m = 6-, 12-$month can be provided upon request. 2. The model is estimated by a two-state Markov-Switching Model. After estimating the parameters by Maximum Likelihood Estimation (MLE), the filtered probabilities are computed by Hamilton filter. 3. Coefficient estimates are reported with the standard errors in parentheses. Asterisks indicate significance levels at 1% (***), 5% (**), and 10% (*) respectively. 4. Adjusted R^2 and AIC are reported.

(Continued)

$$\Delta s_{t+m} = \beta_{0,\xi_t} + \beta_{1,\xi_t} L(D_t) + \beta_{2,\xi_t} S(D_t) + \epsilon_{t+m},\ where\ \epsilon_{t+m}\ i.i.d.N(0, \sigma^2_{\xi_t})$$

	MX	NO	PE	PH	PL	RO	ZA	SE	TH	TR
$\beta_{0,0}$	8.792	-9.806	12.344**	16.989	-13.766*	3.710	-21.006**	-2.227	-5.610	28.997**
	(8.911)	(7.421)	(5.335)	(28.142)	(7.255)	(8.000)	(10.247)	(6.612)	(3.524)	(12.646)
$\beta_{0,1}$	67.938	60.251**	11.925	3.526	90.798	-53.547**	34.683*	47.214	-5.714	179.782
	(49.850)	(30.118)	(26.712)	(4.469)	(131.083)	(22.170)	(19.509)	(39.341)	(5.169)	(223.654)
$\beta_{1,0}$	-0.070***	-0.141	-0.059***	-0.016	-0.052**	-0.028***	-0.068***	-0.203***	-0.001	-0.042*
	(0.015)	(0.154)	(0.018)	(0.024)	(0.025)	(0.010)	(0.025)	(0.077)	(0.019)	(0.023)
$\beta_{1,1}$	-0.013	-0.641	-0.015	-0.008	-0.065	0.142***	-0.021	-0.098	0.004	-0.131
	(0.065)	(0.499)	(0.039)	(0.008)	(0.239)	(0.028)	(0.036)	(0.270)	(0.017)	(0.222)
$\beta_{2,0}$	-0.031	-1.070***	-0.054	0.002	-0.296***	-0.009	-0.119	-0.489*	0.040	0.057
	(0.101)	(0.384)	(0.038)	(0.042)	(0.101)	(0.055)	(0.073)	(0.254)	(0.041)	(0.062)
$\beta_{2,1}$	0.441	0.869	0.074	0.015	0.273	-0.278	-0.080	1.220	-0.129**	0.077
	(0.300)	(1.638)	(0.199)	(0.029)	(0.661)	(0.194)	(0.137)	(1.297)	(0.057)	(0.442)
σ_0	20.146*	29.917**	12.595**	4.497	36.311**	31.275**	26.953*	29.319**	5.495	35.014**
	(11.724)	(12.661)	(5.267)	(3.043)	(15.275)	(13.008)	(15.550)	(13.466)	(3.732)	(14.431)
σ_1	49.380**	56.765*	31.761*	20.966***	77.463	39.139*	53.546**	59.913*	21.454***	74.344
	(23.679)	(31.337)	(16.734)	(7.128)	(48.287)	(21.681)	(23.187)	(34.948)	(8.049)	(59.695)
P_{00}	0.933**	0.947***	0.975***	0.912*	0.943**	0.758***	0.308	0.931**	0.508	0.955**
	(0.459)	(0.286)	(0.338)	(0.527)	(0.431)	(0.283)	(0.472)	(0.393)	(0.358)	(0.452)
P_{11}	0.871*	0.847**	0.925**	0.999	0.645	0.235	0.515	0.777*	0.843***	0.374
	(0.500)	(0.421)	(0.417)	(4.910)	(0.499)	(0.457)	(0.340)	(0.470)	(0.248)	(0.977)
$ad.R^2$	0.249	0.136	0.030	-0.018	0.379	0.522	0.564	0.174	0.091	0.405
AIC	9.784	10.208	8.575	8.840	10.549	10.234	10.840	10.213	8.757	10.335

Note: 1. For robustness checks, I estimate the same model in **Table 2.5** over different horizons. Here, estimation results for $m = 1-$month are reported. Results for $m = 6-, 12-$month can be provided upon request. 2. The model is estimated by a two-state Markov-Switching Model. After estimating the parameters by Maximum Likelihood Estimation (MLE), the filtered probabilities are computed by Hamilton filter. 3. Coefficient estimates are reported with the standard errors in parentheses. Asterisks indicate significance levels at 1% (***), 5% (**), and 10% (*) respectively. 4. Adjusted R^2 and AIC are reported.

Table A.3: Robustness Check with Relative Risk Measures: Structural Break Model

$$\textbf{Model 1:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 D_t^{1,R} + \epsilon_{t+m}$$
$$\textbf{Model 2:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 L(D_t^R) + \epsilon_{t+m}$$
$$\textbf{Model 3:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 S(D_t^R) + \epsilon_{t+m}$$
$$\textbf{Model 4:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 L(D_t^R) + \beta_2 S(D_t^R) + \epsilon_{t+m}$$

	AU	BR	CL	CO	HU	IS	ID	IL	KR	MX
$m = 1$										
Model 1	0.012	0.014	0.120	0.009	0.062	0.228	0.092	0.061	0.185	0.051
Model 2	0.021	0.011	0.179	0.011	0.069	0.234	0.087	0.080	0.189	0.031
Model 3	0.043	0.046	0.116	0.026	0.048	0.266	0.061	0.046	0.051	0.020
Model 4	0.050	0.040	0.175	0.035	0.065	0.256	0.082	0.098	0.180	0.055
$m = 3$										
Model 1	0.047	0.055	0.160	0.029	0.416	0.313	0.179	0.286	0.263	0.066
Model 2	0.078	0.062	0.231	0.043	0.367	0.318	0.177	0.291	0.307	0.047
Model 3	0.100	0.062	0.309	0.074	0.390	0.319	0.093	0.106	0.367	0.040
Model 4	0.095	0.117	0.322	0.079	0.445	0.323	0.268	0.287	0.374	0.054
$m = 6$										
Model 1	0.232	0.241	0.320	0.264	0.612	0.484	0.357	0.423	0.482	0.254
Model 2	0.322	0.261	0.368	0.264	0.588	0.522	0.379	0.366	0.475	0.210
Model 3	0.328	0.251	0.298	0.299	0.526	0.521	0.322	0.281	0.533	0.192
Model 4	0.358	0.322	0.420	0.333	0.616	0.572	0.394	0.359	0.556	0.318
$m = 12$										
Model 1	0.467	0.447	0.429	0.439	0.728	0.701	0.603	0.650	0.600	0.426
Model 2	0.541	0.495	0.475	0.351	0.723	0.710	0.605	0.611	0.587	0.486
Model 3	0.526	0.387	0.263	0.388	0.553	0.630	0.500	0.456	0.610	0.363
Model 4	0.613	0.601	0.476	0.626	0.733	0.753	0.649	0.607	0.670	0.497

Note: 1. For robustness checks, I run the same regression equations in **Table 2.4** for different currency pairs with relative credit risk measures over different horizons ($m = 1-, 3-, 6-, 12-$month) and report the adjusted R^2. 2. Key differences are: (i) the exchange rate is defined as a per-Japanese Yen rate instead of a per-US Dollar rate, (ii) relative sovereign credit risk measures are used as explanatory variables. Relative CDS spreads are defined as $D_t^R = D_t - D_t^*$. Relative level, $L(D_t^R)$ and relative slope, $S(D_t^R)$ are defined accordingly. 3. Bai-Perron(2003) test(with 15% trimming and 5 - 10% significance level) is applied to detect the multiple structural breaks in the regression. After identifying from zero to two break dates, structural break dummy variables for each sub-period are incorporated into the regression.

(Continued)

$$\textbf{Model 1:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 D_t^{1,R} + \epsilon_{t+m}$$
$$\textbf{Model 2:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 L(D_t^R) + \epsilon_{t+m}$$
$$\textbf{Model 3:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 S(D_t^R) + \epsilon_{t+m}$$
$$\textbf{Model 4:}\quad \Delta s_{t+m} = \beta_0 + \beta_1 L(D_t^R) + \beta_2 S(D_t^R) + \epsilon_{t+m}$$

	NO	PE	PH	PL	RO	ZA	SE	TH	TR	US
$m = 1$										
Model 1	0.028	0.055	0.009	0.084	0.067	0.029	0.021	0.071	0.007	0.112
Model 2	0.025	0.053	0.020	0.096	0.106	0.022	0.018	0.064	0.003	0.082
Model 3	0.006	0.034	0.010	0.024	0.059	0.023	0.012	0.032	0.029	0.036
Model 4	0.015	0.069	0.024	0.091	0.136	0.022	0.026	0.056	0.023	0.074
$m = 3$										
Model 1	0.103	0.192	0.081	0.256	0.243	0.050	0.046	0.137	0.004	0.129
Model 2	0.144	0.144	0.132	0.238	0.251	0.027	0.018	0.236	0.096	0.116
Model 3	0.190	0.006	0.107	0.205	0.212	0.262	0.267	0.209	0.362	0.118
Model 4	0.211	0.143	0.157	0.298	0.287	0.325	0.255	0.266	0.364	0.110
$m = 6$										
Model 1	0.244	0.316	0.269	0.448	0.448	0.368	0.254	0.295	0.274	0.349
Model 2	0.129	0.256	0.239	0.443	0.446	0.337	0.278	0.289	0.240	0.395
Model 3	0.294	0.237	0.297	0.338	0.367	0.416	0.347	0.325	0.520	0.444
Model 4	0.350	0.369	0.337	0.470	0.491	0.438	0.421	0.320	0.520	0.438
$m = 12$										
Model 1	0.376	0.369	0.397	0.635	0.658	0.382	0.416	0.398	0.378	0.679
Model 2	0.386	0.421	0.379	0.604	0.659	0.397	0.480	0.398	0.358	0.690
Model 3	0.386	0.569	0.486	0.510	0.556	0.307	0.421	0.406	0.376	0.689
Model 4	0.520	0.595	0.508	0.646	0.721	0.497	0.520	0.405	0.398	0.694

Note: 1. For robustness checks, I run the same regression equations in **Table 2.4** for different currency pairs with relative credit risk measures over different horizons ($m = 1-, 3-, 6-, 12-$month) and report the adjusted R^2. 2. Key differences are: (i) the exchange rate is defined as a per-Japanese Yen rate instead of a per-US Dollar rate, (ii) relative sovereign credit risk measures are used as explanatory variables. Relative CDS spreads are defined as $D_t^R = D_t - D_t^*$. Relative level, $L(D_t^R)$ and relative slope, $S(D_t^R)$ are defined accordingly. 3. Bai-Perron(2003) test(with 15% trimming and 5 - 10% significance level) is applied to detect the multiple structural breaks in the regression. After identifying from zero to two break dates, structural break dummy variables for each sub-period are incorporated into the regression.

