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Essays on Empirical Macroeconomics and International Financial Markets

Tim Leelahaphan

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Acknowledgments

Declaration

I declare the following:

The material contained in this thesis is my own work, with the exception that the contents of Chapter 3 are joint work with Mark P. Taylor.

The thesis has not been submitted for a degree at any another university.

Abstract

The thesis consists of three chapters of self-contained empirical studies.

In Chapter 1, we examine long-run and short-run dynamics of US real trade balance with Canada. In addition to the linear error-correction model, the Markov-switching error-correction model is employed, using quarterly data from 1985 to 2008. We find that real exchange rate, real oil price and real new housing price index have statistically significant effects on US real trade balance with Canada in the long run. We acquire evidence of short-run J-curve. Results show that short-run dynamic effects of real oil price are not so fearful, with statistically insignificant effect on real trade balance following an increase in real oil price. House prices could be argued as being strongly relevant for settlement and adjustment of US trade balance in the long run through the wealth effects. However, the immediate (next-quarter) effect of a change in housing wealth is insignificant, consistent with existing literature. US real trade balance with Canada forecasts from our non-linear VAR model outperform ones from the linear VAR in first difference (DVAR) model and ones from the random walk model. The long-term out-of-sample forecastability is not much improved by the oil price and house price variables, which, nonetheless, actively explain in-sample movement of US real trade balance with Canada in the long run.

In Chapter 2, we examine the effect of monetary policy and exchange rate on stock price movements in Asia. We employ a Bayesian structural vector autoregression model and impose sign restrictions to identify simultaneously and uniquely contractionary monetary policy shocks and exchange rate depreciation shocks in an integrated framework. This study covers the stock markets of Thailand, Malaysia and South Korea, over the period 1989-2008. Our main results acquired using sign restrictions show that monetary policy shocks result in a strongly persistent effect on market index real stock prices whereas the impact of exchange rate shocks is short-lived over the short run. The variance decomposition suggests that the exchange rate is as important as monetary policy for explaining the dynamics of market and financial sector index real stock prices. More precisely, for all the countries examined, real exchange rate developments have been more important in the short run.

In Chapter 3, within the context of a time-varying transition probabilities Markov-switching model of the uncovered interest parity (UIP) condition, we examine if variables measuring fear and volatility have an effect on the probability of switching between the regime where the UIP condition holds and the regime where it does not. The state transition probability depends nonlinearly upon the variables examined. These are the exchange rate volatility, the VIX equity option implied volatility index and the TED spread. Applying this to both US dollar exchange rates and cross (exchange) rates from January 4, 1990 to September 11, 2008, we find that those three variables increase the probability of remaining in the regime where the UIP condition holds. In addition, the probability of switching from the regime where the UIP condition does not hold to the regime where the exchange rate follows the UIP condition decreases as these variables measuring fear and volatility fall, especially the VIX equity option implied volatility index. The smoothed probabilities show that exchange rates examined essentially do not follow the UIP condition except during periods in which the fear and risk variables are increasing, as in the recent global financial crisis in particular.

Introduction

The thesis *Essays on Empirical Macroeconomics and International Financial Markets* consists of three self-contained empirical studies. These are (i) J-Curve, Oil Price, House Price and US-Canada Imbalance, (ii) Monetary Policy, Exchange Rates and Asian Stock Markets and (iii) Fear, Volatility and Uncovered Interest Parity. Here, an introduction to each chapter is addressed.

1 J-Curve, Oil Price, House Price and US-Canada Imbalance

A massive US balance of trade deficit always highlights global imbalances. Figures released by the US Commerce Department in January 2009 show that, as the economic slowdown leads to lower demand for imports, the US trade deficit dropped to its lowest level in more than five years in November 2008. In particular, the US trade deficit shrank by 28.7 percent from October 2008 to 40.4 billion US dollars. Currently, economists have not, however, reached consensus on the direction of US trade balance. Some believe that the US trade deficit

is set to fall while some claim that the US economy could continue to have enormous trade deficits for some time to come. Nevertheless, despite recent improvements, the US trade deficit remains a high priority. It is one of the most striking characteristics of the current global economy. For most economists, this is also one of the most worrying features.

Over the past several years, three developments have been of international macroeconomists' interest, besides the fall of the US dollar. Firstly, large global external imbalances have persisted. These, obviously, include a massive balance of trade deficit in the US. Over the past seven years, we have seen a significant worsening of the aggregate US trade deficit. Evidently, the deficit rose above 6 percent of GDP for the first time in December 2005. This became 5.1 percent in December 2007 and showed a significant increase since March 2001, when it accounted for 3.9 percent of GDP. In addition, this is the largest US trade deficit compared to its levels before March 2004. Secondly, attributed both to strengthening global demand and, most recently, to concerns relating to the issue of supply in years ahead, energy prices have upsurged since 2003. Owing to the limited excess capacity, it is expected that in the medium run we will still realize the remaining very tight balance of supply and demand. Consequently, oil prices would be expected to stay persistently at high levels. Finally, for many of the OECD countries, real housing prices have shown a rapid increase since the mid-1990s. See Figure 1 in Chapter 1 that shows these three important issues.

In Chapter 1, motivated by arguments and findings set out in the chapter,

where we justify how the exchange rate, oil price and asset price relate to some global imbalances and provide reasons for studying the US-Canada imbalance, we would like to examine empirically, both within a linear and non-linear framework, the following issues. These are (i) if evidence of short-run J-curve effects prevail for the US and Canada; (ii) if oil price could be claimed as a major source of global imbalances, the US-Canada imbalance in particular and (iii) if asset price, house price in particular, explains some of the movements of the US trade balance, as recent evidence has shown.[1]

Using quarterly data from 1985 to 2008, we find that real exchange rate, real oil price and real new housing price index have statistically significant effects on US real trade balance with Canada in the long run. We acquire evidence of short-run J-curve within a linear and non-linear framework. Results from both linear and non-linear models show that short-run dynamic effects of real oil price are not so fearful, with statistically insignificant effect on real trade balance following an increase in real oil price. House prices could be argued as being strongly relevant for settlement and adjustment of US trade balance in the long run through the wealth effects. However, the immediate (next-quarter) effect of a change in housing wealth is insignificant, consistent with existing literature. We argue that once J-curve effects disappear, country's trade balance would improve. Nevertheless, policies relating to oil price and house price should be addressed. US real trade balance with Canada forecasts from our

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[1]See a definition of the J-curve phenomenon in Chapter 1.

non-linear VAR model outperform ones from the linear VAR in first difference (DVAR) model and ones from the random walk model. The long-term out-of-sample forecastability is not much improved by the oil price and house price variables, which, nonetheless, actively explain in-sample movement of US real trade balance with Canada in the long run.

2 Monetary Policy, Exchange Rates and Asian Stock Markets

While it is widely believed that the interest rate is an important determinant of stock prices, less attention has been paid to the relationship between the exchange rate and the stock market. Nevertheless, since the 1997 Asian financial crisis, the exchange rate and stock price relationship has received greater attention. This might be due to the fact that, in the aftermath of the crisis, the affected countries suffered turmoil in both foreign exchange and equity markets. In Thailand, the Thai baht reached its lowest point in January 1998 and the stock market fell by 75 percent. In Malaysia, over the period from July 1 to September 30, 1997, the ringgit plunged by 37.4 percent and the stock market fell by 31.4 percent. In South Korea, by the end of the same year, the Korean won dropped dramatically by more than 150 percent and its stock market plunged by more than 50 percent.[2] If the exchange rate explains a large amount of the

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[2]See Baharumshah *et al.* (2002).

dynamics of market and financial sector index real stock prices, it is reasonable to believe that stock market crises and asset price bubbles could potentially be prevented by focusing on the movement of exchange rates. In addition, market participants could use information on exchange rates to predict the stock market behavior.

The link between monetary policy and the stock market is generally realized as the appropriate influence of such policy on the decision-making of the private sector. This is in order to fulfill some objectives. To have low and stable inflation and output near its natural rate are generally believed to be the main objectives of the central bank, achieved by setting and exerting control over the (real) interest rates and by appropriately monitoring the decisions of the private sector. In the framework of the new Keynesian theory, in which prices are not fully flexible in the short run, the real interest rate could be temporarily influenced by the policy of the monetary authority. Consequently, this would affect the real output in addition to nominal prices.

Studying Asian stock markets and economies affected by underlying structural shocks is motivated by the lesson we learn from the Asian crisis of 1997-1998. The severe consequences of the crisis economically and politically destroyed many of the regional economies. It is argued that the primary reason for the Asian financial crisis is attributable to an inappropriate mixture of policies. (See Rajan, Thangavelu and Parinduri (2008)). In particular, this is due to the fact that *regional economies attempted to maintain simultaneously fairly*

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rigid exchange rates (soft US dollar pegs) and monetary policy autonomy in the presence of large-scale capital outflows. Even after more than a decade has passed since the Asian financial crisis, exchange rate issues and monetary policies relevant to Asia, especially those relating to financial issues and asset prices, are still in the focus of economists and market participants. Specifically, the issue of whether monetary policy should actively seek to promote asset price stability might be the most important question central bankers are currently facing.

In Chapter 2, in the context of Asian stock markets and economies, not only is the monetary policy relating to financial issues and asset prices studied, the relationship between exchange rates and stock prices that has received great attention since the 1997 Asian financial crisis also motivates us to have the following two main objectives.[3] Firstly, we would like to examine if there is difference in the influence of the monetary policy actions and of exchange rate developments on the stock market. In particular, the systematic feature, in terms of the persistence of the impact, of such an influence of these two underlying structural shocks is deliberately considered. Secondly, in addition to monetary policy commonly believed to be important determinant of stock prices, we would like to assess quantitatively if the exchange rate has also played an important role in driving the stock market. In particular, we examine and compare the extent to which monetary policy and exchange rate are responsible for the movements in Asian stock prices. In order to reach findings, we adapt standard VAR analyses

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[3]See the relationship between the exchange rate and the stock market in Chapter 2.

to deal with single shocks based upon sign restrictions.

This study covers the stock markets of Thailand, Malaysia and South Korea, over the period 1989-2008. Our main results acquired using sign restrictions show that monetary policy shocks result in a strongly persistent effect on market index real stock prices whereas the impact of exchange rate shocks is short-lived over the short run. The variance decomposition suggests that the exchange rate is as important as monetary policy for explaining the dynamics of market and financial sector index real stock prices. More precisely, for all the countries examined, real exchange rate developments have been more important in the short run. Based purely on our findings, two conclusions are made. Firstly, because of the mistimed and/or persistent effect of monetary policy on both the real economy and financial markets, we argue that one needs to be cautious in using monetary policy to constrain asset price misalignment. Secondly, due to the evidence that exchange rates principally have a contemporaneous impact on equity prices, we suggest that, in the short run, such incorrectly aligned asset prices might potentially be corrected by focusing on exchange rate movements.

3 Fear, Volatility and Uncovered Interest Parity

In international macroeconomics and international finance models, the uncovered interest parity (UIP) condition is commonly assumed. This parity condition

claims that, over some particular time horizon, the foreign exchange's expected gain from holding funds in one currency rather than the other should be compensated by the opportunity cost of holding that currency instead of another. In particular, an interest rate differential between the two currencies must offset expected changes in the exchange rate under the UIP condition. In addition, one might hypothesize that market participants would expect currencies with higher interest rates to fall in value.

However, Fama (1984), a highly influential paper, notes instead that such currencies with high interest rates tend to appreciate. This is inconsistent with the commonly assumed UIP condition. With this result, we could infer that there would be an inverse relationship between the forward premium and the future exchange rate changes.[4] Evidently, for many currencies and periods examined, this remark in line with Fama (1984) has been asserted in the literature.

The deviation from the UIP condition leads to the result that profit could be generated through 'currency carry trade'. This is a strategy in which a market participant takes a short position on a currency with low interest rates, which is called the funding currency, and takes a long position on a currency with high interest rates, which is called the target currency. This is due to the fact that such failure of the UIP conditions indicates that the target currency (with a high interest rate) does not depreciate against the funding currency (with a low interest rate) by a percentage that matches the interest rate differential between

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[4]This is explained in detail in Section 2 in Chapter 3.

the two currencies. Therefore, if that failure of the UIP condition actually prevails, a positive return would then be made.

Recently, one consensus relating to foreign exchange carry trades has been likely to be reached by many market participants and monetary authorities. That is, the failure of the UIP condition (or the appreciation of currencies with high interest rates) has been associated with the currency carry trade activities with a trend lower in volatility. Supporting this view, de Rato (2007), the managing director of the International Monetary Fund (IMF), mentions in his speech that the currency carry trade reflects environments with both low volatilities and large interest rate differentials. Such conditions support market participants acquiring high excess returns per unit of risk (or the Sharpe ratio) measures of such strategy. In addition, this has placed downward pressure on the currencies with low interest rates. The speech addresses rapid reversal movements of exchange rates caused by the unwinding of the foreign exchange carry trade positions.[5]

In Chapter 3, we are motivated by the arguments presented in the chapter, where we describe the crisis in foreign exchange markets affected by the recent global financial crisis and demonstrate its consequence in such markets that is associated with the UIP condition. Within the context of a time-varying transition probabilities Markov-switching model of the UIP condition, we examine if variables measuring fear and volatility, which increase in periods of crisis, have

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[5]See Koyama and Ichiue (2008).

an effect on the probability of switching between the regime where the UIP condition holds and the regime where it does not. The state transition probability depends nonlinearly upon the variables examined. These are the exchange rate volatility, the VIX equity option implied volatility index and the TED spread.

Applying this to both US dollar exchange rates and cross (exchange) rates from January 4, 1990 to September 11, 2008, we find that those three variables increase the probability of remaining in the regime where the UIP condition holds. In addition, the probability of switching from the regime where the UIP condition does not hold to the regime where the exchange rate follows the UIP condition decreases (increases) as these variables measuring fear and volatility fall (rise), especially the VIX equity option implied volatility index. The smoothed probabilities show that exchange rates examined essentially do not follow the UIP condition except during periods in which the fear and risk variables are increasing, as in the recent global financial crisis in particular.

The remainder of the thesis consists of three chapters of self-contained empirical studies. Chapter 1: J-Curve, Oil Price, House Price and US-Canada Imbalance, Chapter 2: Monetary Policy, Exchange Rates and Asian Stock Markets and Chapter 3: Fear, Volatility and Uncovered Interest Parity. Conclusions and implications of each chapter and conclusions of the thesis are addressed in the Conclusion.