Table A.4: Robustness Check with Relative Risk Measures: Markov Switching Model

$$\Delta s_{t+3} = \beta_{0,\xi_t} + \beta_{1,\xi_t} L(D_t^R) + \beta_{2,\xi_t} S(D_t^R) + \epsilon_{t+3},\ where\ \epsilon_{t+3}\ i.i.d.N(0, \sigma_{\xi_t}^2)$$

	AU	BR	CL	CO	HU	IS	ID	IL	KR	MX
$\beta_{0,0}$	-1.500	-3.884	194.793***	12.164*	-1.144	-10.159	8.977	-4.538	-7.048*	7.013
	(5.150)	(6.256)	(4.770)	(6.577)	(4.259)	(82.634)	(5.522)	(5.852)	(4.061)	(8.241)
$\beta_{0,1}$	145.383***	51.372*	-1.544	-18.630	121.121***	-42.605	69.980**	8.573	16.998	61.839
	(35.944)	(27.257)	(5.665)	(23.589)	(31.544)	(14246)	(32.406)	(13.692)	(14.873)	(56.484)
$\beta_{1,0}$	-0.108	-0.016	-0.240***	-0.073***	-0.034***	0.016	-0.032***	-0.028	0.025	-0.024
	(0.164)	(0.012)	(0.015)	(0.019)	(0.009)	(0.075)	(0.012)	(0.019)	(0.018)	(0.027)
$\beta_{1,1}$	-1.271	0.026	-0.042	0.134	-0.114***	-0.061	-0.156***	0.025	-0.016	0.018
	(1.827)	(0.056)	(0.044)	(0.088)	(0.029)	(2.347)	(0.014)	(0.090)	(0.049)	(0.053)
$\beta_{2,0}$	0.264	0.041	0.718***	-0.014	-0.063	0.004	0.000	0.048	0.107	0.108
	(0.561)	(0.076)	(0.209)	(0.064)	(0.049)	(0.389)	(0.003)	(0.151)	(0.088)	(0.107)
$\beta_{2,1}$	0.410	-0.472*	-0.216	-0.158	-0.014	-1.067	-0.685***	0.081	-0.032	0.145
	(5.670)	(0.265)	(0.194)	(0.335)	(0.087)	(68.813)	(0.163)	(0.332)	(0.318)	(0.421)
σ_0	22.621**	24.317**	0.257*	17.910**	21.485**	21.284	21.835***	11.052	10.848**	24.080**
	(9.772)	(9.960)	(0.150)	(7.286)	(8.943)	(20.940)	(8.443)	(6.701)	(4.896)	(10.201)
σ_1	39.421	44.554	23.703**	44.976*	34.237**	102.425	24.980**	35.177*	42.980	45.227
	(38.186)	(30.309)	(9.302)	(24.445)	(15.273)	(1047)	(11.132)	(19.274)	(31.069)	(36.621)
P_{00}	0.993***	0.957***	0.404	0.902***	0.976***	0.989	0.993***	0.905***	0.927***	0.971***
	(0.352)	(0.237)	(0.736)	(0.208)	(0.237)	(8.872)	(0.359)	(0.234)	(0.240)	(0.242)
P_{11}	0.888	0.789***	0.991***	0.793***	0.862**	0.898	0.817*	0.896***	0.865**	0.818**
	(0.830)	(0.293)	(0.308)	(0.209)	(0.352)	(11.933)	(0.478)	(0.263)	(0.391)	(0.352)
$ad.R^2$	0.533	0.567	0.441	0.329	0.576	0.139	0.453	0.092	0.248	0.444
AIC	9.284	9.782	9.181	9.611	9.418	9.514	9.187	9.125	8.993	9.600

Note: 1. For robustness checks, I estimate the same model in **Table 2.5** for different currency pairs with relative credit risk measures over different horizons ($m = 1-, 3-, 6-, 12-$month): (i) the exchange rate is defined as a per-Japanese Yen rate instead of a per-US Dollar rate, (ii) relative sovereign credit risk measures are used as explanatory variables. Relative CDS spreads are defined as $D_t^R = D_t - D_t^*$. Relative level, $L(D_t^R)$ and relative slope, $S(D_t^R)$ are defined accordingly. 2. Here, estimation results for $m = 3-$month are reported. Results for $m = 1-, 6-, 12-$month can be provided upon request.

(Continued)

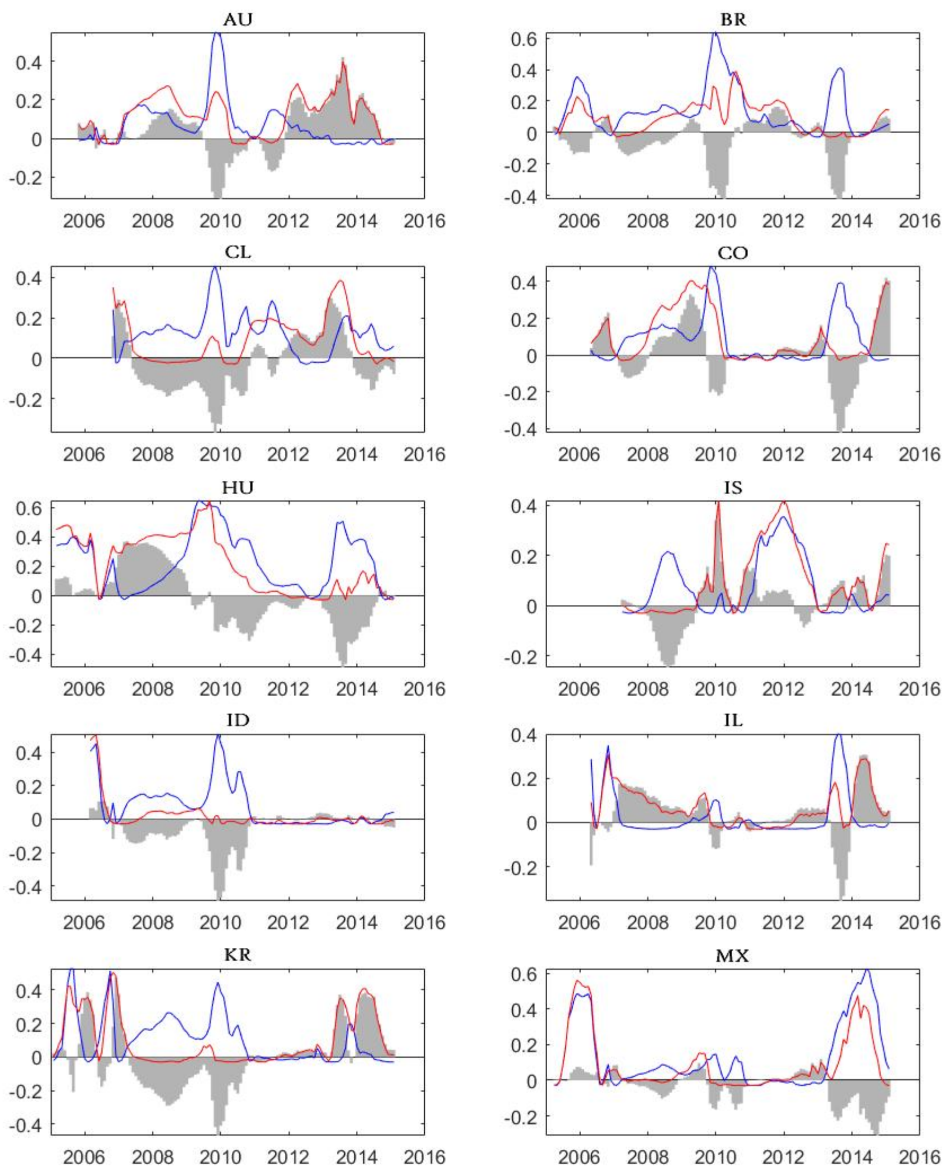
$$\Delta s_{t+3} = \beta_{0,\xi_t} + \beta_{1,\xi_t} L(D_t^R) + \beta_{2,\xi_t} S(D_t^R) + \epsilon_{t+3},\ where\ \epsilon_{t+3}\ i.i.d.N(0, \sigma_{\xi_t}^2)$$

	NO	PE	PH	PL	RO	ZA	SE	TH	TR	US
$\beta_{0,0}$	162.424***	10.026**	-1.182	-5.646	-7.853	5.297	-4.685	-29.825***	36.254***	-4.123
	(3.182)	(4.119)	(14.992)	(6.147)	(6.481)	(9.448)	(4.344)	(8.468)	(9.766)	(2.736)
$\beta_{0,1}$	-1.646	-53.181***	5.532	51.285	23.095	80.858**	59.802	6.674	-142.800***	7.919
	(3.997)	(15.157)	(67.009)	(87.871)	(14.330)	(39.488)	(36.439)	(8.326)	(20.944)	(6.793)
$\beta_{1,0}$	1.010***	-0.003	0.005	-0.063***	-0.002	-0.077***	0.058	0.268***	-0.070***	-0.043
	(0.038)	(0.020)	(0.018)	(0.023)	(0.009)	(0.029)	(0.075)	(0.051)	(0.018)	(0.048)
$\beta_{1,1}$	-0.005	0.053	0.009	0.018	0.057*	-0.095	-0.015	-0.051***	0.175***	0.130
	(0.054)	(0.054)	(0.119)	(0.237)	(0.031)	(0.062)	(0.066)	(0.018)	(0.021)	(0.085)
$\beta_{2,0}$	-1.795***	0.119*	0.081	0.094	0.066	-0.092	-0.069	0.124	0.050	-0.157**
	(0.178)	(0.062)	(0.123)	(0.164)	(0.074)	(0.075)	(0.225)	(0.146)	(0.056)	(0.078)
$\beta_{2,1}$	-0.046	-0.555**	-0.253	0.104	0.069	-0.098	1.559*	-0.126	-0.922***	-0.042
	(0.274)	(0.225)	(0.748)	(0.445)	(0.206)	(0.276)	(0.906)	(0.105)	(0.074)	(0.252)
σ_0	5.680	12.532**	18.654	18.352	18.825**	22.799**	18.289**	20.876*	24.901**	7.925
	(5.165)	(5.045)	(12.014)	(13.008)	(7.913)	(9.518)	(8.878)	(12.580)	(11.505)	(5.142)
σ_1	23.682**	34.505	21.708	43.676	31.302	39.292	36.477*	11.632	25.385	26.369**
	(9.393)	(21.867)	(13.470)	(37.754)	(21.472)	(27.766)	(19.814)	(7.429)	(16.061)	(11.743)
P_{00}	0.590	0.956***	0.972	0.950**	0.924***	0.958***	0.944***	0.880***	0.957***	0.795***
	(0.464)	(0.247)	(1.335)	(0.412)	(0.175)	(0.271)	(0.242)	(0.262)	(0.280)	(0.212)
P_{11}	0.961***	0.860***	0.895	0.811***	0.822***	0.805***	0.827***	0.845***	0.744**	0.874***
	(0.171)	(0.271)	(0.903)	(0.305)	(0.217)	(0.275)	(0.259)	(0.225)	(0.373)	(0.277)
$ad.R^2$	0.500	0.162	0.394	0.592	0.567	0.583	0.497	0.464	0.567	0.066
AIC	9.302	8.683	8.966	9.439	9.481	9.649	9.403	8.797	9.603	8.902

Note: 1. For robustness checks, I estimate the same model in **Table 2.5** for different currency pairs with relative credit risk measures over different horizons ($m = 1-, 3-, 6-, 12-$month): (i) the exchange rate is defined as a per-Japanese Yen rate instead of a per-US Dollar rate, (ii) relative sovereign credit risk measures are used as explanatory variables. Relative CDS spreads are defined as $D_t^R = D_t - D_t^*$. Relative level, $L(D_t^R)$ and relative slope, $S(D_t^R)$ are defined accordingly. 2. Here, estimation results for $m = 3-$month are reported. Results for $m = 1-, 6-, 12-$month can be provided upon request.

Figure A.1: Robustness Check with Relative Risk Measure over Different Horizons: Comparing the Ad.R^2 over Rolling Windows

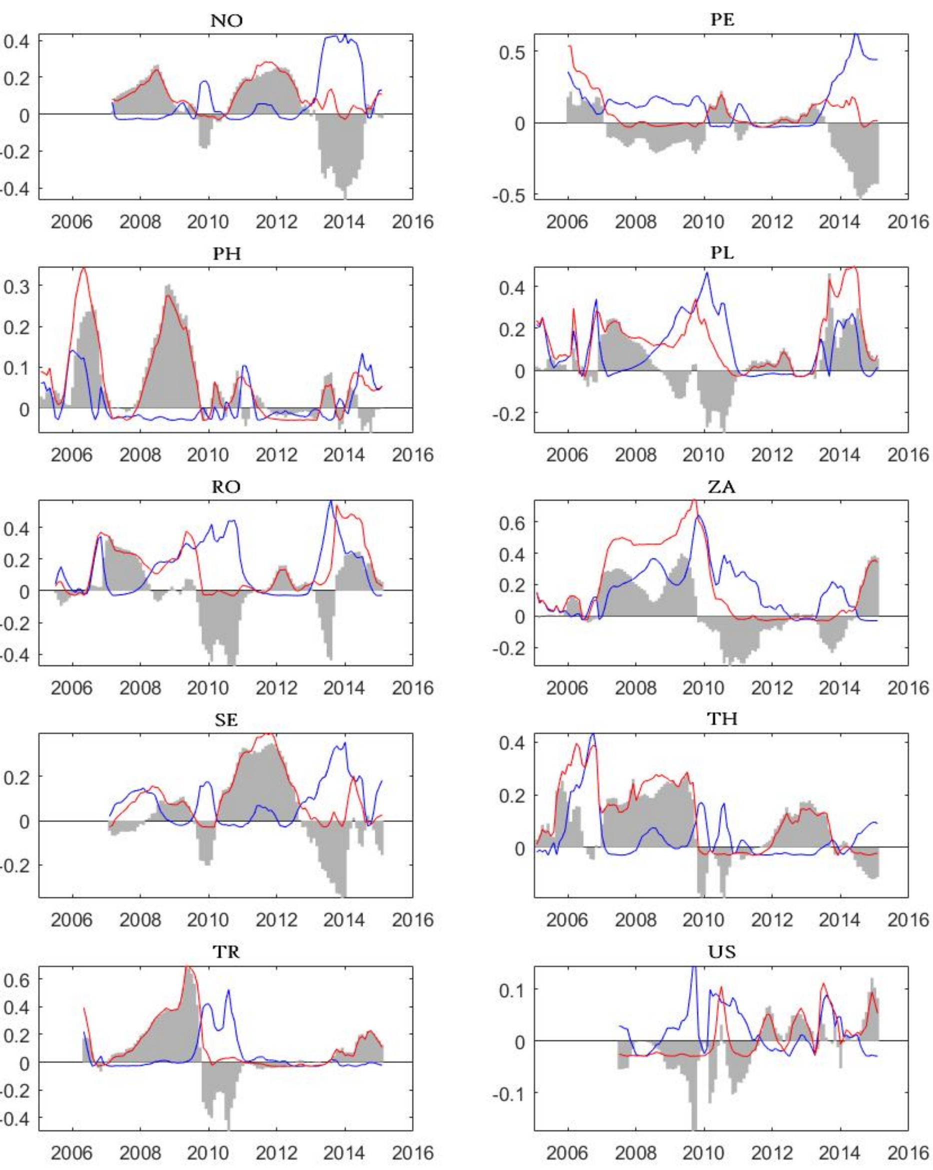
(Adjusted R^2)



Note: 1. This figure is for the robustness check for **Figure 2.4**: 1) Relative level-only model: $\Delta s_{t+m} = \beta_0 + \beta_1 L(D_t^R) + \epsilon_{t+m}$; 2) Relative slope-only model: $\Delta s_{t+m} = \beta_0 + \beta_1 S(D_t^R) + \epsilon_{t+m}$. The results are from JPY as the denominating currency, with relative risk measures as regressors, and for $m = 6-$month. Other results from USD and over different horizons ($m = 1-, 3-, 12-$month) can be provided upon request. 2. Blue line: ad.R^2 from the relative level-only model; Red line: ad.R^2 from the relative slope-only model; Gray bar: ad.R^2 from the relative slope-only model $-$ ad.R^2 from the relative level-only model.

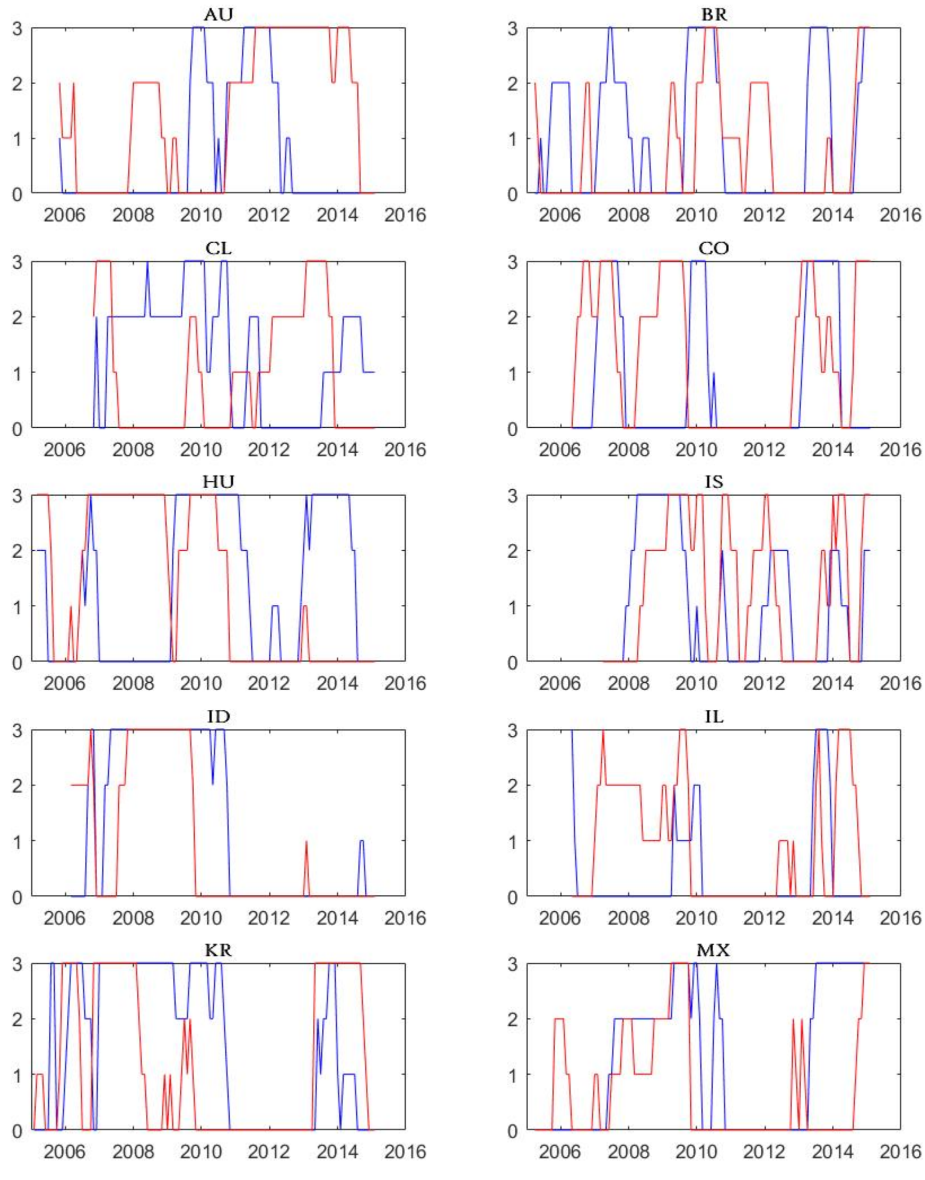
(Continued)

(Adjusted R^2)



Note: 1. This figure is for the robustness check for **Figure 2.4**: 1) Relative level-only model: $\Delta s_{t+m} = \beta_0 + \beta_1 L(D_t^R) + \epsilon_{t+m}$; 2) Relative slope-only model: $\Delta s_{t+m} = \beta_0 + \beta_1 S(D_t^R) + \epsilon_{t+m}$. The results are from JPY as the denominating currency, with relative risk measures as regressors, and for $m = 6-$month. Other results from USD and over different horizons ($m = 1-, 3-, 12-$month) can be provided upon request. 2. Blue line: ad.R^2 from the relative level-only model; Red line: ad.R^2 from the relative slope-only model; Gray bar: ad.R^2 from the relative slope-only model $-$ ad.R^2 from the relative level-only model.