References

Baharumshah, A. Z., Masih, A. M. M., and Azali, M. (2002). The stock market and the ringgit exchange rate: A note. *Japan and the World Economy*, 14(4):471–486.

Bjornland, H. C. and Leitemo, K. (2009). Identifying the interdependence between US monetary policy and the stock market. *Journal of Monetary Economics*, 56(2):275–282.

de Rato, R. (2007). Speech by Rodrigo de Rato. .

Dunaway, S. (2009). *Global imbalances and the financial crisis (council special report).* Council on Foreign Relations Press, New York.

Fama, E. F. (1984). Forward and spot exchange rates. *Journal of Monetary Economics*, 14:319–338.

Fratzscher, M., Juvenal, L., and Sarno, L. (2008). Asset Prices, Exchange Rates and the Current Account. *Research Division, Federal Reserve Bank of St. Louis Working Paper Series*, 31.

Gonzalez-Paramo, J. M. (2006). Speech by Jose Manuel Gonzalez-Paramo: Global imbalances and the need for policy adjustment. .

Koyama, K. and Ichiue, H. (2008). Regime switches in exchange rate volatility and uncovered interest parity. *Bank of Japan Working Paper Series*, (07-E-22).

Melvin, M. and Taylor, M. P. (2009). The crisis in the foreign exchange market. *CESifo Working Paper Series*, (2707).

Rajan, R. S., Thangavelu, S., and Parinduri, R. A. (2008). *Exchange rate, monetary and financial issues and policies in Asia.* World Scientific Publishing Company.

Chapter 1: J-Curve, Oil Price, House Price and US-Canada Imbalance

Tim Leelahaphan[*]

Abstract

We examine long-run and short-run dynamics of US real trade balance with Canada. In addition to the linear error-correction model, the Markov-switching error-correction model is employed, using quarterly data from 1985 to 2008. We find that real exchange rate, real oil price and real new housing price index have statistically significant effects on US real trade balance with Canada in the long run. We acquire evidence of short-run J-curve within a linear and non-linear framework. Results from both linear and non-linear models show that short-run dynamic effects of real oil price are not so fearful, with statistically insignificant effect on real trade balance following an increase in real oil price. House prices could be argued as being strongly relevant for settlement and adjustment of US trade balance in the long run through the wealth effects. However, the immediate (next-quarter) effect of a change in housing wealth is insignificant, consistent with existing literature. We argue that once J-curve effects disappear, country's trade balance would improve. Nevertheless, policies relating to oil price and house price should be addressed. US real trade balance with Canada forecasts from our non-linear VAR model outperform ones from the linear VAR in first difference (DVAR) model and ones from the random walk model. The long-term out-of-sample forecastability is not much improved by the oil price and house price variables, which, nonetheless, actively explain in-sample movement of US real trade balance with Canada in the long run.

JEL Classification: F40
Keywords: forecasting, global imbalance, J-curve, Markov-switching, house price, oil price, trade balance, wealth effects

1 Introduction

A massive US balance of trade deficit always highlights global imbalances. Figures released by the US Commerce Department in January 2009 show that, as the economic slowdown leads to lower demand for imports, the US trade deficit dropped to its lowest level in more than five years in November 2008. In particular, the US trade deficit shrank by 28.7 percent from October 2008 to 40.4 billion US dollars. Currently, economists have not, however, reached consensus on the direction of US trade balance. Some believe that the US trade deficit is set to fall while some claim that the US economy could continue to have enormous trade deficits for some time to come. Nevertheless, despite recent improvements, the US trade deficit remains a high priority. It is one of the most striking characteristics of the current global economy. For most economists, this is also one of the most worrying features.

Over the past several years, three developments have been of international macroeconomists' interest, besides the fall of the US dollar. Firstly, large global external imbalances have persisted. These, obviously, include a massive balance of trade deficit in the US. Over the past seven years, we have seen a significant worsening of the aggregate US trade deficit. Evidently, the deficit rose above 6 percent of GDP for the first time in December 2005. This became 5.1 percent in December 2007 and showed a significant increase since March 2001, when it accounted for 3.9 percent of GDP. In addition, this is the largest US trade

deficit compared to its levels before March 2004. Secondly, attributed both to strengthening global demand and, most recently, to concerns relating to the issue of supply in years ahead, energy prices have upsurged since 2003. Owing to the limited excess capacity, it is expected that in the medium run we will still realize the remaining very tight balance of supply and demand. Consequently, oil prices would be expected to stay persistently at high levels. Finally, for many of the OECD countries, real housing prices have shown a rapid increase since the mid-1990s. Figure 1 shows these three important issues.[1] These three developments play an important role in motivation, set out as follows.

1.1 Motivation

We are motivated by the following arguments and findings regarding (i) trade between the US and Canada; (ii) J-curve effects and global imbalance; (iii) oil price and global imbalance; (iv) asset price and global imbalance and (v) business cycle and non-linear model. These are explained as follows.

Firstly, we are motivated by the fact that the US and Canada conduct the world's largest bilateral trade relationship, with the North American Free Trade Agreement implemented in 1994.[2] In addition, Canada is the US' first-ranked trading partner, not China (the second-ranked) often misunderstood. For the month of November 2008, this accounts for 41.77 billions of US dollar. For year

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[1]Sources are Bureau of Economic Analysis, IMF International Financial Statistics and Bureau of the Census, the US Department of Commerce.

[2]This is with total merchandise trade (imports and exports added together) exceeding 533.7 billion US dollars in 2006. See Fergusson (2008).

to date, this is 559.86 billions of US dollar. With China, this is 33.5 billion US dollars in November 2008. For year to date, this is 378.95 billion US dollars. The values given are for imports and exports added together. Table 1 shows this.[3] In 2008, Canada represented 15.88 percent of US imports and 20.24 percent of US exports in goods. In addition, Canada's main trading partner is the US. In particular, approximately three-quarters of trade and the majority of capital moving in and out of Canada are associated with its trade with the US. From these facts provided, which motivate us to perform empirical analysis emphasizing the international trade between these two closely linked countries, we believe that international trade between the US and Canada is a very good case study of bilateral trade due to the fact that they are the each other's biggest trade partner.[4] We are also motivated by the fact that Canada is the top source of US crude oil imports, that is, 2.055 million barrels per day for October 2008, and by the continuing US trade imbalance with Canada, despite the weak US dollar.[5] See Table 2 and Figure 2.[6]

Secondly, J-curve effects, that is, worsening trade balance measured in local

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[3]Source is Bureau of the Census, the US Department of Commerce.

[4]Note that empirical research on balance of trade could be usefully classified into two groups. The first one, on the one hand, is the domestic and rest-of-the-world framework. A second group of research, on the other hand, focuses on trade between two partners only. While our analysis falls into the second group of research, we argue that it also represents the first group well, due to the fact that they are each other's biggest trading partner.

[5]Among the G7 countries, Canada is the only economy that has a surplus current account and government balance. This might be due to the fact that soaring global crude oil prices have benefited Canada's overall trade balance.

[6]Sources are Energy Information Administration, Official Energy Statistics from the US Government, US Census Bureau, Foreign Trade Division and IMF International Financial Statistics.

currency in response to exchange rate depreciation in the impact period, are widely believed to be able to produce a temporary increase in the nominal balance of trade deficit. This could help explain the fact that US current account imbalance has stayed rigidly high despite the decline in the US dollar.[7,8]

Thirdly, we are motivated by the argument that a rise in energy prices has exacerbated some of global imbalances. Clearly, the trade balances of fuel importers would be worsened by the rise in world oil prices.[9] For the US, we have seen that the increase in oil prices since 2003 has directly worsened its current account deficit, in particular, by over 1 percent of GDP.[10]

Fourthly, recent studies have started showing the link between asset prices and the trade balance. This is, essentially, through wealth effects. In particular, an increase in asset prices, especially when it is expected to be permanent, raises expected household income and, therefore, consumption. In addition, this situation helps firms to finance opportunities for investment easily. Consequently, it

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[7]In other words, the Marshall-Lerner condition, that is, improving trade balance following a devaluation, is not met.

[8]Note that the current account is comprised of (i) merchandise trade balance, (ii) service trade balance and (iii) income balance. Following Kilian *et al.* (2007), the trade balance should be recognized as referring to the merchandise trade balance. In fact, the trade in services is not included in the trade balance. This is due to data availability. The quality of service trade data is another concern. In addition, the income balance, which is generally calculated as the difference between the current account and trade balance, is excluded. This is due to the fact that, without further knowledge pertaining to the country's asset position, it might be hard to provide an interpretation of the income balance. Also, it could not be computed accurately. For example, even when service trade data of decent quality are available, the income balance could not be separated from transfer payments. The merchandise trade balance in our analysis, therefore, could reasonably stand for the current account.

[9]Note that we focus on trade balance (not oil or non-oil trade balance). This is due to data availability between the US and Canada. Also, we emphasize that our objective is to examine whether oil price could be claimed as a major source of the US-Canada (trade) imbalance.

[10]See Rebucci and Spatafora (2006).

causes a deterioration in a country's balance of trade. Fratzscher *et al.* (2008) show that declines in US asset prices, such as housing and equities, have an important role in alleviating its trade imbalances. In particular, such declines account for up to 35 percent of the movements of its balance of trade, which represents a larger extent than changes in the US dollar exchange rate over the period 1974-2005.

Finally, we are motivated by a lack of research on possible nonlinearities in the response of the balance of trade to its determinants. In particular, most international trade studies (on J-curve effects in particular) have been carried out using linear vector error-correction models, assuming that macroeconomic variables have linear effects, independent of the magnitudes of the variables. Nevertheless, it is reasonable to believe that the trade balance may differ across the business cycle. Specifically, on the one hand, in export-led growth economies (as in oil and early industrialized countries), the trade balance would improve during an expansion of the economy. On the other hand, in domestic demand-led growth economies (such as in the US and Australia), their balance of trade would worsen at the same stage in the business cycle.[11] Inspired by research on the business cycle, we believe that linear models might not be able to explain its behavior well and, consequently, we might need to use non-linear models to cope with such behavior.[12]

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[11]For the US, there is clear evidence that the US trade deficit drops to its lowest level in more than five years in November 2008 as the economic slowdown leads to lower demand for imports.

[12]See Clements and Krolzig (2002) and Clements and Krolzig (2004).

1.2 Questions We Ask

Motivated by arguments and findings set out in the previous section, where we justify why the exchange rate, oil price and asset price might be important, and (additional) controls in the empirical trade balance equation, we would like to examine empirically, both within a linear and non-linear framework, the following issues. These are (i) if evidence of short-run J-curve effects prevail for the US and Canada; (ii) if oil price could be claimed as a major source of global imbalances, the US-Canada imbalance in particular and (iii) if asset price, house price in particular, explains some of the movements of the US trade balance, as recent evidence has shown. Our novel features include the use of non-linear model in the balance of trade analysis and the use of variables, such as oil price and house price, which have been claimed recently as sources of the US trade deficit, to augment traditional functional relationships for the trade balance, having only domestic income, trading partner's income and real bilateral exchange rate as explanatory variables.[13]

To explore the dynamic relationship between US real trade balance with Canada and real exchange rate, real oil price and real new housing price index and its adjustment, both in the short and long run, empirically, we use cointegration analysis and a linear and non-linear vector error-correction model. This is with quarterly data from quarter one 1985 to quarter one 2005. In addi-

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[13]To our knowledge, no one, to date, has formally examined the (traditional) functional relationship for the trade balance within a non-linear framework, except Moura and Silva (2005).

tion, we would like to evaluate our oil price and asset price-augmented model by performing simple forecasting exercises, compared with forecastability of the traditional functional relationship for the trade balance. We would like to make an implication about global imbalance at the end of the chapter.

We find that real exchange rate, real oil price and real new housing price index have statistically significant effects on the US real trade balance with Canada in the long run. We acquire evidence of short-run J-curve with a percentage change in real trade balance equal to -0.45 and -0.54, following a 1 percent deprecia-tion within linear and non-linear frameworks respectively. Results from both the linear and non-linear models show that short-run dynamic effects of real oil price are not so fearful, with statistically insignificant effects on real trade bal-ance following an increase in real oil price. House price can be argued as being strongly relevant for settlement and adjustment of US trade balance in the long run through the wealth effects, with a distinguished coefficient of US real new housing price index. However, the immediate (next-quarter) effect of a change in housing wealth is insignificant, consistent with existing literature. With the transition probability matrix showing that moving to a regime presenting per-sistent correction is more likely than the opposite, we believe, reasonably, that a (small) chance to correct the US-Canada imbalance prevails.

From a multi-step ahead forecasting exercise, we conclude that the US real trade balance with Canada forecasts from our non-linear VAR model outperform ones from the linear VAR in first difference (DVAR) model and ones from the

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random walk model. Furthermore, our results indicate that the long-term out-of-sample forecastability is not much improved by the additional variables, which nonetheless actively explain in-sample movement of US real trade balance with Canada in the long run.

The remainder of the chapter is set out as follows. In Section 2, we describe the methodology, including variables of interest, cointegration analysis and the linear and non-linear vector error-correction models. In Section 3, we describe the data and report the estimation results, including results from cointegration analysis and results from the linear and non-linear models. This is in addition to discussion. In Section 4, we perform forecasting exercise and offer some concluding comments in Section 5.

2 Methodology

We divide this section into three subsections. These are (i) variables of interest; (ii) cointegration analysis and linear vector error-correction model and (iii) non-linear vector error-correction model.

2.1 Variables of Interest

The selection of variables is usually based on economic theory considerations. This also requires good judgement and sound intuitive analysis. For our empirical analysis, chosen variables are based mainly on existing literature, country

specifics and motivation set out in the Motivation subsection.