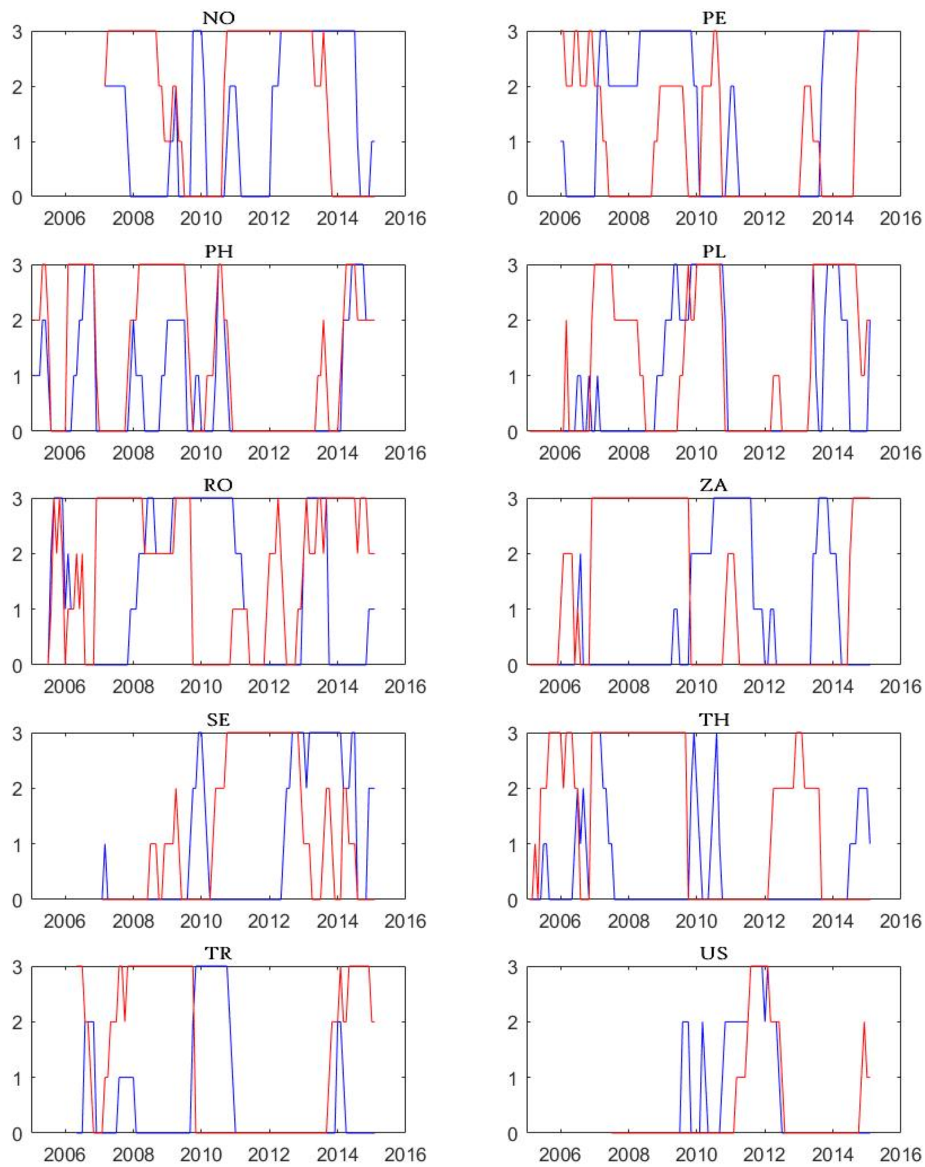
Figure A.2: Robustness Check with Relative Risk Measure over Different Horizons: Comparing the $P-values$ over Rolling Windows
$(P-values)$



Note: 1. This figure is for the robustness check for **Figure 2.5**. From the regression of $\Delta s_{t+m} = \beta_0 + \beta_1 L(D_t^R) + \beta_1 S(D_t^R) + \epsilon_{t+m}$, I obtain $P-values$ for the level and slope factors. If $P-value \leq 0.01$, I give "3". If $0.01 < P-value \leq 0.05$, I give "2". If $0.05 < P-value \leq 0.1$, I give "1". Otherwise, "0". Here, I show the results from JPY as the denominating currency, with relative risk measures as regressors, and for $m = 6-$month. The results from USD and/or over different horizons ($m = 1-, 3-, 12-$month) can be provided upon request. 2. Blue line for the relative level factor and Red line for the relative slope factor.

(Continued)

$(P - values)$



Note: 1. This figure is for the robustness check for **Figure 2.5**. From the regression of $\Delta s_{t+m} = \beta_0 + \beta_1 L(D_t^R) + \beta_1 S(D_t^R) + \epsilon_{t+m}$, I obtain $P - values$ for the level and slope factors. If $P - value \leq 0.01$, I give "3". If $0.01 < P - value \leq 0.05$, I give "2". If $0.05 < P - value \leq 0.1$, I give "1". Otherwise, "0". Here, I show the results from JPY as the denominating currency, with relative risk measures as regressors, and for $m = 6-$month. The results from USD and/or over different horizons ($m = 1-, 3-, 12-$month) can be provided upon request. 2. Blue line for the relative level factor and Red line for the relative slope factor.

Appendix B

APPENDIX TO CHAPTER 3

B.1 Decomposing the Yield Curve into Expectations and Risk

B.1.1 Nelson-Siegel and Nelson-Siegel augmented with Macro model

The Nelson-Siegel latent factor framework (Nelson and Siegel, 1987) provides a succinct summary of the few sources of systematic risks that underlie the pricing of various tradable financial assets.[1] The classic Nelson-Siegel model summarizes the shape of the yield curve using three factors: L_t (*level*), S_t (*slope*), and C_t (*curvature*). Compared to the no-arbitrage affine or quadratic factor models, these factors are easy to estimate, can capture the various shapes of the empirically observed yield curves, and have simple intuitive interpretations.[2] The three factors typically account for most of the information in a yield curve, with the R^2 for cross-sectional fits around 0.99. While the more structural no-arbitrage factor models also fit cross-sectional data well, they do not provide as good a description of the dynamics of the yield curve over time.[3]

We extract *relative expected yields* and *relative term premiums* from the cross-country yield curves using the Nelson-Siegel framework, based on Diebold, Rudebusch, and Aruoba (2006). First, assuming that the factors follow a VAR(1), we simultaneously fit the yield

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[1]Since the Nelson-Siegel framework is by now well-known, we refer interested readers to Chen and Tsang (2013) and references therein for a more detailed presentation of it.

[2]The *level* factor L_t, with its loading of unity, has equal impact on the entire yield curve, shifting it up or down. The loading on the *slope* factor S_t equals 1 when $m = 0$ and decreases down to zero as maturity m increases. The *slope* factor thus mainly affects yields on the short end of the curve; an increase in the *slope* factor means the yield curve becomes flatter, holding the long end of the yield curve fixed. The *curvature* factor C_t is a "medium" term factor, as its loading is zero at the short end, increases in the middle maturity range, and finally decays back to zero. It captures the *curvature* of the yield curve is at medium maturities. See Chen and Tsang (2013) and references therein.

[3]See, e.g., Diebold, Rudebusch, and Aruoba (2006) and Duffee (2002).

curve for each country at each point in time and estimate the underlying dynamics of Nelson-Siegel factors by employing the state-space and Kalman filter approach. As discussed in Diebold, Rudebusch, and Aruoba (2006), this one-step approach improves upon the two-step estimation procedure of Diebold and Li (2006).[4] In this study, we explore two different dynamics of yield factors: one is a dynamics of yield factors only (NS model) and the other is a joint dynamics of yield factors and macro fundamentals such as unemployment rate and inflation rate (NSM model). Specifically, even though two models share the same measurement equation, which relates a set of N yields to state vector, the transition equation follows different dynamics. As such, the state-space model representation is as follows.

$$\begin{aligned} (Measurement Equation) \quad & i_t && = \Lambda f_t + e_t \\ (Transition Equation) \quad & f_t - \mu && = A(f_{t-1} - \mu) + v_t \\ & \begin{pmatrix} e_t \\ v_t \end{pmatrix} && \sim i.i.d.N\left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} H & 0 \\ 0 & Q \end{pmatrix} \right) \end{aligned} \tag{B.1}$$

$$\begin{aligned} where\ i_t &= (i_t^{m_1}, i_t^{m_2}, ..., i_t^{m_N})', \\ f_t &= (L_t, S_t, C_t)' in\ the\ NS\ model, \\ f_t &= (u_t, \pi_t, L_t, S_t, C_t)' in\ the\ NSM\ model, \\ \Lambda &= \begin{pmatrix} 1 & \frac{1-\exp(-\lambda m_1)}{\lambda m_1} & \frac{1-\exp(-\lambda m_1)}{\lambda m_1} - \exp(-\lambda m_1) \\ 1 & \frac{1-\exp(-\lambda m_2)}{\lambda m_2} & \frac{1-\exp(-\lambda m_2)}{\lambda m_2} - \exp(-\lambda m_2) \\ \vdots & \vdots & \vdots \\ 1 & \frac{1-\exp(-\lambda m_N)}{\lambda m_N} & \frac{1-\exp(-\lambda m_N)}{\lambda m_N} - \exp(-\lambda m_N) \end{pmatrix} in\ the\ NS\ model, \\ \Lambda &= \begin{pmatrix} 0 & 0 & 1 & \frac{1-\exp(-\lambda m_1)}{\lambda m_1} & \frac{1-\exp(-\lambda m_1)}{\lambda m_1} - \exp(-\lambda m_1) \\ 0 & 0 & 1 & \frac{1-\exp(-\lambda m_2)}{\lambda m_2} & \frac{1-\exp(-\lambda m_2)}{\lambda m_2} - \exp(-\lambda m_2) \\ \vdots & \vdots & \vdots & \vdots & \vdots \\ 0 & 0 & 1 & \frac{1-\exp(-\lambda m_N)}{\lambda m_N} & \frac{1-\exp(-\lambda m_N)}{\lambda m_N} - \exp(-\lambda m_N) \end{pmatrix} in\ the\ NSM\ model \end{aligned}$$

[4] In Diebold and Li (2006), with λ held fixed at 0.0609, the *level*, *slope*, and *curvature* parameters for each monthly yield curve are estimated. This process is repeated for all observed yield curves, and provides a time series of estimates of the unobserved *level*, *slope*, and *curvature* factors. Then, it fits a first-order autoregressive model to the time series of factors derived in the first step.