In order to make our empirical analysis consistent with the existing literature, we start with a traditional functional relationship for the trade balance. Following Bahmani-Oskooee and Artatrana (2004a), the traditional functional relationship for the trade balance is

$$
\begin{aligned}
Real\ Trade\ Balance = &f(Domestic\ income,\ Trading\ partner's\ income, \\
&Real\ bilateral\ exchange\ rate).
\end{aligned} \tag{1}
$$

We then introduce oil price and asset price to our analysis due to the discussion in the Motivation subsection. We call our specification the oil price and asset price-augmented trade balance equation. Hence, variables are (i) US real trade balance with Canada, (ii) US real industrial production, (iii) Canada real industrial production, (iv) US (domestic) dollar real exchange rate relative to Canadian (foreign) dollar, (v) real oil price, (vi) US real new housing price index and (vii) Canada real new housing price index. Note that such variables are chosen for our empirical analysis, but variables are not necessarily limited to these. Our functional relationship for the oil price and asset price-augmented trade balance is

$$Real\ Trade\ Balance = f(US\ income,\ CA\ income,\ Real\ exchange\ rate,\ Real\ oil\ price,$$

$$US\ real\ housing\ price\ index,\ CA\ real\ housing\ price\ index).$$

$$(2)$$

For the US real trade balance with Canada, we follow Moura and Silva (2005). In particular, we take the export-import ratio to represent trade balance, which is common in literature. The advantage of doing so is that it allows us to take logarithms of trade balance and, hence, to get growth rates (see Brada *et al.* (1997)). With the export-import ratio, remaining constant when measurement units change is another gain. In addition, the export-import ratio is able to serve as either nominal or real trade balance (see Bahmani-Oskooee and Tatchawan (2001)). Alternatively, with the difference between exports and imports to represent trade balance, a deflator is needed in order to calculate the real trade balance. However, this measurement is sensitive to different deflators. The US real trade balance with Canada (RTB) is

$$RTB_t = ln\left(Export\ import\ ratio\right)_t = ln\left(\frac{Export}{Import}\right)_t = ln\left(Export\right)_t - ln\left(Import\right)_t.$$

$$(3)$$

For US and Canada real GDP ($USRY$ and $CARY$), we use industrial production and consumer price index. That is

$$USRY_t = ln\left(Real\ US\ industrial\ production\right)_t = ln\left(\frac{US\ industrial\ production}{US\ CPI}\right)_t$$

(4)

and

$$CARY_t = ln\left(Real\ CA\ industrial\ production\right)_t = ln\left(\frac{CA\ industrial\ production}{CA\ CPI}\right)_t.$$

(5)

For the US dollar real exchange rate relative to the Canadian dollar, we use (domestic) US dollar unit to (foreign) Canadian dollar exchange rate, US consumer price index and Canada consumer price index. The US dollar real exchange rate relative to Canadian dollar (RER), in our context, is defined as the natural logarithm of (domestic) US dollars to one (foreign) Canadian dollar real exchange rate. Hence, a decrease (increase) in exchange rates corresponds to an appreciation (depreciation) of the (domestic) US dollar against the (foreign) Canadian dollar. We write

$$RER_t = ln\left(US\ dollars\ per\ Canadian\ dollar\right)_t + ln\left(CA\ CPI\right)_t - ln\left(US\ CPI\right)_t.$$

(6)

For real oil price, we use the price of West Texas intermediate crude in dollars

per barrel. We use the US consumer price index as the deflator to derive the real oil price (ROP). That is

$$ROP_t = ln\,(Oil\ price)_t - ln\,(US\ CPI)_t.\tag{7}$$

For US and Canada real new housing price index ($USRHPI$ and $CARHPI$), we use the US price index of new one-family houses under construction, Canada new housing price index and consumer price index. That is

$$USRHPI_t = ln\,(Real\ US\ new\ housing\ price\ index)_t =$$
$$= ln\left(\frac{US\ price\ index\ of\ new\ one-family\ houses\ under\ construction}{US\ CPI}\right)_t \tag{8}$$

and

$$CARHPI_t = ln\,(Real\ CA\ new\ housing\ price\ index)_t =$$
$$= ln\left(\frac{CA\ new\ housing\ price\ index}{CA\ CPI}\right)_t. \tag{9}$$

2.2 Cointegration Analysis and Linear Vector Error-Correction Model

To examine the dynamic relationship between US real trade balance with Canada and real exchange rate, real oil price and real new housing price, claimed recently to be sources of the US trade deficit, both in the short and long run, we first perform multivariate cointegration analysis and linear vector error-correction model.

With cointegration analysis, we aim for stationary linear combinations of stochastic process integrated of order d, $I(d)$, non-stationary time series; that stationary combination would be called a cointegrating equation. This could be described as an equilibrium relationship between the variables of interest in the long run (see Engle and Granger (1987)). For our empirical analysis, this might be written as

$$RTB_t = \beta_0 + \beta_1 RER_t + \beta_2 ROP_t + \beta_3 USRHPI_t + \beta_4 CARHPI_t + \beta_5 USRY_t + \beta_6 CARY_t + u_t,$$

$$(10)$$

where u_t is the random error term. Each cointegrating coefficient, β_i, shows the percentage change in US real trade balance with Canada for one unit percentage change in each of the explanatory variables in the long run.

The coefficients of these variables of interest should enter with a sign according to the channels discussed in the Motivation subsection. That is, we expect

that $\beta_1 > 0$ for real exchange rate depreciation to improve real trade balance. This also satisfies the Marshall-Lerner condition. In particular, due to the fact that exchange rate devaluation leads to a fall in export prices, there would be an increase in demanded export quantities. In addition, import prices would rise, diminishing their demanded import quantities. In order to realize the net effect on the balance of trade, the price elasticities would have to be considered. Specifically, export quantities demanded would rise proportionately less than the fall in their prices if exported goods are inelastic to their prices. This would not result in an increase in export total revenue. Likewise, a decrease in import total expenditure would not occur if imported goods are also inelastic to their prices. Overall, both changes that are caused by the devaluation of the exchange rate would diminish the balance of trade position.[14] Without any doubt, β_2 is expected to be negative, due to the fact that in fuel-importing countries the increase in world oil prices worsens the balance of trade and Canada is the top source of US crude oil imports. Because of the wealth effects described in the Motivation subsection, we expect that $\beta_3 < 0$ and $\beta_4 > 0$. For real GDP, following existing literature, we do not have *priori* expectations for β_5 and β_6; coefficients of these two variables are purely empirical.[15]

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[14]It has been found empirically that in the short run goods are likely to be inelastic. This is due to the fact that it takes time for patterns of consumption to change, following changes in price. As a result, the Marshall-Lerner condition is not satisfied, and a devaluation of a currency does not tend to improve the balance of trade initially; this effect is called the J-curve effect. In the long term, consumers would adjust to the new prices, and, consequently, balance of trade would improve.

[15]It should be expected that $\beta_5 < 0$, since an increase in US real GDP usually leads to an increase in imports from Canada. However, it would also be possible that $\beta_5 > 0$ if the increase in US real GDP is attributed to a rise in the import-substitution goods production. This would

To find the number of cointegrating relations and their estimation, we employ the maximum likelihood procedure; see Johansen (1988) and Johansen and Juselius (1990). Assuming Gaussian errors, this method is applied to a vector autoregressive (VAR) model. Specifically, this is based on a pth-order structural and dynamic VAR model involving the variables examined. In association with the Granger representation theorem, an unrestricted vector error-correction model (VECM) with a maximum lag of p could be written. In particular, this is represented in the following form

$$\Delta x_t = \mu + \sum_{i=1}^{p} \Gamma_i \Delta x_{t-i} + \Pi x_{t-1} + \xi_t, \tag{11}$$

where Δ is the first difference operator and x_t is a vector of variables of interest. That is $x_t = [RTB, RER, ROP, USRHPI, CARHPI, USRY, CARY]'$, in our context. μ is a constant term and Γ_i is a vector of coefficients of lagged variables in first difference dynamic terms to be estimated. Π contains an information on a speed-of-adjustment coefficient (α) to the long-run equilibrium and the vector of cointegrating coefficients (β). That is, $\Pi = \alpha\beta'$. ξ_t is a purely white noise term.

We explain the practical procedure for the cointegration analysis and the linear vector error-correction model estimation. Firstly, we consider whether each of the seven time series of interest (US real trade balance with Canada,

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in turn result in fewer imports from Canada as the economy grows. Hence, depending on the magnitude of demand side factors and supply side factors, β_5 could be either negative or positive. This argument also applies to β_6.

US dollar real exchange rate relative to Canadian dollar, real oil price, US real new housing price index, Canada real new housing price index, US real GDP and Canada real GDP) is integrated, that is, the number of difference before achieving stationarity, of the same order. In particular, we test for unit roots to examine the stationary properties of the data we use. To do this, we consider the standard Augmented Dickey-Fuller (ADF) test. Secondly, we use the lag selection criterion as a leading indicator for selecting the robust and stable cointegration test specification. To search for the number of multivariate cointegrating relationships, we use the Johansen (1988) and Johansen and Juselius (1990) procedure. The cointegration rank, r, is the set from zero to $k-1$, where k is the number of endogenous variables and equal to 7 in our context. r is then determined with the trace and maximum eigenvalue test. The trace statistic, λ_{trace}, tests the hypothesis that there are, at most, r cointegrating vectors, the maximum eigenvalue statistic, λ_{max}, tests that there are r cointegrating vectors against the alternative that there are $r+1$. The asymptotic critical values are given in Johansen and Juselius (1990) and MacKinnon $et\ al.$ (1999). The forms of the test statistics are

$$LR_{tr}\left(r|k\right) = -T \sum_{i=r+1}^{k} log\left(1 - \lambda_i\right), \tag{12}$$

where λ_i denotes the i-th largest eigenvalue of the Π matrix and

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$$LR_{max}\left(r|r+1\right) = -T\,log\left(1-\lambda_{r+1}\right) = LR_{tr}\left(r|k\right) - LR_{tr}\left(r+1|k\right), \qquad (13)$$

for $r = 0,\ 1,\ ...\ k-1$. Note that lag selection for cointegration test specification would also be used in the vector autoregressive (VAR) in first difference part when estimating the vector error-correction model. This is in order to obtain the same long-run (error-correction mechanism) relationship. Thirdly, we specify the long-run relationship between US real trade balance with Canada and the other six variables; in particular, this is the normalized cointegrating relationship estimated. Finally, we estimate the vector error-correction model by the Johansen procedure and examine the estimated coefficients of lagged variables in first difference dynamic terms in Γ_i and the coefficient estimate of the error-correction term, that is, the adjustment parameter, α.

2.3 Non-Linear Vector Error-Correction Model

We estimate a non-linear model for the reasons given in the Motivation subsection. This is in order to investigate the regime changes of our six variables and, in particular, US real trade balance with Canada. In addition, this is in order to take into account the cyclical pattern in the time series examined. A cointegrated VAR(p) with M Markovian regimes model is named MS(M)-VECM(p). This is the model we employ in particular. Generally, the MS-VECM model is

a vector error (equilibrium)-correction model with shifts in the drift $v(s_t)$ and in the long-run equilibrium $\mu(s_t)$. Note that the linear VECM(p) collapses to the particular case where $M = 1$. We write the MS-VECM as

$$\Delta x_t - \mu(s_t) = \alpha(s_t)\beta x_{t-1} + v(s_t) + \sum_{k=1}^{p} A(s_t)_k (\Delta x_{t-k} - \mu(s_{t-k})) + u_t \quad (14)$$

and the innovations u_t are conditionally Gaussian, $u_t|s_t \sim \text{NID}(0, \sum(s_t))$. The (autoregressive) parameters are time-varying and also depend upon changes in a stochastic, unobservable regime variable, $s_t \in \{1, \dots M\}$. This is with a finite number of states. The stochastic process for generating changes in the unobservable regimes is an ergodic Markov chain defined by the transition probabilities

$$p_{ij} = \Pr(s_{t+1} = j|s_t = i), \sum_{j=1}^{M} p_{ij} = 1 \ \forall i,j \in \{1, \dots M\}. \quad (15)$$

where $p_{iM} = 1 - p_{i1} - \dots - p_{iM-1}$ for $i = 1, \dots M$. The regimes could be reconstructed. This is with the unobserved regimes probabilities conditional on information set available; see Krolzig (1997).

The important issue is how to decide upon the proper number of regimes in, and specification of, an MS-VECM. Nevertheless, this could not be found directly. With the presence of nuisance parameters, the likelihood ratio could not be used to serve as a criterion. This is due to the fact that it would turn out to be with an unknown distribution, that is, it does not have a standard

asymptotic chi-squared distribution. With such situations, Hansen (1992) and Garcia (1998) have developed an approach by which to attain the asymptotic distribution. Nevertheless, for practical purposes, this might not be useful, since the distribution depends on both data and parameters of the model. As a result, every time a test is performed, asymptotic distribution needs to be obtained (see Krolzig (1997)). However, in line with literature, both the Akaike information criterion (AIC) and Schwarz criterion (SC) are believed to well serve as criteria. In particular, they do not underestimate the minimum number of regimes to be examined in the MS-VECM (see Ryden (1995)). We use such criteria for our analysis, therefore.

3   Data and Results

We divide this section into four subsections. These are (i) data; (ii) results from cointegration analysis; (iii) results from linear model and (iv) results from non-linear model.

3.1   Data

As in the previous section, variables are (i) US real trade balance with Canada, (ii) US real GDP, (iii) Canada real GDP, (iv) US dollar real exchange rate relative to Canadian dollar, (v) real oil price, (vi) US real new housing price index and (vii) Canada real new housing price index. Our dataset is from quarter

one 1985 to quarter one 2008 (93 observations). Figure 3 shows plots of these variables of interest, transformed as set out in Variables of Interest subsection, in level and in first difference.

For US real trade balance with Canada, monthly US nominal exports and imports (with Canada) data in millions of US dollars are taken from the US Census Bureau, Foreign Trade Division. We calculate quarterly averages from monthly data and derive the export-import ratio representing the real trade balance.

For US and Canada real GDP, we use quarterly US industrial production, quarterly Canada industrial production, quarterly US consumer price index and quarterly Canada consumer price index data from IMF International Financial Statistics. Note that they are in the same unit, price index, quarter two 2000=100.

For US dollar real exchange rate relative to Canadian dollar, quarterly US dollar unit to Canadian dollar average exchange rate, quarterly US consumer price index and quarterly Canada consumer price index data are taken from IMF International Financial Statistics. Note that US consumer price index and Canada consumer price index are in the same unit, index, quarter two 2000=100.

For real oil price, the quarterly current price of West Texas intermediate crude in dollars per barrel is taken from IMF International Financial Statistics. For deflator, we use quarterly US consumer price index. Note that for US consumer price index, quarter two 2000=100.

For US and Canada real new housing price index, the quarterly US price index of new one-family houses under construction and Canada new housing price index are from Bureau of the Census, the US Department of Commerce and Cansim - Statistics Canada - respectively. Note that both are in the same unit, price index. Note also that, from the original dataset where 2005=100 for the US price index of new one-family houses under construction and 1997=100 for Canada new housing price index, we construct time series of interest, where quarter two 2000=100.

Note that, in estimating the model, we use data from quarter one 1985 to quarter one 2005 (81 observations). The last three-year observations (12 observations) are kept for forecasting exercise.

3.2 Results from Cointegration Analysis

We analyze the stationarity in our seven variables, US real trade balance with Canada, US real GDP, Canada real GDP, US dollar real exchange rate relative to Canadian dollar, real oil price, US real new housing price index and Canada real new housing price index. We use the Augmented Dickey-Fuller (ADF) test, with an intercept. The maximum lag length for the Schwarz Information Criterion is 15. The results are shown in Table 3, which report the ADF test statistic with the unit root null hypothesis and MacKinnon (1996) one-sided p-values. We find that our variables have a unit root or, alternatively, are integrated of order one, $I(1)$, at the 5 percent significance level. This is especially clear if we

use the Phillips and Perron test.[16]

We analyze cointegration among our seven variables of interest, for which we use the Johansen cointegration test. We choose the cointegration test specification with 1 lag.[17] Moreover, we allow for linear deterministic trend in the data and for an intercept. The results with MacKinnon *et al.* (1999) p-values are presented in Table 4.

The results from the Trace and Maximum Eigenvalue statistics suggest (rank) $r \leq 0$ is strongly rejected but $r \leq 1$ is not rejected (for Maximum Eigenvalue statistics) at the 5 percent significance level. We conclude, therefore, that there is one cointegrating equation at the 5 percent significance level.[18]

3.3   Results from Linear Model

We provide results from the linear model and form an analysis. Table 5 reports the estimated cointegrating vector, β, normalized cointegrating coefficients, on

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[16]There is evidence that US real GDP and US real new housing price index may be integrated of an order higher than one, $I(2)$ in particular. With the Phillips and Perron test, all variables are integrated of order one, $I(1)$. The significance level for this test for the level (first difference) of *RTB*, *RER*, *ROP*, *USRHPI*, *CARHPI*, *USRY* and *CARY*, under the null hypothesis of unit root is 0.1919 (0.0000), 0.8269 (0.0000), 0.7369 (0.0000), 0.8195 (0.0000), 0.8110 (0.0068), 0.5118 (0.0003) and 0.5441 (0.0002), respectively.

[17]The lag length selection criteria we use for this specification are Schwarz Information Criterion and Hannan-Quinn Information Criterion, selecting one lag order. In addition, we choose the specification with one lag suggested due also to sound economic intuition achieved in the following subsection. Furthermore, with regime-dependent (autoregressive) parameters in the non-linear model, this might relieve the problem of too few degrees of freedom (too many parameters) that might occur. Akaike Information Criterion selects seven lags order.

[18]The Trace statistic suggests there may be two cointegrating relations, but the subsequent analysis was more economically meaningful if we assumed a single cointegrating equation. This will be emphasized again in the following subsection.

US real trade balance with Canada.[19]

From results in Table 5, we could write our cointegrating relationship as

$$RTB_t = -0.46 + 0.45RER_t - 0.20ROP_t-$$

$$0.70USRHPI_t + 0.59CARHPI_t + 0.20USRY + 0.01CARY. \tag{16}$$

We check whether all the coefficients of the variables of interest are correctly signed. Note that US dollar real exchange rate relative to Canadian dollar is natural logarithm of (domestic) US dollars to one (foreign) Canadian dollar real exchange rate. That is, a decrease (increase) in exchange rates corresponds to an appreciation (depreciation) of (domestic) US dollar against (foreign) Canadian dollar.

A 1 percent depreciation of US dollar real exchange rate relative to Canadian dollar would lead to a 0.45 percent increase of the US real trade balance with Canada in the long run, with the t-statistic of 2.80. This implies that the Marshall-Lerner condition for a currency devaluation to have a positive impact on trade balance between the US and Canada holds in the long run.

A 1 percent increase in real oil price would lead to a 0.20 percent decrease of US real trade balance with Canada in the long run. The estimate of this coefficient is statistically significant, with the t-statistic of -6.16.

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[19]We select only one cointegrating equation after experimenting also two cointegration equations. The latter yields a real trade balance relationship that appears to counter the *priori* expectations explained in the Cointegration Analysis and Linear Vector Error-Correction Model subsection. We also find that no additional gain is achieved, quantitatively and qualitatively, examining two cointegration equations. In addition, we, in this chapter, would like to examine only the adjustment of US real trade balance with Canada in the long run.

For the real new housing price index, supporting the wealth effects discussed in the Motivation subsection, results are of interest and are noteworthy. Following a 1 percent increase in US real new housing price index, US real trade balance with Canada deteriorates by as much as 0.70 percent in the long run. On the contrary, US real trade balance with Canada improves, in the long run, 0.59 percent with a 1 percent increase in Canada real new housing price index. The estimates of these two countries' real new housing price index coefficients are statistically significant, with the t-statistic of -2.09 and 4.03, respectively. House price in the US could be argued, therefore, as being strongly relevant for settlement and adjustment of US trade balance in the long run due to the fact that US real new housing price index has the highest effect on US real trade balance with Canada, compared with other variables.

For US and Canada real GDP, without a *priori* expectations for signs, as discussed in the Cointegration Analysis and Linear Vector Error-Correction Model subsection, the estimates show that an increase in US real GDP leads to a decrease of imports from its trading partner, Canada, whereas a rise in Canada real GDP would lead to an increase of imports from its trading partner, the US. However, the interpretation should be treated with caution since these estimated real GDP coefficients of the US and Canada are not statistically significant, with the t-statistic of 0.86 and 0.04 respectively.

Comparing these six explanatory variables of interest, we believe that all variables, but US and Canada real GDP, explain the long-run dynamics of US

real trade balance with Canada statistically significantly. Noticeably, real new housing price index, the US one in particular, could be argued as being strongly relevant for settlement and adjustment of US trade balance in the long run. In other words, the meaning and validity of the wealth effects are strongly supported in the long run.

We now analyze the short-run dynamics of US real trade balance with Canada. The existence of at least one cointegrating vector among the seven variables considered implies that a vector error-correction model could be estimated to investigate the short-run relationship.

We estimate the vector error-correction model. The specification capturing the short-run dynamics with 1 lag is exactly the same as one for the Johansen cointegration test. Looking at the system, we find that US real trade balance with Canada shows an adjustment due to the statistically significant estimate of the error-correction term coefficient, with the t-statistic of -2.85.[20] Table 6 reports the error-correction model estimates (the lagged variables in first difference dynamic terms) for one cointegrating vector and standard errors.

From results in Table 6, we find that only past values of the US dollar real exchange rate relative to the Canadian dollar at lag 1 offer a statistically significant explanation of the movement of US real trade balance with Canada

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[20]With regard to US dollar real exchange rate relative to Canadian dollar, the estimate of the error-correction term coefficient (-0.14) is also statistically significant at the 5 percent significance level, with the t-statistic of -2.34. This empirical evidence could be taken as a reason for why we could extend our analysis to examine also the US dollar real exchange rate relative to Canadian dollar error-correction model, in addition to US real trade balance with Canada, which is the main interest.

in the short run at the 10 percent significance level; this is with the t-statistics of -1.88. Past value of real oil price, US real new housing price index, Canada real new housing price index, US real GDP and Canada real GDP, however, do not explain statistically significantly the short-run dynamics of real trade balance between the US and Canada; this is seen with the t-statistics of 1.47, -0.27, -0.61, 0.83 and 0.08 respectively. The magnitude of 0.45 for lagged US dollar real exchange rate relative to Canadian dollar in first difference is noteworthy. In addition, the fact that our results from the linear model provide statistically insignificant positive estimate of lagged real oil price in first difference could be interpreted as short-run dynamic effects of real oil price on US real trade balance with Canada not being so fearful.

We would like to make an argument that results from our linear vector error-correction model show evidence of the J-curve effect, where, in response to exchange rate depreciation, the trade balance measured in (local) US dollars worsens in the impact period. This effect could, therefore, be thought of as a short-run departure from the Marshall-Lerner condition discussed above. This is due to the fact that the coefficient estimate of the significant past value (in first difference) of US dollar real exchange rate relative to Canadian dollar is negative, with the magnitude of 0.45 as reported earlier. Note also that, in our context, where US dollar real exchange rate relative to Canadian dollar is defined as the natural logarithm of (domestic) US dollars to one (foreign) Canadian dollar real exchange rate, an increase in exchange rates corresponds to a

depreciation of (domestic) US dollar against (foreign) Canadian dollar.

The impulse-response function is a convenient way of showing an estimate of a variable coefficient of a VAR model. In our context, it is particularly useful to capture the short-run dynamics of the response of US real trade balance with Canada to shocks in relevant variables. Figure 4 shows the response of US real trade balance with Canada to innovations of one standard deviation in US dollar real exchange rate with Canadian dollar and in real oil price.

Note that the impulse-response functions shown maintain their long-run equilibrium. This is in contrast to those that gradually decrease their response magnitude in the long run commonly found in the vector autoregression model. However, this is not surprising due to the fact that the long-run cointegrating relationship is included in the vector error-correction model. In other words, the long-run behavior of the endogenous variables is restricted, converging to their long-run cointegrating relationship consequently; see Hsing (2005).

We emphasize our impulse-response analysis on the short-run response of real trade balance to real exchange rate innovations. We see that US real trade balance with Canada decreases in the aftermath of shocks to US dollar real exchange rate relative to Canadian dollar. Note the second quarter. From accumulated figure, the minimum occurs after nearly four quarters. We argue, therefore, that the impulse-response function shows the evidence of the J-curve effect in this linear vector error-correction model; afterwards, the Marshall-Lerner condition holds. This is due to the fact that the temporary negative impact of US dollar

real exchange rate with Canadian dollar shocks to US real trade balance with Canada is believed to be fully faded away after eight quarters have vanished from the accumulated figure. We also note, in the two quarters immediately following the shocks, the insignificant response of US real trade balance with Canada to real oil price shocks, from both accumulated and non-accumulated figures.

The estimate of the error-correction term coefficient (-0.34) is statistically significant, with the t-statistic of -2.85 as reported earlier. This is with the appropriate (negative) sign indicating existence of a long-run relationship. Short-run US real trade imbalance with Canada is corrected quarterly at a rate of 34 percent.

We assess the performance of our linear model, performing the presence of serial correlation test and heteroskedasticity test.[21] The model does not have autocorrelations but a problem of heteroskedasticity. This might be due to the very long period of data analyzed.

3.4 Results from Non-Linear Model

We estimate the non-linear model. Since we are interested to see whether our six variables of interest have a different effect on US real trade balance with Canada

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[21]The multivariate LM test statistics (probability) for residual serial correlation up to 12 order, under the null hypothesis of no serial correlation of order h, are 0.6804, 0.4605, 0.7830, 0.4334, 0.7971, 0.9920, 0.7138, 0.2342, 0.0572, 0.3449, 0.3127 and 0.3441. With the White heteroskedasticity test, under the null of no heteroskedasticity, the non-constant regressors are jointly significant, with probability 0.0008.

in separate regimes, we should estimate only models that allow for regime-dependent parameters. Hence, we choose the specification from the two specifications, namely, MSIA(2)-VECM(1) and MSIAH(2)-VECM(1). Both specifications are with regime-dependent intercept term and with regime-dependent (autoregressive) parameters. The first specification is homoskedastic and the second one heteroskedastic. A regime-dependent covariance structure of the process could be regarded as an added model feature (Krolzig (1998)). We are interested only in the two-regime specification. This is due mainly to the fact that the following results show that duration is very short; consequently, it might not be reasonable to increase the number of regimes. In addition, we use the same lag length as one in the linear model, that is, the specification with 1 lag. This is to relieve the problem of too few degrees of freedom (too many parameters) that might occur. This is also for dynamic comparison reasons.

We estimate both specifications and report results from both AIC and SC criteria in Table 7. Note that the error-correction term (RTB_{t-1}) in the MS-VECM is the same one as in the linear VECM. In particular, we use the error-correction term (cointegrating equation as (16)) from the linear VECM as an exogenous variable when estimating the MS-VECM. This is the two-stage procedure proposed in Krolzig (1999).

From these results, since the second specification MSIAH(2)-VECM(1) shows lower values for both the AIC criterion (-37.1867) and the SC criterion (-31.6679), we should look at such specification for our analysis. Nevertheless, we also put

an eye on the first specification MSIA(2)-VECM(1). After deliberately considering both specifications, we reasonably select the simpler one of MSIA(2)-VECM(1) specification for our analysis. This is due mainly to smoother probabilities (shown below) obtained from this specification, compared with ones from the MSIAH(2)-VECM(1); results pertaining to the MSIAH(2)-VECM(1) are also reported in Table 9. Nevertheless, we find that no additional gain is achieved and that no qualitative impact on the results is found, when examining the MSIAH(2)-VECM(1).[22] Our chosen model MSIA(2)-VECM(1) is written as

$$\Delta x_t = \alpha\left(s_t\right)\beta x_{t-1} + v\left(s_t\right) + \sum_{k=1}^{1} A\left(s_t\right)_k \Delta x_{t-k} + u_t, \qquad (17)$$

where $x_t = [RTB, RER, ROP, USRHPI, CARHPI, USRY, CARY]'$, $\alpha\left(s_t\right)$ is regime-dependent speed-of-adjustment coefficient to long-run equilibrium, β is vector of cointegrating coefficients, $v\left(s_t\right)$ is regime-dependent intercept term, $A\left(s_t\right)$ is vector of regime-dependent (autoregressive) parameters, $s_t \in \{1,\ 2\}$ and u_t is conditionally Gaussian for the MSIAH(2)-VECM(1) specification, $u_t|s_t \sim\text{NID}(0, \sum\left(s_t\right))$. The stochastic process for generating changes in the unobservable regimes is an ergodic Markov chain defined by the transition probabilities $p_{ij} = \Pr\left(s_{t+1} = j|s_t = i\right)$, $\sum_{j=1}^{2} p_{ij} = 1 \ \forall i, j \in \{1,\ 2\}$, where $p_{i2} = 1 - p_{i1}$

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[22]The MSIA(2)-VECM(1) specification provides longer durations for both regime one and regime two, compared with ones from the MSIAH(2)-VECM(1). We believe that specification providing longer durations might be meaningful in forecasting exercise in the next section. Applying this here, we are confident that the selected specification MSIA(2)-VECM(1) providing longer durations in two regimes might not lose forecastability. This supports the specification we select for our analysis. Regime one duration from the MSIA(2)-VECM(1) is 4.27 quarters, regime two 5.42 quarters. Regime one duration from the MSIAH(2)-VECM(1) is 2.13 quarters, regime two 3.52 quarters.

for $i = 1, 2$. We normalize the model to make US real trade balance with Canada dependent variable. We try to construct an analysis from results from the non-linear model reported in Table 8.