By applying the Kalman filter, we obtain the maximum likelihood estimates for parameters including VAR coefficients and optimal filtered and smoothed estimates of yield factors.[5] Next, iterating the estimated VAR, we can calculate in-sample forecasts of the factors:

$$\widehat{f_{t+j}} - \hat{\mu} = \hat{A}^j(\hat{f}_t - \hat{\mu}), \; for \; j = 1, 2, ..., \infty, \; for \; all \; t \tag{B.2}$$

Using the Nelson-Siegel formula with $m = 1$, we obtain the predicted 1-month yield over future horizons:

$$E_t i^1_{t+j} = \hat{L}_{t+j} + \hat{S}_{t+j}\left(\frac{1-\exp(-\lambda)}{\lambda}\right) + \hat{C}_{t+j}\left(\frac{1-\exp(-\lambda)}{\lambda} - \exp(-\lambda)\right) \tag{B.3}$$

The *expected yield* of maturity m is defined as the average of the expected 1-month yields over $m-1$ maturities and the *term premiums* of the same maturity is calculated by subtracting the *expected yield* from the fitted yield:

$$E_t i^m_t = \frac{1}{m}\sum_{j=0}^{m-1} E_t\left[i^1_{t+j}\right] \tag{B.4}$$

$$\theta^m_t = \hat{i}^m_t - E_t i^m_t \tag{B.5}$$

Finally, *relative expected yields* and *relative term premiums* are calculated by the differences in *expected yields* and *term premiums* between each country-pair:

$$E_t i^{R,m}_t = E_t i^m_t - E_t i^{m,*}_t \tag{B.6}$$

$$\theta^{R,m}_t = \theta^m_t - \theta^{m,*}_t \tag{B.7}$$

[5]For details of Kalman filtering and related issues, see Diebold, Rudebusch, and Aruoba (2006).

B.1.2 Affine model

We follow the discrete-time affine Gaussian term structure model, based on Ang and Piazzesi (2003). Let P_t^m denote the price at time t of an m-period zero-coupon bond. For a zero-couple bond, we know $i_t^m = -\log(P_t^m)/m$. Under no-arbitrage, the price of the bond should be consistent with the pricing kernel that $P_t^m = E_t(\prod_{j=1}^m M_{t+j})$, where the pricing kernel M_{t+1} is conditionally lognormal:

$$M_{t+1} = \exp\left(-i_t^1 - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t'\epsilon_{t+1}\right),\ \epsilon_t \sim i.i.d.N(0, I) \tag{B.8}$$

The term $\lambda_t = \lambda_0 + \lambda_1 X_t$ is a time-varying market price of risk, where X_t is the state variables, which in our case are the three principal components of the yields. The one-period yield i_t^i is also an affine function of the state variables, $i_t^1 = \delta_0 + \delta_1' X_t$. The state variables follow a first-order Gaussian VAR. Under objective probability measure P:

$$X_{t+1} = \mu + \Phi X_t + \Sigma\epsilon_{t+1} \tag{B.9}$$

Under risk-neutral probability measure Q:

$$X_{t+1} = \mu^Q + \Phi^Q X_t + \Sigma\epsilon_{t+1}^Q \tag{B.10}$$

where $\mu^Q = \mu - \Sigma\lambda_0,\ \Phi^Q = \Phi - \Sigma\lambda_1$.

Using the log-normality assumption and the VAR model for the state variables X_t, we can express bond prices and yields as a function of the state variables and other parameters. The model-implied bond prices and yields are:

$$P_t^m = \exp\left(A_m + B_m' X_t\right),\ i_t^m = -\log(P_t^m)/m \tag{B.11}$$

where $A_{m+1} = A_m + B_m'\mu^Q + \frac{1}{2}B_m'\Sigma\Sigma' B_m - \delta_0, B_{m+1} = \Phi^Q B_m - \delta_1$

The risk-neutral bond prices and yields are:

$$\tilde{P}_t^m = \exp\left(A_m + B_m' X_t\right),\ \tilde{i}_t^m = -\log(\tilde{P}_t^m)/m \tag{B.12}$$

where $\tilde{A}_{m+1} = \tilde{A}_m + \tilde{B}_m'\mu + \frac{1}{2}\tilde{B}_m'\Sigma\Sigma'\tilde{B}_m - \delta_0, \tilde{B}_{m+1} = \Phi\tilde{B}_m - \delta_1$

That is, bond prices and yields are determined recursively through A_m and B_m. And the model-implied bond prices and yields are calculated as if agents are risk-neutral ($\lambda_0 = \lambda_1 = 0$) but the state variables follow a different law of motion.

When estimating the model, considering small sample bias due to high persistence of interest rate, we follow the bias correction method proposed by Bauer et al. (2012). The idea is that after estimating μ and Φ under P measure by ordinary least squares (OLS), it corrects the possible mean bias by bootstrap bias correction and indirect inference. Then, given bias-corrected μ and Φ, the remaining parameters $(\mu_Q, \Phi_Q, \Sigma, \delta_0, \delta_1)$ are estimated by maximum likelihood estimation (MLE).[6] The estimated risk-neutral yields $(\tilde{i}_t^m)$ are used as *expected yields* and the differences between model-implied yields and risk-neutral yields $(i_t^m - \tilde{i}_t^m)$ are defined as *term premiums*. Again, for our cross-country analysis, *relative expected yields* and *relative term premiums* are obtained by the cross-country differences.

B.1.3 Cochrane-Piazzesi model

Following Cochrane and Piazzesi (2005) we look at annual excess returns. Holding period return of a 12n-month bond from now to next year can be calculated as:

$$r_{t+12}^{12n} \equiv n i_t^{12n} - (n-1) i_{t+12}^{12(n-1)}, where\ n = 2, 3, ..., 10 \tag{B.13}$$

That is, you buy the $12n$-month bond now and sell it as a $12(n-1)$-month bond next year. The above defines the return of such a transaction. Excess return is then defined as

[6] For details of bias correction estimation for the Affine model, see bauer2012correcting.

$$rx_{t+12}^{12n} \equiv r_{t+12}^{12n} - i_t^{12} \tag{B.14}$$

The term tells you the extra return you get from the transaction over a riskless $12n$-month bond. In the data, we have 12, 24, ... , 120-month bonds. The ten yields allow us to define $rx_{t+12}^{24}, ..., rx_{t+12}^{120}$, a total of nine excess returns. The forward rate at time t for loans between time $t+n-1$ and $t+n$ is also defined as

$$f_t^{12n} \equiv n i_t^{12n} - (n-1) i_t^{12(n-1)} \tag{B.15}$$

The CP regression involves regressing the bond excess returns on the 12-month yield and all the forward rates:[7]

$$rx_{t+12}^{12n} = \gamma_0 + \gamma_1 i_t^{12} + \gamma_2 f_t^{24} + ... + \gamma_{10} f_t^{120} + \epsilon_{t+12}, \ for\ n = 2, 3, ..., 10 \tag{B.16}$$

We borrow the concept of the CP regression, but modify it to a VAR(1) representation in order to obtain the predicted excess returns in the future periods. Specifically, we estimate the following VAR(1) system:

$$CP_{t+1} - \mu = A(CP_t - \mu) + e_t \tag{B.17}$$

where $CP_t = (rx_{t+12}^{12n}, i_t^{12}, f_t^{24}, ..., f_t^{120})'$, for $n = 2, 3, ..., 10$

After estimating the parameters, we iterate the VAR to obtain in-sample forecasts of the excess returns at any horizons:

$$\widehat{CP}_{t+12h} - \hat{\mu} = \hat{A}^{12h}(\widehat{CP}_t - \hat{\mu}), \ for\ h = 1, 2, ..., \infty, \ for\ all\ t \tag{B.18}$$

[7] In Cochrane and Piazzesi (2005), it runs a regression of the average excess return on the 12-month yield and all the forward rates, in the sense that expected excess returns of all maturities share the same single factor, and calls this single-factor model as a restricted model. Here, since we want to derive the term premiums at different maturities from the excess returns, we rely on an unrestricted model in which the excess returns at different maturities are assumed to have different factors.

The expected excess returns with $12n-month$ maturity bond at $t+12h$ is the first column of $\widehat{CP}_{t+12h}$. Next, by using the relationship between expected excess returns and *term premiums*, we calculate the *term premiums*:

$$\theta_t^{12n} = \frac{1}{n}\sum_{h=0}^{n-2} E_t\left(rx_{t+12(h+1)}^{12(n-h)}\right) \tag{B.19}$$

We can easily prove this relationship by the definitions of excess returns and term premiums:

$$\begin{aligned}
< proof > RHS &= \frac{1}{n}\left[E_t(rx_{t+12}^{12n}) + E_t(rx_{t+24}^{12(n-1)}) + ... + E_t(rx_{t+12(n-1)}^{24})\right] \\
&= \frac{1}{n}\left[\left(ni_t^{12n} - (n-1)E_t(i_{t+12}^{12(n-1)}) - i_t^{12}\right)\right] \\
&+ \frac{1}{n}\left[\left((n-1)E_t(i_{t+12}^{12(n-1)}) - (n-2)E_t(i_{t+24}^{12(n-2)}) - E_t(i_{t+12}^{12})\right)\right] \\
&+ ... + \frac{1}{n}\left[\left(2E_t(i_{t+12(n-2)}^{24}) - E_t(i_{t+12(n-1)}^{12}) - E_t(i_{t+12(n-2)}^{12})\right)\right] \\
&= \frac{1}{n}\left[ni_t^{12n} - i_t^{12} - E_t(i_{t+12}^{12}) - ... - E_t(i_{t+12(n-2)}^{12}) - E_t(i_{t+12(n-1)}^{12})\right] \\
&= i_t^{12n} - \frac{1}{n}\sum_{h=0}^{n-1} E_t(i_{t+12h}^{12}) = LHS
\end{aligned}$$

The *expected yields* then are defined as the difference between the *term premiums* and the actual yields at the same maturity. Then, *relative expected yields* and *relative term premiums* are calculated by the cross-county differences.

B.1.4 Decomposed Results

In order to describe the shape of *relative expected yields* and *relative term premiums* at all maturities in a parsimonious way, we summarize them with three factors: *level*, *slope* and *curvature*.[8]. **Table B.1** and **B.2** reports summary statistics for *relative expected yield* and *relative term premiums* factors. Tables first show the mean and standard deviation of three factors from four models and then report the correlation in each of three factors from different models. We first confirm the resemblance among NS, NSM and CP model and the distinctiveness of the Affine model. Correlation in the *level, slope and curvature* factor is high for pairs from NS, NSM and CP model, but low for pairs with Affine model. We conjecture that this unique characteristics of the Affine model is due to the concept of risk neutral yields as well as imposition of no-arbitrage condition and correction of small sample bias. Second, the *slope* factor of *relative expected yields* from NS, NSM and CP model is smaller than the *slope* factor of entire *relative yields*. This observation is consistent with the idea that the yield curve is the sum of both *expected yields* and *term premiums*, so separating out the *term premiums* results in the flatter slope.

[8]The choice of three factors is consistent with the idea of three Nelson-Siegel factors. Recall the definition of *relative expected yields* of maturity m, $E_t i_t^{R,m} = E_t i_t^m - E_t i_t^{m,*}$ and *relative term premiums* of maturity m, $\theta_t^{R,m} = \theta_t^m - \theta_t^{m,*}$. The *level*, *slope* and *curvature* factor of the *relative expected yields* are constructed as follows: $L(E_t i_t^R) = E_t i_t^{R,120}$, $S(E_t i_t^R) = E_t i_t^{R,3} - E_t i_t^{R,120}$, $C(E_t i_t^R) = 2E_t i_t^{R,24} - (E_t i_t^{R,3} + E_t i_t^{R,120})$. The *level*, *slope* and *curvature* factor of the *relative term premiums* are defined as follows: $L(\theta_t^R) = \theta_t^{R,120}$, $S(\theta_t^R) = \theta_t^{R,3} - \theta_t^{R,120}$, $C(\theta_t^R) = 2\theta_t^{R,24} - (\theta_t^{R,3} + \theta_t^{R,120})$

Table B.1: Summary Statistics for Relative Expected Yield Factors

		AU	CA	DK	JP	NZ	SE	CH	UK
1) NS model									
$L(E_t i_t^R)$	$Mean$	1.939	-0.847	-0.253	-2.540	2.438	-0.231	-1.523	0.381
	SD	1.014	2.086	0.888	1.301	0.875	1.137	0.951	1.309
$S(E_t i_t^R)$	$Mean$	0.229	1.069	0.208	0.121	0.376	0.342	0.135	0.567
	SD	0.904	2.000	0.685	1.048	1.089	1.104	0.869	1.149
$C(E_t i_t^R)$	$Mean$	0.138	0.350	0.066	0.259	0.215	-0.022	0.183	0.261
	SD	0.706	0.897	0.637	0.724	0.600	0.844	0.621	0.583
2) NSM model									
$L(E_t i_t^R)$	$Mean$	1.553	-0.738	0.046	-2.552	2.573	0.696	-1.684	0.921
	SD	0.937	0.976	0.961	1.164	1.381	1.020	0.811	1.007
$S(E_t i_t^R)$	$Mean$	0.606	0.959	-0.109	0.194	0.291	-0.493	0.289	0.072
	SD	1.146	1.019	1.328	1.248	1.432	1.461	1.019	1.146
$C(E_t i_t^R)$	$Mean$	0.463	0.788	0.367	0.582	0.585	0.900	0.394	0.512
	SD	0.869	1.326	1.256	0.803	0.835	2.047	0.752	0.801
3) Affine model									
$L(E_t i_t^R)$	$Mean$	5.521	2.958	2.091	1.132	4.479	3.236	1.601	2.991
	SD	2.273	1.654	1.862	2.706	1.784	2.108	2.125	1.431
$S(E_t i_t^R)$	$Mean$	-3.304	-2.721	-2.148	-3.530	-1.608	-3.035	-2.962	-1.981
	SD	1.016	1.013	0.834	0.678	1.341	1.316	0.689	0.770
$C(E_t i_t^R)$	$Mean$	-3.206	-2.662	-2.102	-3.398	-1.606	-2.979	-2.874	-1.917
	SD	0.636	0.614	0.597	0.244	0.633	0.893	0.409	0.263
4) CP model									
$L(E_t i_t^R)$	$Mean$	1.558	-0.323	-0.856	-2.112	2.641	-1.176	-1.412	0.398
	SD	0.982	0.575	0.904	1.327	0.890	0.866	0.870	0.675
$S(E_t i_t^R)$	$Mean$	0.440	0.513	0.712	-0.249	-0.032	1.057	0.071	0.402
	SD	0.734	0.525	0.775	0.916	0.733	0.756	0.697	0.607
$C(E_t i_t^R)$	$Mean$	0.262	0.277	0.440	0.021	-0.014	0.546	0.110	0.224
	SD	0.398	0.315	0.309	0.410	0.380	0.439	0.316	0.403

Note: 1. The *relative expected yields*, $E_t i_t^{R,m}$ for $m = 3, 6, 12, ..., 120$, are computed by four different models: 1) the Nelson-Siegel model (NS model), the Nelson-Siegel model augmented with macro variables (NSM model), the Affine model and Cochrane-Piazzesi approach (CP model). See text and Appendix for details about how we obtained "*relative expected yields*" and "*relative term premiums*" from the cross-country yield curves. 2. The *level*, *slope* and *curvature* factors of the *relative expected yields* are defined as follows: $L(E_t i_t^R) = E_t i_t^{R,120}$, $S(E_t i_t^R) = E_t i_t^{R,3} - E_t i_t^{R,120}$, $C(E_t i_t^R) = 2E_t i_t^{R,24} - (E_t i_t^{R,3} + E_t i_t^{R,120})$. 3. Sample period is from January 1995 to December 2015.

(Continued)

	AU	CA	DK	JP	NZ	SE	CH	UK
$L(E_t i_t^R)$								
corr(NS, NSM)	0.688	-0.123	0.432	0.971	0.098	0.259	0.931	0.823
corr(NS, Affine)	0.934	0.439	0.951	0.996	0.939	0.729	0.960	0.821
corr(NS, CP)	0.960	0.080	0.710	0.922	0.768	0.838	0.931	0.552
corr(NSM, Affine)	0.671	0.720	0.220	0.958	-0.012	0.822	0.961	0.627
corr(NSM, CP)	0.743	0.637	0.827	0.952	0.082	0.444	0.883	0.727
corr(Affine, CP)	0.915	0.720	0.541	0.903	0.866	0.724	0.878	0.701
$S(E_t i_t^R)$								
corr(NS, NSM)	0.734	-0.132	0.682	0.974	0.264	0.596	0.945	0.799
corr(NS, Affine)	-0.013	0.282	0.104	-0.202	0.669	0.234	-0.223	0.713
corr(NS, CP)	0.842	-0.372	0.686	0.833	0.688	0.841	0.847	0.184
corr(NSM, Affine)	-0.392	0.368	-0.496	-0.296	-0.366	0.510	-0.185	0.411
corr(NSM, CP)	0.638	0.615	0.921	0.857	0.084	0.557	0.880	0.712
corr(Affine, CP)	-0.164	-0.517	-0.584	-0.602	0.354	-0.118	-0.568	-0.280
$C(E_t i_t^R)$								
corr(NS, NSM)	0.868	0.474	0.721	0.960	0.779	0.674	0.845	0.119
corr(NS, Affine)	-0.087	0.732	-0.036	-0.440	0.098	0.211	-0.426	0.438
corr(NS, CP)	0.661	0.282	0.756	0.733	0.724	0.764	0.746	0.073
corr(NSM, Affine)	-0.195	0.698	0.245	-0.299	-0.374	0.354	-0.168	-0.451
corr(NSM, CP)	0.531	0.574	0.440	0.665	0.449	0.761	0.710	0.703
corr(Affine, CP)	0.548	0.129	-0.023	-0.043	0.371	0.304	-0.122	-0.349

Note: We report the correlation in *relative expected yield* factors estimated by four different models: 1) NS model, 2) NSM model, 3) Affine model and 4) CP model.