In order to justify the use of the non-linear model, we report the p-value of likelihood ratio (LR) test. Under nuisance parameters, Davies (1977) derives an upper bound for the likelihood ratio (LR) test statistic significance level. This is employed due to the fact that, in presence of nuisance parameters, a standard chi-square distribution could not be derived with the likelihood ratio (LR) statistics even asymptotically. See Garcia (1998). We find that the linearity null hypothesis is strongly rejected. In other words, the non-linear model of MSIA(2)-VECM(1) seems to outperform its linear counterpart.[23]

From the chosen MSIA(2)-VECM(1) specification, intercept is negative both in regime one and in regime two. The intercept in regime two does not prove to be statistically significant with the t-statistic of -0.17. However, the negative intercept in regime one is statistically significant at the 10 percent significance level with the t-statistic of -1.87. The standard deviation of both regimes is 0.04.

The coefficient estimates of the past value (in first difference) of US dollar real exchange rate relative to Canadian dollar explain US real trade balance with Canada movement statistically significantly only in regime two, at the 5

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[23]Supporting the fact that non-linear model is preferred to its linear counterpart, lower Akaike information criterion statistic and higher log-likelihood are achieved using the non-linear model.

percent significance level. It is worth mentioning that this statistically significant negative relationship with magnitude of 0.54 is bigger than that acquired from the linear model.[24] As in Results from Linear Model subsection, we argue that this is strong evidence of J-curve effects. While statistically insignificant, the effect of US dollar real exchange rate relative to Canadian dollar on US real trade balance with Canada in regime one is -0.46, also supporting the evidence of J-curve effects (regardless of the regime). In our context, these results are particularly useful in explaining the short-run dynamics of the response of US real trade balance with Canada to US dollar real exchange rate with Canadian dollar shocks in this non-linear vector error-correction model. That is, there is an impact-period temporary worsening of US real trade balance with Canada.

For real oil price, US real new housing price index and Canada real new housing price index, the coefficient estimates of the past value (in first difference) do not prove to be statistically significant either in regime one or in regime two. These results, obtained from the non-linear model, are in line with those obtained from the linear model. That is, short-run dynamic effects of real oil price on US real trade balance with Canada are not so fearful. It is interesting that the coefficient estimate of the past value (in first difference) of these explanatory variables is negative in regime one, but positive in regime two. For real GDP, the coefficient estimates of the past value (in first difference) of US real GDP are

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[24]Results pertaining to the coefficient estimate of the past value (in first difference) of US dollar real exchange rate relative to Canadian dollar from the linear model prove to be statistically significant at the 10 percent significance level.

statistically significant at the 5 percent significance level in both regimes, being negative in regime one and positive in regime two with an elastic coefficient. Note that we could not see this in the linear model.

The estimate of the error-correction term coefficient in regime two (-0.38) is statistically significant at the 5 percent significance level, with the t-statistic of -2.82. This is, with the appropriate (negative) sign, indicating the existence of a long-run relationship. Hence, it is reasonable to call regime two an imbalance-correction regime.[25] Short-run US real trade imbalance with Canada is corrected quarterly at a rate of 38 percent.

Table 10 reports the transition probability matrix and the expected duration of both regimes. In addition, Figure 5 shows smoothed probabilities of the regimes.[26] Regime one persistency (0.7656) is smaller than that of regime two (0.8157). Note also that moving from regime two to regime one (0.1843) is less likely than the opposite (0.2344). Both regimes last about five quarters (4.27 for regime one and 5.42 for regime two). Regime two has a slightly longer expected duration. 35.9 quarters of 79 quarters included in this analysis (45.44 percent) are categorized in regime one and 43.1 quarters (54.56 percent) in regime two. Note that, although our dataset for estimation is from quarter one 1985 to

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[25]This interpretation is justified by the evidence that the expansion of the difference between the actual value of US real trade balance with Canada and the long-run equilibrium value implied by the cointegrating equation corresponds to regime two (later explained), which presents an imbalance correction. In particular, such difference becomes more negative in the period of the worsening balance of trade deficit, following a short period of improved trade balance with Canada (quarter two 2000 to quarter four 2001).

[26]Figure 6 shows smoothed probabilities of the regimes of MSIAH(2)-VECM(1).

quarter one 2005 (81 observations), we miss the two initial data points. This is because of the 1 lag and differentiation of data.

Based on the smoothed and the filtered probabilities for each regime illustrated in Figure 5, samples are peaks that swing between both regimes. Nevertheless, classification of the regimes roughly captures National Bureau of Economic Research (NBER) US business cycle expansions and contractions. This supports the argument set out in the Motivation subsection that the balance of trade is likely to differ across the business cycle. In particular, from quarter one 1994, the period of worsening balance of trade deficit corresponds to regime two, which presents an imbalance correction, roughly spanning from quarter one 1994 to quarter one 2000 and from quarter one 2002 to quarter one 2005. In addition, regime one, roughly spanning from quarter two 2000 to quarter four 2001, is at a time when the US was enjoying a short period of improving trade balance with Canada;[27] see also Figure 2. We elaborate this in Table 11, showing periods of improving and worsening US trade balance with Canada, NBER US business cycle and direction of our variables of interest for the corresponding period.

According to results from both linear and non-linear models, we would like to make note of the following interesting findings. Firstly, real exchange rate, real oil price and real new housing price index have statistically significant effects on US real trade balance with Canada in the long run. Secondly, we find evidence

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[27]For the US, clear and more recent evidence relating to this is that US trade deficit drops to its lowest level in more than five years in November 2008 as the economic slowdown leads to lower demand for imports.

of short-run J-curve with a percentage change in real trade balance equal to -0.45 and -0.54 following a 1 percent depreciation within linear and non-linear frameworks respectively. Thirdly, results from both linear and non-linear models show that short-run dynamic effects of real oil price are not so fearful, with statistically insignificant effects on real trade balance following an increase in real oil price. Fourthly, house price could be argued as being strongly relevant for settlement and adjustment of US trade balance in the long run through the wealth effects, with a distinguished coefficient of US real new housing price index. However, the immediate (next-quarter) effect of a change in housing wealth is insignificant, consistent with existing literature.[28] Finally, with the transition probability matrix showing that moving to a regime presenting persistent correction is more likely than the opposite, we believe, reasonably, that a (small) chance to correct US-Canada imbalance prevails.

4   Forecasting

We perform a multi-step-ahead forecasting exercise, examining out-of-sample forecasts for US real trade balance with Canada from 1 to 12-quarters ahead.[29] That is, we forecast the observations quarter two 2005 to quarter one 2008,

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[28]See Carroll *et al.* (2006).

[29]We are aware that one-step-ahead forecasts might be more informative than multi-step-ahead forecasts in some situations. However, we believe that multi-step-ahead forecasts up to 12 quarters ahead are worth examining. In addition, if our objective is to see the direction of the US-Canada trade balance, the multi-step-ahead forecasting exercise might be more appropriate than the one-step-ahead forecasting, providing the general long-run behavior of the actual series. We are aware that the short-run fluctuations in the actual series might not be captured.

assuming the forecast origin is quarter one 2005.

In using multi-step-ahead forecast technique, the initial observation in the forecast sample would use the actual values of the lagged endogenous variables, whereas forecasts for subsequent periods would use the previously forecasted values of the lagged endogenous variables. Hence, information based on data through quarter one 2005 is used when computing such multi-step-ahead forecasts. This generates a sequence of 1 to 12-step-ahead forecasts.

In evaluating the forecastability, the criterion we use is the root mean squared error (RMSE).[30] The RMSE statistics are used as a relative measure by which to compare the forecasts across different models. The smaller the forecast error, the better the forecast performance of that model according to that criterion. The RMSE is

$$RMSE = \sqrt{\sum_{t=T+1}^{T+h} (\hat{y}_t - y_t)^2 / h} \tag{18}$$

With four models, namely, linear VECM, linear VAR in first differences (DVAR) model, non-linear VAR (MSI(2)-VAR(1)) model and random walk (in first difference) model, we present the forecast results for our variable of interest in this chapter, US real trade balance with Canada, in Table 12.[31] These are for the 12-quarter-ahead forecast horizon. The reason for using the non-linear VAR

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[30]Note that our evidence on the forecast performance is based on this criterion. We do not calculate or perform other forecast accuracy measures for multi-step-ahead forecasts - the Diebold-Mariano test, for example. This is because it is difficult to calculate them with very few forecasts generated.

[31]Instead of using MSIA(2)-VAR(1) model, we use the simpler one of MSI(2)-VAR(1).

model, instead of the non-linear VECM, is shown in what follows, providing evidence that, in our context, cointegration does not help forecast US real trade balance with Canada.

From results in Table 12, we start by comparing the forecast performance of the linear VECM against one of the DVAR model. This is in order to examine whether cointegration does help forecast US real trade balance with Canada in our context (or whether the differences in forecastability are small) over the examined forecast horizon.[32]

We find that the DVAR model provides more accurate forecasts than the linear VECM. This evidence indicates that cointegration does not help forecast US real trade balance with Canada in our context. This implies that, at least in our empirical work, the cointegrating relationship is informative in explaining the dynamic adjustment of the variable of interest, but not informative in forecasting. In addition, this explains why we could reasonably use non-linear VAR model, instead of the non-linear VECM, for generating forecasts.

We then compare the forecastability of the DVAR model against one of the non-linear VAR model. Results pertaining to the root mean squared error support our expectation that the non-linear VAR model *might* provide more accurate forecasts than the DVAR model.

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[32]Different empirical works on the contribution of cointegration to a forecastability of time series models provide some conflicting results. Nevertheless, one common conclusion seems to be achieved. That is, for long-horizon forecasting exercise, error correction models taking into account cointegrating restrictions should result in forecasts of higher accuracy, compared with a model excluding those restrictions. See Engle and Yoo (1987), Clements and Hendry (1995), Hoffman and Rasche (1996) and Lin and Tsay (1996).

We look at the error ratio which is calculated as the RMSE for the model being examined (the non-linear VAR model) divided by the RMSE for the random walk model.[33] In our non-linear VAR model only, the error ratio of less than one means that this model performs better in forecasting when compared with the random walk model, according to the RMSE criterion. We provide 12-quarter-ahead forecasts generated from the non-linear VAR model and actual US real trade balance with Canada in Figure 7, which shows that the multi-step-ahead forecasts could generate the long-run movement of the actual series of interest. In particular, the upward trend of US real trade balance with Canada is correctly viewed. However, the dramatic fall in US real trade balance with Canada between quarter two 2005 and quarter two 2006 is not captured. This might be unsurprising, given that such a large deterioration was rare historically. In addition, there is systematic under-forecasting throughout 2007. Again, this might not be fatal, since that period of under-forecasting corresponds to the time when the US has the recent continued improvement of its trade balance with Canada, which is owing to an unexpected economic downturn in the wake of the recent global financial crisis.

Relating to our results from the multi-step-ahead forecasting exercise, two possible explanations are mentioned.

Firstly, our results show that the model based on cointegration faces forecast-

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[33]Note that random walk forecasts are from the random walk in first difference model. This model yields more accurate forecasts than the random walk in level model. In addition, we use that model as the benchmark, as we are evaluating forecasts in terms of the first difference of US real trade balance with Canada, not the level.

ing problems. Literature related to this focus on structural breaks or (unforeseen location) shifts in the underlying equilibrium mean, which is viewed as a permanent change in the parameter vector of a model, as the reason.[34] Clements and Hendry (2006) predict that model without cointegration could outperform the VECM if there were structural breaks. This would still apply even some time after a shift. Visualizing the US real trade balance with Canada plotted in Figure 3, we observe a downward mean shift in 2001. This might explain the forecasting problems from which the VECM suffers.

Secondly, treating the time series as being governed by a non-linear model would have implications for forecasting. That is, the possibility of a switch in regime, or state, during the future is included in forecasts from non-linear models. Studies comparing the forecastability of linear and non-linear models, nevertheless, provide mixed results.[35] Clements and Smith (1999) review a number of reasons why apparent nonlinearities in the data might not be able to be exploited to generate more accurate forecasts. One of those that might relate to our results is that forecast performance might depend on the regime which the forecast origin falls in. For our case, this is regime two, which is the regime with more data points. In particular, approximately 55 percent of the sample data points fall in this particular regime. Therefore, we could expect the linear model to match that regime model closely. In other words, this would result in

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[34]See Clements and Hendry (2006) for forecasting with breaks.

[35]See, for example, Stock and Watson (1999) and De Gooijer and Kumar (1992) for evidence and issues pertaining to the comparison of the forecast performance of linear and non-linear models.

the broadly similar forecastability of the two models when the dynamics of the US real trade balance with Canada is in that particular state. Consequently, our results demonstrating that the non-linear model does not forecast markedly more accurately than the linear model are, perhaps, not surprising.

From this multi-step-ahead forecasting exercise, we conclude that US real trade balance with Canada forecasts from our non-linear VAR model outperform ones from the DVAR model and ones from the random walk model.

We also compare the accuracy of forecasts from our oil price and asset price-augmented trade balance equation, which includes real oil price, US real new housing price index and Canada real new housing price index, against the traditional functional relationship for the trade balance, having only US real GDP, Canada real GDP and US dollar real exchange rate relative to Canadian dollar, as explanatory variables. We call the traditional functional relationship the baseline model.

As a formal comparison, results pertaining to the error ratio calculated as the RMSE for the oil price and asset price-augmented model divided by the RMSE for the baseline model are provided in Table 13. With the value close to one for both the linear and non-linear cases, these results indicate that, in terms of a loss function RMSE, the long-term out-of-sample forecastability is not much improved by the additional variables, which nonetheless actively explain in-sample movement of US real trade balance with Canada in the long-run. The reason for that might be the unusual oil price rise and asset price crisis during

the forecast period. If the oil price and asset price-augmented model could have forecasted that event itself, it might have fared better.

With results from the multi-step-ahead forecasting exercise, we have learned that, while empirical evidence found appears to agree with economic intuition and *priori* expectations in-sample, out-of-sample forecast performance might be questioned in some particular circumstances. Nevertheless, evaluating the model by performing forecasting exercise should not be neglected in an empirical work.