Table B.2: Summary Statistics for Relative Term Premium Factors

		AU	CA	DK	JP	NZ	SE	CH	UK
1) NS model									
$L(\theta_t^R)$	$Mean$	-0.696	0.908	0.016	-0.368	-0.922	0.135	-0.418	-0.187
	SD	0.734	2.504	1.021	0.920	0.661	0.821	0.617	1.083
$S(\theta_t^R)$	$Mean$	0.678	-0.903	-0.014	0.329	0.894	-0.104	0.381	0.194
	SD	0.688	2.414	0.978	0.897	0.608	0.786	0.591	1.019
$C(\theta_t^R)$	$Mean$	0.333	-0.033	-0.034	-0.273	0.422	0.367	-0.105	0.198
	SD	0.260	1.000	0.368	0.531	0.313	0.378	0.301	0.493
2) NSM model									
$L(\theta_t^R)$	$Mean$	-0.308	0.800	-0.299	-0.337	-1.054	-0.793	-0.255	-0.728
	SD	0.780	0.958	0.535	0.708	1.496	1.665	0.518	0.617
$S(\theta_t^R)$	$Mean$	0.297	-0.799	0.278	0.236	0.973	0.734	0.223	0.684
	SD	0.755	0.943	0.474	0.663	1.474	1.549	0.470	0.519
$C(\theta_t^R)$	$Mean$	0.013	-0.463	-0.313	-0.609	0.053	-0.546	-0.313	-0.041
	SD	0.620	1.156	0.750	0.546	0.447	1.442	0.506	0.662
3) Affine model									
$L(\theta_t^R)$	$Mean$	-4.257	-2.869	-2.318	-3.984	-2.931	-3.337	-3.491	-2.797
	SD	2.068	1.779	2.006	2.203	1.453	2.213	1.821	1.383
$S(\theta_t^R)$	$Mean$	4.229	2.884	2.309	3.956	2.889	3.336	3.464	2.775
	SD	2.015	1.721	1.957	2.146	1.414	2.176	1.777	1.337
$C(\theta_t^R)$	$Mean$	3.621	2.930	2.112	3.315	2.134	3.238	2.901	2.298
	SD	1.094	0.922	1.112	1.186	1.004	1.274	1.032	0.664
4) CP model									
$L(\theta_t^R)$	$Mean$	-0.361	0.341	0.548	-0.720	-1.069	0.932	-0.488	-0.262
	SD	0.544	0.415	0.496	0.497	0.423	0.603	0.422	0.474
$S(\theta_t^R)$	$Mean$	0.278	-0.289	-0.574	0.499	0.906	-0.767	0.225	0.205
	SD	0.434	0.267	0.361	0.415	0.378	0.395	0.397	0.384
$C(\theta_t^R)$	$Mean$	-0.053	-0.126	-0.499	-0.191	0.329	-0.393	-0.136	0.022
	SD	0.324	0.186	0.263	0.251	0.281	0.305	0.248	0.350

Note: 1. The *relative term premiums*, $\theta_t^{R,m}$ for $m = 3, 6, 12, ..., 120$, are computed by four different models: 1) the Nelson-Siegel model (NS model), the Nelson-Siegel model augmented with macro variables (NSM model), the Affine model and Cochrane-Piazzesi approach (CP model). See text and Appendix for details about how we obtained "*relative expected yields*" and "*relative term premiums*" from the cross-country yield curves. 2. The *level*, *slope* and *curvature* factor of the *relative term premiums* are defined as follows: $L(\theta_t^R) = \theta_t^{R,120}$, $S(\theta_t^R) = \theta_t^{R,3} - \theta_t^{R,120}$, $C(\theta_t^R) = 2\theta_t^{R,24} - (\theta_t^{R,3} + \theta_t^{R,120})$. 3. Sample period is from January 1995 to December 2015.

(Continued)

	AU	CA	DK	JP	NZ	SE	CH	UK
$L(\theta_t^R)$								
corr(NS, NSM)	0.482	0.291	0.333	0.945	0.123	0.627	0.819	0.739
corr(NS, Affine)	0.893	0.620	0.942	0.905	0.721	0.891	0.883	0.791
corr(NS, CP)	0.883	0.683	0.665	0.777	0.356	0.649	0.796	0.312
corr(NSM, Affine)	0.468	0.745	0.247	0.826	-0.230	0.770	0.826	0.570
corr(NSM, CP)	0.357	0.442	0.462	0.711	-0.056	0.728	0.662	0.352
corr(Affine, CP)	0.874	0.860	0.651	0.946	0.741	0.806	0.809	0.643
$S(\theta_t^R)$								
corr(NS, NSM)	0.419	0.247	0.234	0.945	0.077	0.654	0.809	0.720
corr(NS, Affine)	0.896	0.614	0.939	0.912	0.712	0.892	0.879	0.794
corr(NS, CP)	0.819	0.492	0.738	0.664	0.037	0.767	0.685	0.086
corr(NSM, Affine)	0.419	0.734	0.182	0.841	-0.289	0.781	0.856	0.576
corr(NSM, CP)	0.141	0.600	0.244	0.607	-0.451	0.789	0.762	0.205
corr(Affine, CP)	0.854	0.844	0.671	0.861	0.594	0.911	0.753	0.447
$C(\theta_t^R)$								
corr(NS, NSM)	0.829	0.406	0.054	0.915	0.085	-0.296	0.624	-0.284
corr(NS, Affine)	0.438	0.826	0.889	0.867	0.620	0.634	0.643	0.651
corr(NS, CP)	0.219	0.289	0.661	0.358	0.422	0.041	0.443	-0.483
corr(NSM, Affine)	0.123	0.541	0.160	0.734	-0.440	0.056	0.436	0.026
corr(NSM, CP)	0.156	0.623	-0.208	0.336	-0.473	0.260	0.530	0.643
corr(Affine, CP)	0.715	0.429	0.623	0.546	0.665	0.539	0.573	-0.172

Note: We report the correlation in *relative term premium* factors estimated by four different models: 1) NS model, 2) NSM model, 3) Affine model and 4) CP model.

B.2 VAR Multi-Period Predictions

To compute the partial R^2 for each variable and their total contribution in the VAR, we follow the procedure as described in Hodrick (1992). The method is also adopted in Campbell and Shiller (1988b) and Campbell (1991), among others. The VAR models can be written as:

$$f_t = Af_{t-1} + e_t \tag{B.20}$$

where the constant term μ is omitted for notational convenience. Denote the information set at time t as I_t, which includes all current and past values of f_t. A forecast of horizon m can be written as $E_t\left(f_{t+m}|I_t\right) = A^m f_t$. By repeated substitution, first-order VAR can be expressed in its MA(∞) representation:

$$f_t = \sum_{j=0}^{\infty} A^j e_{t-j} \tag{B.21}$$

where the covariance matrix of e_t is Q. Then, the unconditional variance of f_t can then be expressed as:

$$C(0) = \sum_{j=0}^{\infty} A^j Q A^{j'} \tag{B.22}$$

Denoting $C(j)$ as the jth-order covariance of f_t, which is calculated as $C(j) = A^j C(0)$, the variance of the sum of m consecutive $f_t s$, denoted as V_m, is then:

$$V_m = mC(0) + \sum\nolimits_{j=1}^{m-1} (m-j)\left[C(j) + C(j)'\right] \tag{B.23}$$

We are not interested in the variance of the whole vector but only that of the long-horizon exchange rate change, Δs_t, which is the third element in the vector f_t. We can define $e_3' = (0,0,1,0,0,0)$, and express the variance of the m-period exchange rate change as $e_3' V_m e_3$.

To assess whether a variable in f_t, say the *relative level* factor L_t^R, explains exchange rate change $\Delta s_{t+m} = s_{t+m} - s_t$, we run a long-horizon regression of Δs_{t+m} on L_t^R. The VAR model for f_t allows us to calculate the coefficient from this regression based on only the VAR coefficient estimates. Since the *relative level* factor is the fourth element in f_t, the coefficient is defined as:

$$\beta_4(m) = \frac{e_3' \left[C(1) + ... + C(m)\right] e_4}{e_4' C(0) e_4} \tag{B.24}$$

where vector e_4 is defined as $e_4 = (0, 0, 0, 1, 0, 0)$. The numerator is the covariance between Δs_{t+m} and L_t^R, and the denominator is the variance of L_t^R. Finally, the R^2 as reported in the results is calculated as:

$$R_4^2(m) = \beta_4(m)^2 \frac{e_4' C(0) e_4}{e_3' V_m e_3} \tag{B.25}$$

The R^2 for all other variables in the vector f_t can be suitably obtained by replacing e_4 with e_1, e_2, e_3, e_5, e_6.

To calculate the total R^2 for all explanatory variables, we calculate the innovation variance of the exchange rate change as $e_3' W_m e_3$, where

$$W_m = \sum_{j=1}^{m} (I - A)^{-1} (I - A^j) Q (I - A^j)' (I - A)^{-1\prime} \tag{B.26}$$

The total R^2 is then:

$$R^2(m) = 1 - \frac{e_3' W_m e_3}{e_m' V_m e_m} \tag{B.27}$$

For the calculation to be valid, we need A to be stationary.

B.3 Robustness Checks

This Appendix presents the robustness check results of our main results. First, the effect of the ZLB on the significance of risk factors can be shown more clearly with multiple structural breaks. **Table B.3** reports the Wald test results that risk factors have become relevant only after the second break around 2011 - 2012 for European countries such as Denmark, Sweden, and Switzerland because these countries hit the ZLB not around 2008 but around 2011 - 2012. This observation supports our argument that the unusual monetary policy condition might be one good reason why risk channel has become relatively strong.

Second, our main exchange rates are defined against the USD. **Table B.4** shows the Wald test results on which of expectations or risk explains more of the subsequent exchange rate against different denominating currency. Similarly, **Table B.5** shows whether the macro-fundamentals deliver additional information to the yield factors in explaining the exchange rate against different denominating currency.

Third, concerned about the possibility of high adjusted R^2 due to overlapping data in multi-month-ahead predictions, we report the goodness of fit measures from the non-overlapping regressions in **Table B.6**.

Table B.3: Robustness Check with Multiple Breaks: Expectations vs. Risk

$$\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \beta_4 S(\theta_t^R) + \beta_5 C(\theta_t^R) + \epsilon_{t+3}$$

	AU	CA	DK	JP	NZ	SE	CH	UK
Wald test								
No Expectation?								
Full	0.001	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Pre-Crisis	0.000	0.003	0.001	0.008	0.001	0.000	0.149	0.671
Crisis	0.562	0.000	0.000	0.000	0.000	0.001	0.000	0.000
Post-Crisis	0.096		0.000		0.066	0.000	0.000	0.064
No Risk?								
Full	0.042	0.729	0.022	0.000	0.000	0.000	0.000	0.009
Pre-Crisis	0.248	0.982	0.805	0.000	0.256	0.366	0.635	0.025
Crisis	0.008	0.168	0.382	0.000	0.000	0.459	0.400	0.039
Post-Crisis	0.667		0.002		0.225	0.000	0.000	0.184
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Adj. R^2	0.342	0.228	0.276	0.245	0.409	0.353	0.220	0.388
*Breakdate*1	May,08	Jan,07	May,08	Sep,08	May,08	May,08	Jun,07	Jun,08
*Breakdate*2	Dec,12		Aug.11		Jun,11	Dec,12	Jun,11	Jul,11

Note: 1. We run the same regression equation in **Table 3.5** with multiple breaks, instead of a single break. 2. The multiple structural breaks are tested by the Bai and Perron (2003) test (with 15% trimming and 5% significance level.) One or two breaks are chosen to identify the Great Recession period. The first break date is around 2008 when the Global Financial Crisis has been triggered. The second break date is around 2011 - 2012. Each breakdate indicates the first month of subsequent sub-period. 3. After identifying the breakdates, structural break dummy variables for each sub-period are incorporated into the regression. 4. The row labeled "No Expectation" reports the p-values of the Wald tests for the null hypothesis that the *relative expected yield* factors have no explanatory power ($\beta_1 = \beta_2 = \beta_3 = 0$, for each sub-period), and the "No Risk" row tests the null hypothesis that the *relative term premium* factors do not matter ($\beta_4 = \beta_5 = 0$, for each sub-period). 5. P-value is for the Wald test that factors jointly have no explanatory power ($H_0 : \beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = 0,\ \forall\ sub-periods$).

Table B.4: Robustness Check with Different Currency Pairs: Expectations vs. Risk

$$\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \beta_4 S(\theta_t^R) + \beta_5 C(\theta_t^R) + \epsilon_{t+3}$$

Panel A	AU	CA	DK	NZ	SE	CH	UK	US
No Expectation?								
Full	0.000	0.217	0.001	0.000	0.150	0.000	0.000	0.000
Pre-Break	0.000	0.636	0.235	0.106	0.293	0.768	0.423	0.005
Post-Break	0.024	0.095	0.000	0.000	0.124	0.000	0.000	0.000
No Risk?								
Full	0.012	0.002	0.315	0.076	0.012	0.048	0.011	0.000
Pre-Break	0.111	0.292	0.151	0.065	0.574	0.218	0.493	0.000
Post-Break	0.014	0.001	0.610	0.234	0.003	0.038	0.003	0.000
p-value	0.000	0.002	0.001	0.000	0.003	0.000	0.000	0.000
Adj. R^2	0.226	0.180	0.170	0.267	0.204	0.125	0.377	0.264
Breakdate	May,08	Jul,08	Jul,08	Jul,08	Jun,08	Jun,08	Jul,08	Sep,08
Panel B	AU	CA	DK	JP	NZ	SE	UK	US
No Expectation?								
Full	0.000	0.000	0.000	0.000	0.000	0.009	0.002	0.000
Pre-Break	0.000	0.001	0.015	0.746	0.002	0.002	0.003	0.334
Post-Break	0.001	0.001	0.000	0.000	0.013	0.506	0.059	0.000
No Risk?								
Full	0.000	0.007	0.000	0.040	0.001	0.580	0.008	0.000
Pre-Break	0.000	0.167	0.172	0.173	0.064	0.923	0.017	0.330
Post-Break	0.003	0.007	0.000	0.039	0.002	0.259	0.056	0.000
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.001	0.000
Adj. R^2	0.230	0.176	0.267	0.127	0.218	0.183	0.186	0.149
Breakdate	May,08	Oct,98	Mar,11	Jun,08	May,04	Jul,08	May,00	Jun,11

Note: 1. We run the same regression equation in **Table 3.5** with different denominating currencies: in **Panel A** with Japanese Yen, in **Panel B** with Swiss Franc. 2. The *relative expected yield* and *relative term premium* factors are estimated by NSM model. 3. After identifying the unknown breakpoint around 2008 by the Quants-Andrews test (Andrews, 1993; Hansen, 1997) with 15% trimming and 5% significance level, structural break dummy variables for each sub-period are incorporated into the regression. 4. The row labeled "No Expectation" reports the p-values of the Wald tests for the null hypothesis that the *relative expected yield* factors have no explanatory power ($\beta_1 = \beta_2 = \beta_3 = 0$, for each sub-period), and the "No Risk" row tests the null hypothesis that the *relative term premium* factors do not matter ($\beta_4 = \beta_5 = 0$, for each sub-period). 5. P-value is for the Wald test that factors jointly have no explanatory power ($H_0 : \beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = 0, \ \forall \ sub-periods$). 6. Adjusted R^2 is also reported.

Table B.5: Robustness Check with Different Currency Pairs: Yields vs. Macro

$$\Delta s_{t+3} = \beta_0 + \beta_1 L_t^R + \beta_2 S_t^R + \beta_3 C_t^R + \beta_4 u_t^R + \beta_5 \pi_t^R + \epsilon_{t+3}$$

Panel A	AU	CA	DK	NZ	SE	CH	UK	US
No Yields?								
Full	0.005	0.228	0.000	0.000	0.119	0.000	0.000	0.000
Pre-Break	0.013	0.988	0.970	0.785	0.343	0.246	0.243	0.044
Post-Break	0.047	0.047	0.000	0.000	0.079	0.000	0.000	0.000
No Macro?								
Full	0.076	0.048	0.086	0.014	0.277	0.917	0.009	0.310
Pre-Break	0.203	0.353	0.335	0.003	0.237	0.718	0.003	0.679
Post-Break	0.072	0.023	0.049	0.596	0.328	0.864	0.353	0.130
p-value	0.001	0.102	0.000	0.000	0.228	0.000	0.000	0.000
Adj. R^2	0.183	0.149	0.145	0.269	0.175	0.096	0.368	0.150
Breakdate	May,08	Jul,08	Jun,08	Jul,08	Jun,08	Jul,12	Jul,08	Jul,08
Panel B	AU	CA	DK	JP	NZ	SE	UK	US
No Yields?								
Full	0.000	0.000	0.026	0.000	0.000	0.000	0.170	0.000
Pre-Break	0.000	0.146	0.010	0.235	0.206	0.001	0.576	0.000
Post-Break	0.797	0.000	0.411	0.000	0.000	0.000	0.069	0.020
No Macro?								
Full	0.009	0.000	0.060	0.912	0.006	0.003	0.749	0.000
Pre-Break	0.009	0.552	0.386	0.705	0.294	0.925	0.533	0.001
Post-Break	0.129	0.000	0.028	0.867	0.002	0.000	0.715	0.000
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.045	0.000
Adj. R^2	0.078	0.127	0.154	0.098	0.161	0.185	0.126	0.182
Breakdate	Apr,02	Dec,08	Oct,07	Jul,12	Oct,07	Jun,08	Aug,07	Sep,03

Note: 1. We run the same regression equation in **Table 3.6** with different denominating currencies: in **Panel A** with Japanese Yen, in **Panel B** with Swiss Franc. 2. After identifying the unknown breakpoint around 2008 by the Quants-Andrews test (Andrews, 1993; Hansen, 1997) with 15% trimming and 5% significance level, structural break dummy variables for each sub-period are incorporated into the regression. 3. The row labeled "No Yields" reports the p-values of the Wald tests for the null hypothesis that the *relative yield* factors have no explanatory power ($\beta_1 = \beta_2 = \beta_3 = 0$, for each sub-period), and the "No Macro" row tests the null hypothesis that macroeconomic fundamentals do not matter ($\beta_4 = \beta_5 = 0$, for each sub-period). 4. P-value is for the Wald test that factors jointly have no explanatory power ($H_0 : \beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = 0, \; \forall \; sub-periods$). 5. Adjusted R^2 is also reported.

Table B.6: Robustness Check with Non-Overlapping Regressions

	AU	CA	DK	NZ	SE	CH	UK	US
Panel A: $\Delta s_{t+3} = \beta_0 + \beta_1 L_t^R + \beta_2 S_t^R + \beta_3 C_t^R + \epsilon_{t+3}$								
p-value	0.023	0.006	0.104	0.008	0.000	0.000	0.072	0.000
Adj. R^2	0.128	0.167	0.075	0.158	0.293	0.281	0.087	0.373
Breakdate	Q2,08	Q1,07	Q2,08	Q3,08	Q2,08	Q2,08	Q3,11	Q2,08
Panel B: $\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \epsilon_{t+3}$								
p-value	0.003	0.001	0.002	0.034	0.000	0.000	0.026	0.000
Adj. R^2	0.183	0.218	0.203	0.114	0.329	0.319	0.122	0.395
Breakdate	Q4,07	Q1,07	Q3,08	Q4,02	Q3,08	Q3,08	Q4,00	Q3,08
Panel C: $\Delta s_{t+3} = \beta_0 + \beta_1 S(\theta_t^R) + \beta_2 C(\theta_t^R) + \epsilon_{t+3}$								
p-value	0.005	0.008	0.042	0.007	0.001	0.124	0.047	0.232
Adj. R^2	0.155	0.144	0.095	0.145	0.190	0.060	0.090	0.037
Breakdate	Q3,08	Q4,07	Q3,08	Q4,02	Q2,08	Q3,08	Q3,11	Q4,08
Panel D: $\Delta s_{t+3} = \beta_0 + \beta_1 L(E_t i_t^R) + \beta_2 S(E_t i_t^R) + \beta_3 C(E_t i_t^R) + \beta_4 S(\theta_t^R) + \beta_5 C(\theta_t^R) + \epsilon_{t+3}$								
p-value	0.010	0.007	0.015	0.010	0.000	0.000	0.054	0.000
Adj. R^2	0.181	0.190	0.167	0.179	0.318	0.353	0.118	0.411
Breakdate	Q3,08	Q1,07	Q3,08	Q2,08	Q3,08	Q3,08	Q3,98	Q3,08
Panel E: $\Delta s_{t+3} = \beta_0 + \beta_1 L_t^R + \beta_2 S_t^R + \beta_3 C_t^R + \beta_4 u_t^R + \beta_5 \pi_t^R + \epsilon_{t+3}$								
p-value	0.010	0.014	0.004	0.016	0.000	0.000	0.101	0.000
Adj. R^2	0.181	0.171	0.212	0.163	0.428	0.326	0.092	0.448
Breakdate	Q3,08	Q1,07	Q3,08	Q3,98	Q3,08	Q3,08	Q4,00	Q3,08

Note: 1. Considering that high adjusted R^2's from our regressions may be due to overlapping data with high persistence, we check the robustness with non-overlapping data; we regress the quarterly exchange rate changes with quarterly data. 2. **Panel A** is non-overlapping regression for **Table 3.2**; **Panel B** for **Table 3.3-Panel A** from NSM model; **Panel C** for **Table 3.3-Panel B** from NSM model; **Panel D** for **Table 3.3-Panel C** from NSM model; **Panel E** for **Table 3.6**. 3. After identifying the unknown breakpoint around 2008 by the Quants-Andrews test (Andrews, 1993; Hansen, 1997) with 15% trimming and 5% significance level, structural break dummy variables for each sub-period are incorporated into the regression. 4. P-value is for the Wald test that factors jointly have no explanatory power ($H_0 : \beta_i = 0,\ \forall\ i,\ \forall\ sub-periods$).