5   Conclusion

In this chapter, we provide answers to questions relating to the imbalance in US-Canada bilateral trade, real exchange rate, real oil price and real asset price. We find that real exchange rate, real oil price and real new housing price index have statistically significant effects on US real trade balance with Canada in the long run. We acquire evidence of short-run J-curve effect with a percentage change in real trade balance equal to -0.45 and -0.54, following a 1 percent depreciation within linear and non-linear framework, respectively. Results from both linear and non-linear models show that short-run dynamic effects of real oil price are not so fearful, with statistically insignificant effects on real trade balance following an increase in real oil price. House price could be argued as being strongly relevant for settlement and adjustment of US trade balance in the long run through wealth effects, with a distinguished coefficient of US real new

housing price index. However, the immediate (next-quarter) effect of a change in housing wealth is insignificant, consistent with existing literature. With the transition probability matrix showing that moving to a regime presenting persistent correction is more likely than the opposite, we believe, reasonably, that a (small) chance to correct US-Canada imbalance prevails.

From the multi-step ahead forecasting exercise, we conclude that US real trade balance with Canada forecasts from our non-linear VAR model outperform ones from the linear VAR in first difference (DVAR) model and ones from the random walk model. Furthermore, our results indicate that the long-term out-of-sample forecastability is not much improved by the additional variables, which nonetheless actively explain in-sample movement of US real trade balance with Canada in the long run.

From our findings, we support the argument that J-curve effects from the most recent dollar decline, higher oil prices and asset prices could produce a temporary increase in the nominal and real trade balance deficits. We also believe reasonably that, once J-curve effects disappear, a country's trade balance would improve. Nevertheless, besides exchange rate being manipulated so as to gain balance, policies relating to our empirical analysis should be addressed. Since it is less likely that oil-consuming nations could have an influence on world oil prices, the US could increase refining capacity and provide an effective means to restrain the demand for oil in the medium term. Improving conservation and a shift towards higher energy efficiency are suggested. This could also be good

for the environment. With regard to asset price, private savings associated with wealth effects should increase in general. This is in addition to fiscal consolidation and could be assisted mainly by US tax system reforms. Our empirical analysis argues that current asset (housing) price bubble reflation would not help improve US trade deficit (with Canada in particular) over the long run. With appropriate actions, US trade balance deficit (if it still exists) might, then, not be so bad, stimulating its demand and global, in particular Asian, exports.

Figure 1: (a) US International Balance of Trade in Goods and Services (Percent of GDP), (b) Texas Spot Oil Price (US Dollar per Barrel) and (c) US Real New Housing Price Index

(a)

(b)

(c)
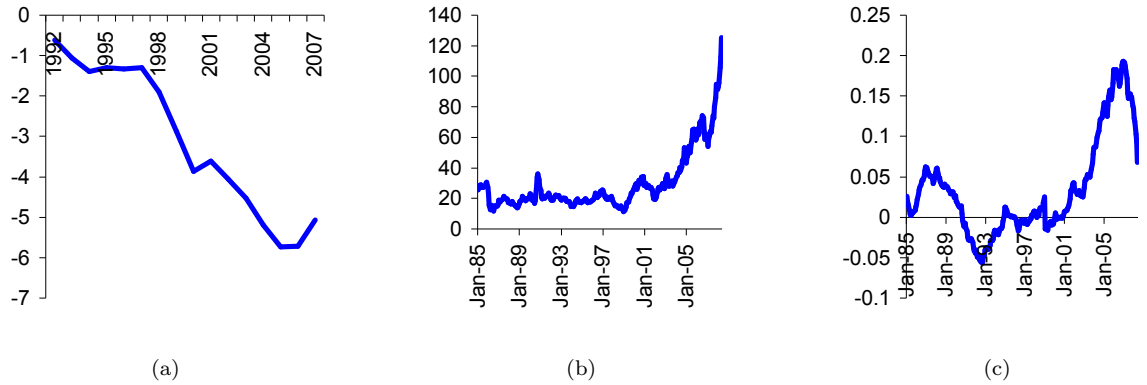
Table 1: Top Five Countries with which the US Trades, for the Month of November 2008

Country name	Total in billions of US dollar	Total in billions of US dollar, Year to date
Canada	41.77	559.86
China	33.5	378.95
Mexico	27.45	343.04
Japan	15.16	191.46
Federal Republic of Germany	11.54	141.27

Table 2: US Crude Oil Imports Top Five Countries (Thousand Barrels per Day)

Country	08-Oct	08-Sep	YTD 2008	07-Oct	YTD 2007
Canada	2,055	1,923	1,910	1,898	1,894
Saudi Arabia	1,427	1,429	1,519	1,370	1,416
Mexico	1,254	890	1,180	1,322	1,419
Venezuela	1,014	944	1,037	1,221	1,131
Nigeria	908	508	941	1,184	1,055

Figure 2: Continuing US' Trade Imbalance with Canada (Millions of US Dollar) and US Dollars to One Canadian Dollar
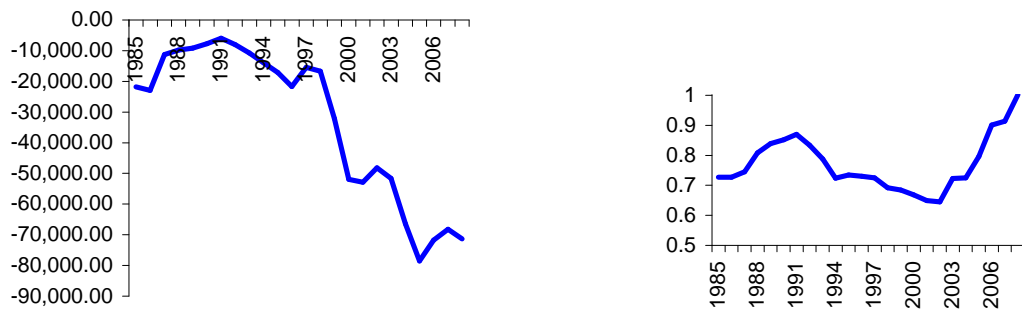
Figure 3: Plots of Variables of Interest in Level (Grey) and in First Difference (Blue)

(*RTB*)

(*RER*)

(*ROP*)

(*USRHPI*)

(*CARHPI*)

(*USRY*)

(*CARY*)

58
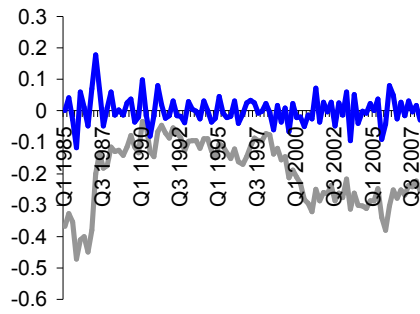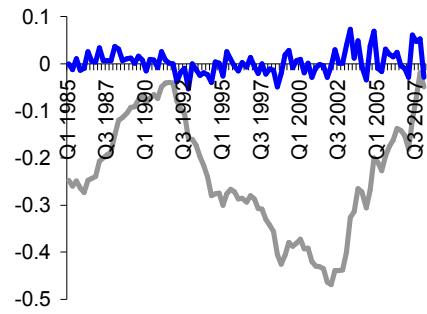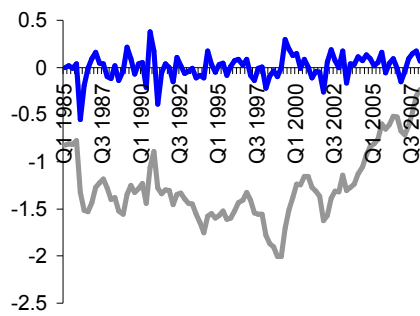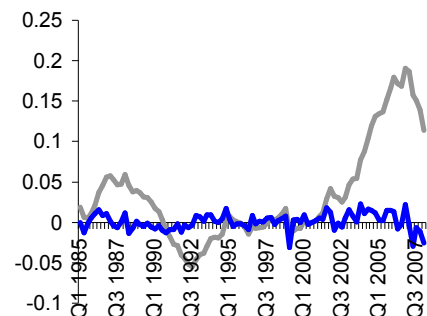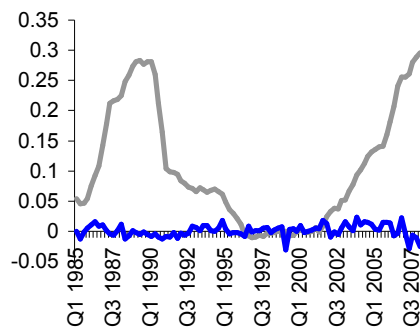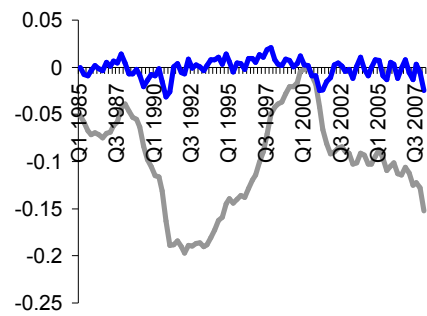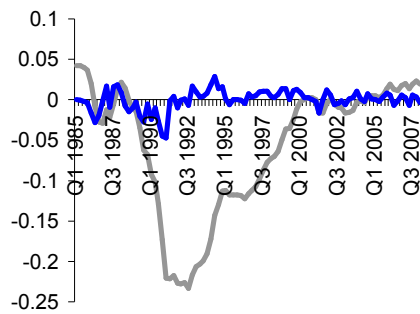
Table 3: Results from Augmented Dickey-Fuller (ADF) Test

Variables	Prob.	Variables	Prob.
RTB	0.0539	ΔRTB	0.0052*
RER	0.8004	ΔRER	0.0000*
ROP	0.8078	ΔROP	0.0000*
$USRHPI$	0.3962	$\Delta USRHPI$	0.0535
$CARHPI$	0.5590	$\Delta CARHPI$	0.0064*
$USRY$	0.2682	$\Delta USRY$	0.0782
$CARY$	0.4556	$\Delta CARY$	0.0002*

Note: * indicates statistical significance at 5 percent significance level.

Table 4: Results from Johansen Cointegration Analysis

No. hypothesized CEs	Trace statistic	Max-eigen statistic	Trace 5 percent critical value	Max-eigen 5 percent critical value	Trace prob.	Max-eigen prob.
None	160.8873	53.61057	125.6154	46.23142	0.0001*	0.0069*
At most 1	107.2767	38.84848	95.75366	40.07757	0.0064*	0.0683
At most 2	68.4282	28.33729	69.81889	33.87687	0.0642	0.1984
At most 3	40.09092	17.78209	47.85613	27.58434	0.2193	0.5133
At most 4	22.30883	14.59297	29.79707	21.13162	0.2816	0.3185
At most 5	7.715855	6.489533	15.49471	14.2646	0.4963	0.5512
At most 6	1.226321	1.226321	3.841466	3.841466	0.2681	0.2681

Note: * indicates statistical significance at 5 percent significance level.

Table 5: The Estimated Cointegrating Vector (β) on US Real Trade Balance with Canada

				Cointegrating vector, $\hat{\beta}$				
	RTB	*RER*	*ROP*	*USRHPI*	*CARHPI*	*USRY*	*CARY*	*t*
$\hat{\beta}$	1.000	-0.448*	0.203*	0.700*	-0.591*	-0.197	-0.011	-0.002
(s.e.)	(-)	(0.160)	(0.033)	(0.335)	(0.147)	(0.229)	(0.312)	(0.000)

Table 6: Estimates from the Vector Error-Correction Model for One Cointegrating Vector

Dependent variable	Adjustment coefficient, $\hat{\alpha}$	Constant term	Lagged variables in first difference dynamic terms			
ΔRTB	-0.343*	0.002				
(s.e.)	(0.120)	(0.005)				
			ΔRTB_{t-1}	ΔRER_{t-1}	ΔROP_{t-1}	
			0.128	-0.452**	0.056	
			(0.142)	(0.241)	(0.038)	
			$\Delta USRHPI_{t-1}$	$\Delta CARHPI_{t-1}$	$\Delta USRY_{t-1}$	$\Delta CARY_{t-1}$
			-0.169	-0.319	0.58	0.046
			(0.630)	(0.522)	(0.697)	(0.598)

Note: The value of R^2 is 0.167. * (**) indicate statistical significance at 5 (10) percent significance level.

Figure 4: The Response of US Real Trade Balance with Canada to Shocks
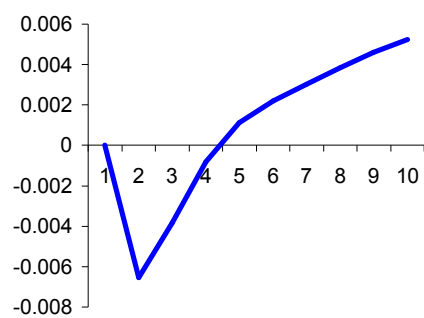
(*RER* Shocks, Not Accumulated)
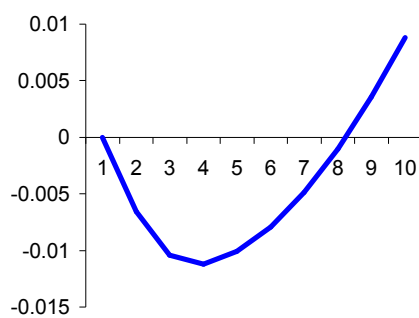
(*RER* Shocks, Accumulated)
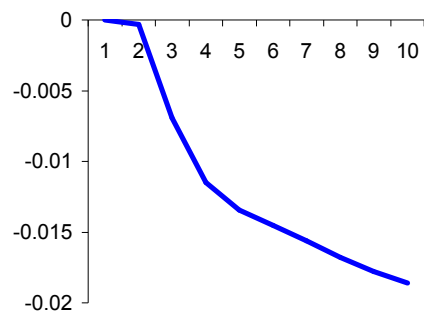
(*ROP* Shocks, Not Accumulated)
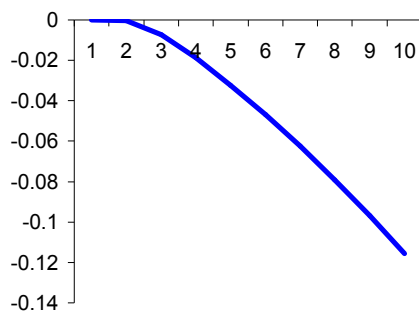
(*ROP* Shocks, Accumulated)

Table 7: Results from Both AIC and SC Criteria

Specification	AIC Criterion	SC Criterion
MSIA(2)-VECM(1)	-36.1927	-31.5138
MSIAH(2)-VECM(1)	-37.1867	-31.6679

Table 8: Estimates from the Non-Linear Vector Error-Correction Model, MSIA(2)-VECM(1)

Dependent variable	Estimates			
ΔRTB				
(s.e.)				
Regime 1	Adjustment coefficient, $\hat{\alpha}$	Intercept	σ	
	-0.369	-0.015**	0.036	
	(0.205)	(0.008)		
	ΔRTB_{t-1}	ΔRER_{t-1}	ΔROP_{t-1}	
	-0.022	-0.456	-0.001	
	(0.199)	(0.368)	(0.049)	
	$\Delta USRHPI_{t-1}$	$\Delta CARHPI_{t-1}$	$\Delta USRY_{t-1}$	$\Delta CARY_{t-1}$
	-0.971	-0.503	-1.934*	1.027
	(1.038)	(0.668)	(0.923)	(0.938)
Regime 2	Adjustment coefficient, $\hat{\alpha}$	Intercept	σ	
	-0.379*	-0.001	0.036	
	(0.134)	(0.007)		
	ΔRTB_{t-1}	ΔRER_{t-1}	ΔROP_{t-1}	
	0.109	-0.540*	0.022	
	(0.193)	(0.250)	(0.051)	
	$\Delta USRHPI_{t-1}$	$\Delta CARHPI_{t-1}$	$\Delta USRY_{t-1}$	$\Delta CARY_{t-1}$
	0.057	0.324	2.644*	-0.454
	(0.655)	(0.779)	(0.934)	(0.669)
	Non-Linear	Linear		
AIC Criterion	-36.1927	-35.8401		
HQ Criterion	-34.3182	-34.7467		
SC Criterion	-31.5138	-33.1108		
Log-Likelihood	1585.6129	1506.6853		
LR Linearity Test	157.8552			
DAVIES	0.0000*			

Note: * (**) indicates statistical significance at 5 (10) percent significance level.

Table 9: Estimates from the Non-Linear Vector Error-Correction Model, MSIAH(2)-VECM(1)

Dependent variable		Estimates		
ΔRTB (s.e.)				
Regime 1	Adjustment coefficient, $\hat{\alpha}$	Intercept	σ	
	-0.357 (0.243)	-0.009 (0.010)	0.042	
	ΔRTB_{t-1}	ΔRER_{t-1}	ΔROP_{t-1}	
	-0.039 (0.232)	-0.284 (0.508)	-0.005 (0.059)	
	$\Delta USRHPI_{t-1}$	$\Delta CARHPI_{t-1}$	$\Delta USRY_{t-1}$	$\Delta CARY_{t-1}$
	-0.764 (1.254)	-0.620 (0.762)	-1.473 (1.100)	1.050 (1.100)
Regime 2	Adjustment coefficient, $\hat{\alpha}$	Intercept	σ	
	-0.407* (0.127)	-0.003 (0.006)	0.035	
	ΔRTB_{t-1}	ΔRER_{t-1}	ΔROP_{t-1}	
	0.180 (0.179)	-0.548* (0.227)	0.030 (0.048)	
	$\Delta USRHPI_{t-1}$	$\Delta CARHPI_{t-1}$	$\Delta USRY_{t-1}$	$\Delta CARY_{t-1}$
	0.025 (0.605)	0.091 (0.736)	2.401* (0.868)	-0.191 (0.618)

	Non-Linear	Linear
AIC Criterion	-37.1867	-35.8401
HQ Criterion	-34.9757	-34.7467
SC Criterion	-31.6679	-33.1108
Log-Likelihood	1652.8727	1506.6853
LR Linearity Test	292.3749	
DAVIES	0.0000*	

Note: * (**) indicates statistical significance at 5 (10) percent significance level.

Table 10: Transition Probabilities and Regime Properties

	Regime 1	Regime 2	No. of obs.	Probability	Duration
MSIA(2)-VECM(1)					
Regime 1	0.7656	0.2344	35.9	0.4402	4.27
Regime 2	0.1843	0.8157	43.1	0.5598	5.42
MSIAH(2)-VECM(1)					
Regime 1	0.5305	0.4695	30.0	0.3771	2.13
Regime 2	0.2842	0.7158	49.0	0.6229	3.52

Figure 5: Smoothed and Filtered Probabilities of Non-Linear Model for US Real Trade Balance with Canada, MSIA(2)-VECM(1)
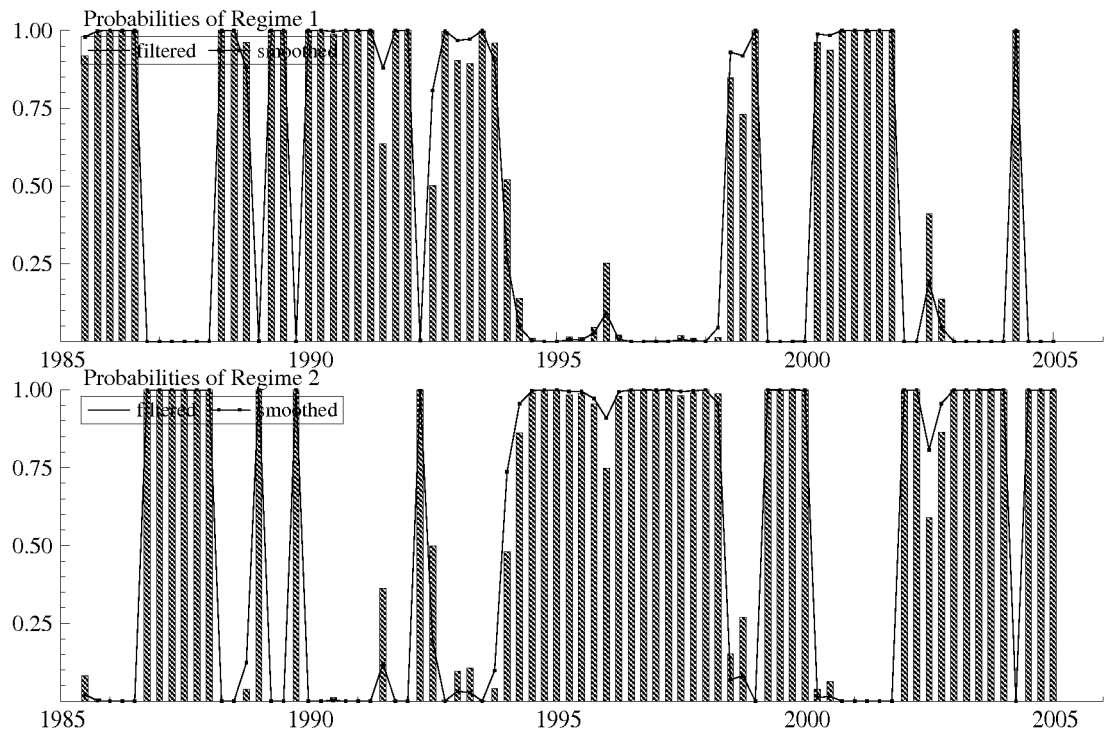
Figure 6: Smoothed and Filtered Probabilities of Non-Linear Model for US Real Trade Balance with Canada, MSIAH(2)-VECM(1)
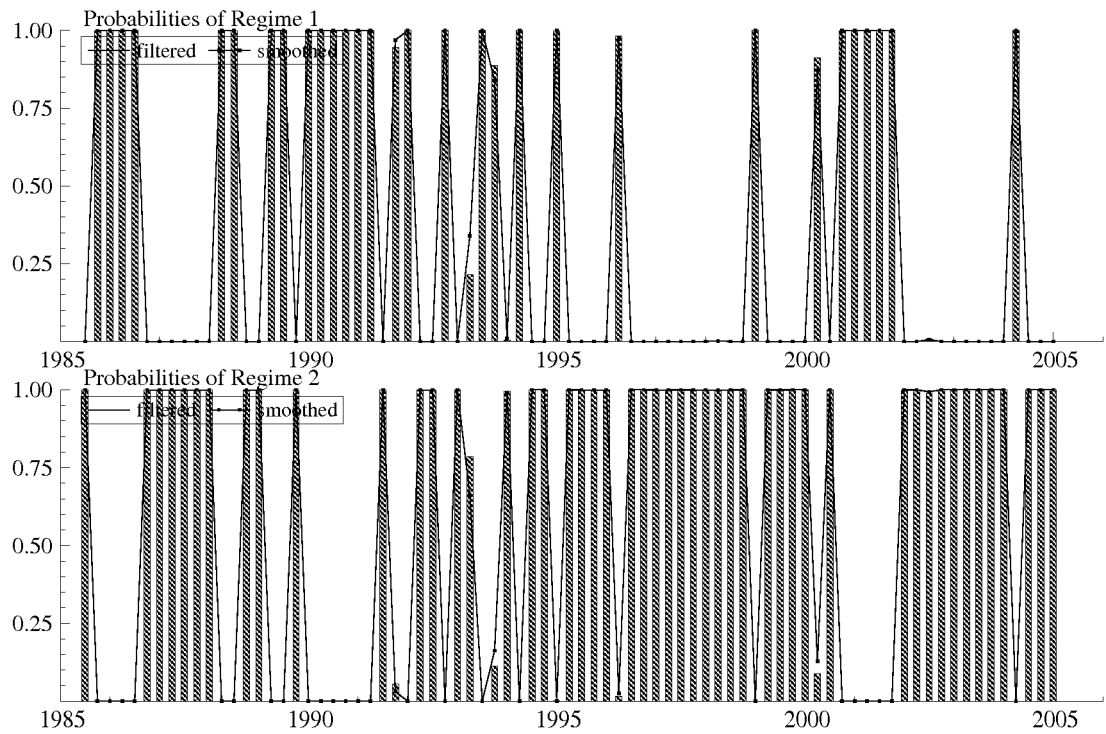
Table 11: US Trade Balance with Canada, US Business Cycle and Variables of Interest

Period and regime (roughly)	US trade balance with Canada	NBER business cycle	Canadian dollar	Oil price	House price
Q1 94 - Q1 00 (regime 2)	Worsening	Expansion (Mar 91 - Mar 01)	Weak*	Stable	Stable
Q2 00 - Q4 01 (regime 1)	Improving	Contraction (Mar 01 - Nov 01)	Weak*		
Q1 02 - Q1 05 (regime 2)	Worsening	Expansion (Nov 01 - Dec 07)	Strong	Rising*	Rising*

Note: * shows negative effect on US trade balance with Canada.

Table 12: The Forecast Results for US Real Trade Balance with Canada

Model	RMSE	Error Ratio
Linear VECM	0.051	1.186
Linear DVAR	0.043	1.000
Non-linear VAR	0.042	0.977
Random Walk	0.043	

Figure 7: Forecasts from Non-Linear VAR Model (Dashed) and Actual US Real Trade Balance with Canada (Solid), Quarter Two 2005 to Quarter One 2008
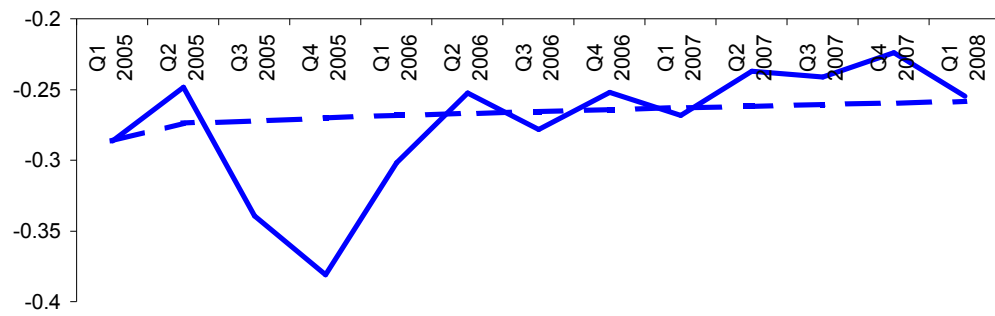
Table 13: The Forecast Results for US Real Trade Balance with Canada, Augmented and Baseline Model

Model	Error Ratio
Linear augmented VAR	0.978
Non-linear augmented VAR	0.992

References

Amano, R. A. and van Norden, S. (1998). Oil Prices and the Rise and Fall of the US Real Exchange Rate. *Journal of International Money and Finance*, 17(2):299–316.

Backus, D. (1998). The Japanese Trade Balance: Recent History and Future Prospects. *Japan and the World Economy*, 10(4):409–420.

Bahmani-Oskooee, M. and Artatrana, R. (2004a). Dynamics of the US Trade with Developing Countries. *Journal of Developing Areas*, 37(2):1–11.

Bahmani-Oskooee, M. and Artatrana, R. (2004b). The J-Curve: a literature review. *Applied Economics*, 36(13):1377–1398.

Bahmani-Oskooee, M. and Tatchawan, K. (2001). Bilateral J-Curve Between Thailand and her Trading Partners. *Journal of Economic Development*, 26(2):107–117.

Brada, J. C., Kutan, A. M., and Zhou, S. (1997). The Exchange Rate and the Balance of Trade: The Turkish Experience. *Journal of Development Studies*, 33(5):675–692.

Bryant, R. C., Hooper, P., and Mann, C. L. (1993). *Evaluating Policy Regimes*. The Brookings Institution, Washington D.C.

Carroll, C. D., Otsuka, M., and Slacalek, J. (2006). How Large Is the Housing Wealth Effect? A New Approach. *National Bureau of Economic Research Working Paper Series*, (12746).

Clements, M. P. and Hendry, D. F. (1995). Forecasting in Cointegrated Systems. *Journal of Applied Econometrics*, 10(2):127–146.

Clements, M. P. and Hendry, D. F. (2002). Modelling Methodology and Forecast Failure. *The Econometrics Journal*, 5(2):319–344.

Clements, M. P. and Hendry, D. F. (2006). Forecasting with breaks. volume 1 of *Handbook of Economic Forecasting*, pages 605–657. Elsevier, Amsterdam.

Clements, M. P. and Krolzig, H.-M. (2002). Can oil shocks explain asymmetries in the US Business Cycle. *Empirical Economics*, 27(2):185–204.

Clements, M. P. and Krolzig, H.-M. (2004). Can regime-switching models reproduce the business cycle features of US aggregate consumption, investment and output. *International Journal of Finance and Economics*, 9(1):1–14.

Clements, M. P. and Smith, J. (1999). A Monte Carlo Study of the Forecasting Performance of Empirical SETAR Models. *Journal of Applied Econometrics*, 14(2):123–141.

Davies, R. B. (1977). Hypothesis Testing When a Nuisance Parameter is Present Only Under the Alternative. *Biometrika*, 64(2):247–254.

De Gooijer, J. G. and Kumar, K. (1992). Some Recent Developments in Non-Linear Time Series Modelling, Testing and Forecasting. *International Journal of Forecasting*, 8(2):135–156.

Diebold, F. X. and Mariano, R. S. (1995). Comparing Predictive Accuracy. *Journal of Business and Economic Statistics*, 13(3):253–263.

Engle, R. F. and Granger, C. W. J. (1987). Co-Integration and Error Correction: Representation, Estimation, and Testing. *Econometrica*, 55(2):251–276.

Engle, R. F. and Yoo, B. S. (1987). Forecasting and testing in co-integrated systems. *Journal of Econometrics*, 35(1):143–159.

Ericsson, N. R. and Marquez, J. (1993). Encompassing the Forecasts of US Trade Balance Models. *The Review of Economics and Statistics*, 75(1):19–31.

Fergusson, I. F. (2008). United States-Canada Trade and Economic Relationship: Prospects and Challenges. *CRS Report for Congress*.

Fratzscher, M., Juvenal, L., and Sarno, L. (2008). Asset Prices, Exchange Rates and the Current Account. *Research Division, Federal Reserve Bank of St. Louis Working Paper Series*, 31.

Garcia, R. (1998). Asymptotic Null Distribution of the Likelihood Ratio Test in Markov Switching Models. *International Economic Review*, 39(3):763–788.

Georgopoulos, G. (2008). The J-curve Revisited: An Empirical Analysis for Canada. *Atlantic Economic Journal*, 36(3):315–332.

Hansen, B. E. (1992). The Likelihood Ratio Test under Nonstandard Conditions: Testing the Markov Switching Model of GNP. *Journal of Applied Econometrics*, 7(S):S61–S82.

Hoffman, D. L. and Rasche, R. H. (1996). Assessing Forecast Performance in a Cointegrated System. *Journal of Applied Econometrics*, 11(5):495–517.

Hsing, H.-M. (2005). Re-examination of J-curve Effect for Japan, Korea and Taiwan. *Japan and the World Economy*, 17(1):43–58.

Issa, R., Lafrance, R., and Murray, J. (2008). The Turning Black Tide: Energy Prices and the Canadian Dollar. *Canadian Journal of Economics*, 41(3):737–759.

Johansen, S. (1988). Statistical analysis of cointegration vectors. *Journal of Economic Dynamics and Control*, 12(2-3):231–254.

Johansen, S. and Juselius, K. (1990). Maximum Likelihood Estimation and Inference on Cointegration-With Application to the Demand for Money. *Oxford Bulletin of Economics and Statistics*, 52(2):169–210.

Kilian, L., Rebucci, A., and Spatafora, N. (2007). Oil Shocks and External Balances. *CEPR Discussion Papers*, 6303.

Krolzig, H.-M. (1997). *Markov-Switching Vector Autoregressions, Modelling, Statistical Inference and Application to Business Cycle Analysis*. Springer, Berlin.

Krolzig, H.-M. (1998). Econometric Modelling of Markov-Switching Vector Autoregressions using MSVAR for Ox. .

Krolzig, H.-M. (1999). Statistical Analysis of Cointegrated VAR Processes with Markovian Regime Shifts. *Computing in Economics and Finance, Society for Computational Economics*, (1113).

Krolzig, H.-M. (2005). MSVAR, MSVAR1.32a for Ox 3.4. .

Lin, J.-L. and Tsay, R. S. (1996). Co-integration Constraint and Forecasting: An Empirical Examination. *Journal of Applied Econometrics*, 11(5):519–538.

MacKinnon, J. G. (1996). Numerical Distribution Functions for Unit Root and Cointegration Tests. *Journal of Applied Econometrics*, 11(6):601–618.

MacKinnon, J. G., Haug, A. A., and Michelis, L. (1999). Numerical Distribution Functions of Likelihood Ratio Tests for Cointegration. *Journal of Applied Econometrics*, 14(5):563–77.

Marwah, K. and Klein, L. R. (1996). Estimation of J-Curves: United States and Canada. *The Canadian Journal of Economics*, 29(3):523–539.

Moura, G. and Silva, S. D. (2005). Is There a Brazilian J-Curve? *Economics Bulletin*, 6(10):1–17.

Phillips, P. C. B. and Perron, P. (1988). Testing for a Unit Root in Time Series Regression. *Biometrika*, 75(2):335–346.

Rebucci, A. and Spatafora, N. (2006). Oil Prices and Global Imbalances. *IMF World Economic Outlook, April*, pages 71–96.

Ryden, T. (1995). Estimating the Order of Hidden Markov Models. *Statistics*, 26(4):345–354.

Stock, J. H. and Watson, M. W. (1999). A comparison of linear and nonlinear univariate models for forecasting macroeconomic time series. Cointegration, Causality and Forecasting: A Festschrift in Honour of Clive Granger, pages 1–44. Oxford University Press, Oxford.

Yousefi, A. and Wirjanto, T. S. (2003). Exchange rate of the US dollar and the J curve: the case of oil exporting countries. *Energy Economics*, 25(6):741–765.

Chapter 2: Monetary Policy, Exchange Rates and Asian Stock Markets

Tim Leelahaphan[*]

Abstract

We examine the effect of monetary policy and exchange rate on stock price movements in Asia. We employ a Bayesian structural vector autoregression model and impose sign restrictions to identify simultaneously and uniquely contractionary monetary policy shocks and exchange rate depreciation shocks in an integrated framework. This study covers the stock markets of Thailand, Malaysia and South Korea, over the period 1989-2008. Our main results acquired using sign restrictions show that monetary policy shocks result in a strongly persistent effect on market index real stock prices whereas the impact of exchange rate shocks is short-lived over the short run. The variance decomposition suggests that the exchange rate is as important as monetary policy for explaining the dynamics of market and financial sector index real stock prices. More precisely, for all the countries examined, real exchange rate developments have been more important in the short run. Based purely on our findings, two conclusions are made. Firstly, because of the mistimed and/or persistent effect of monetary policy on both the real economy and financial markets, we argue that one needs to be cautious in using monetary policy to constrain asset price misalignment. Secondly, due to the evidence that exchange rates principally have a contemporaneous impact on equity prices, we suggest that, in the short run, such incorrectly aligned asset prices might potentially be corrected by focusing on exchange rate movements.

JEL Classification: C32, E52, G12
Keywords: Asia, Bayesian structural vector autoregression model, exchange rate, monetary policy, sign restrictions, stock prices

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1 Introduction

While it is widely believed that the interest rate is an important determinant of stock prices, less attention has been paid to the relationship between the exchange rate and the stock market. Nevertheless, since the 1997 Asian financial crisis, the exchange rate and stock price relationship has received greater attention. This might be due to the fact that, in the aftermath of the crisis, the affected countries suffered turmoil in both foreign exchange and equity markets. In Thailand, the Thai baht reached its lowest point in January 1998 and the stock market fell by 75 percent. In Malaysia, over the period from July 1 to September 30, 1997, the ringgit plunged by 37.4 percent and the stock market fell by 31.4 percent. In South Korea, by the end of the same year, the Korean won dropped dramatically by more than 150 percent and its stock market plunged by more than 50 percent.[1] If the exchange rate explains a large amount of the dynamics of market and financial sector index real stock prices, it is reasonable to believe that stock market crises and asset price bubbles could potentially be prevented by focusing on the movement of exchange rates. In addition, market participants could use information on exchange rates to predict the stock market behavior.

The link between monetary policy and the stock market is generally realized as the appropriate influence of such policy on the decision-making of the private sector. This is in order to fulfill some objectives. To have low and stable inflation

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[1]See Baharumshah *et al.* (2002).

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and output near its natural rate are generally believed to be the main objectives of the central bank, achieved by setting and exerting control over the (real) interest rates and by appropriately monitoring the decisions of the private sector. In the framework of the new Keynesian theory, in which prices are not fully flexible in the short run, the real interest rate could be temporarily influenced by the policy of the monetary authority. Consequently, this would affect the real output in addition to nominal prices.

Studying Asian stock markets and economies affected by underlying structural shocks is motivated by the lesson we learn from the Asian crisis of 1997-1998. The severe consequences of the crisis economically and politically destroyed many of the regional economies. It is argued that the primary reason for the Asian financial crisis is attributable to an inappropriate mixture of policies. (See Rajan, Thangavelu and Parinduri (2008)). In particular, this is due to the fact that *regional economies attempted to maintain simultaneously fairly rigid exchange rates (soft US dollar pegs) and monetary policy autonomy in the presence of large-scale capital outflows.* Even after more than a decade has passed since the Asian financial crisis, exchange rate issues and monetary policies relevant to Asia, especially those relating to financial issues and asset prices, are still in the focus of economists and market participants. Specifically, the issue of whether monetary policy should actively seek to promote asset price stability might be the most important question central bankers are currently facing.

There are two main objectives of this chapter. Firstly, we would like to

examine if there is difference in the influence of the monetary policy actions and of exchange rate developments on the stock market. In particular, the systematic feature, in terms of the persistence of the impact, of such an influence of these two underlying structural shocks is deliberately considered. Secondly, in addition to monetary policy commonly believed to be important determinant of stock prices, we would like to assess quantitatively if the exchange rate has also played an important role in driving the stock market. In particular, we examine and compare the extent to which monetary policy and exchange rate are responsible for the movements in Asian stock prices.

We are aware that monetary policy shocks and exchange rate shocks are definitely not able to capture sufficiently the full movements of the stock market in Asian countries. That is, other shocks, oil price shocks and fiscal policy shocks, for example, have also certainly played a role in the dynamics of the stock market. Nevertheless, we reasonably choose to examine only these two underlying structural shocks. This is due to the fact that they are of interest to the monetary authority setting the official interest rate, which is used to manage both inflation and output, and managing the country's exchange rate.[2]

In order to differentiate monetary policy shocks and exchange rate shocks in the data, we employ a Bayesian structural vector autoregression model and impose sign restrictions to identify simultaneously and uniquely contractionary

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[2]This is in line with the literature. For example, Uhlig (2005) does not aim at an entire decomposition of the one-step ahead prediction error into all the components caused by underlying structural shocks. Rather, the study concentrates on identifying only one underlying structural shock.

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monetary policy and exchange rate depreciation shocks in an integrated frame-work. In particular, the approach we pursue stems from the methodology developed by Canova and de Nicolo (2002), Uhlig (2005) and Mountford and Uhlig (2008). The identification of these two underlying structural shocks is neat. This is due to the fact that it is realized by verifying whether the signs of the corresponding impulse responses are accepted by priori consensual considerations regarding the effects of monetary policy shocks and exchange rate shocks on key macroeconomic variables.

The sign restrictions approach requires only a minimal set of economically meaningful restrictions to identify contractionary monetary policy shocks and exchange rate depreciation shocks. The impact of these two structural shocks on a number of variables of interest, market and financial sector index real stock prices in this chapter, is left unrestricted. In addition, the significance testing, sign and persistence of the impact of these two underlying structural shocks on the unrestricted variables of interest are achieved by this approach.

Several advantages of the sign restrictions approach are worth mentioning. Firstly, compared with the traditional structural vector autoregression model, the sign restrictions approach makes restrictions that are often used implicitly and are in line with the conventional considerations more explicitly. Secondly, in a methodological context, the results from this approach are not dependent on the chosen decomposition of a variance-covariance matrix. Consequently, such results are not altered by reordering the variables and by a consequent selection

of a different Cholesky decomposition.

This study covers the stock markets of Thailand, Malaysia and South Korea, over the period 1989-2008. Often referred to as semi- or newly industrialized economies and considered to be emerging markets with respect to the maturity of financial markets, the three Asian economies examined were affected by the 1997 Asian financial crisis. Thailand and South Korea were the countries most hurt by the slump, while Malaysia was fairly affected by the crisis. In addition, all of these three Asian countries saw their currencies fall significantly relative to the United States dollar, though extended currency losses were realized by the harder hit economies. Regarding capital market liberalization and capital control, these three Asian countries do not completely open their equity markets to foreign investors.

In order to reach findings, we adapt standard VAR analyses to deal with these single shocks based upon sign restrictions. In particular, for each country examined, impulse responses, variance decomposition and historical decomposition are carefully considered.

With impulse responses, we observe the following. For the countries examined, we can confidently conclude that contractionary monetary policy shocks have a statistically significant effect on the market index real stock prices in general. In addition, the findings relating to real stock prices and exchange rate depreciation shocks prove to be statistically significant within a year immediately following the shocks for all three countries examined. Nevertheless,

the significance of the impact is sensitive to the choice of horizons for sign restrictions to be imposed. In addition, financial sector index real stock prices react similarly to market index real stock prices, but with a greater magnitude. Our main results acquired using sign restrictions show that monetary policy shocks result in a strongly persistent effect on market index real stock prices whereas the impact of exchange rate shocks is short-lived over the short run. The variance decomposition suggests that the exchange rate is as important as monetary policy for explaining the dynamics of market and financial sector index real stock prices. More precisely, for all the countries examined, real exchange rate developments have been more important in the short run.

This chapter briefly explains the link between monetary policy and the stock market and the relationship between the exchange rate and the stock market and provides a review of recent empirical literature in Section 2. Section 3 outlines the econometric approach to identifying the monetary policy shocks and the exchange rate shocks. The data set used for the estimations, our empirical results and an analysis pertaining to impulse responses and to variance decomposition are presented in Section 4. Section 5 and Section 6 provide issues regarding historical decomposition and robustness, respectively. Section 7 concludes.

2 Monetary Policy, Exchange Rates and Stock Markets

We divide this section into two subsections. These are (i) monetary policy and stock prices and (ii) exchange rates and stock prices.

2.1 Monetary Policy and Stock Prices

For equity prices, an approach for determining stock prices is commonly assumed to be forward-looking. This is in order to reflect the discounted expected future sum of return on the assets of the private sector. For monetary policy, the monetary authority affects both the current and the expected future real interest rate. This in turn has an effect on the decisions of households and of firms to consume and to invest, respectively, in terms of timing. To establish the relationship, following Bjørnland and Leitemo (2009), changes in the expected future interest rate serving as a discount rate are directly attributed to the link between monetary policy and the stock market. In addition, changes in all the factors affecting aggregate demand, the path of profit and the expected dividends, if firms are in a monopolistic competition and mark-up pricing market, are indirectly attributed to the relationship between monetary policy and the stock market.[3]

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[3]In addition to the determinants of expected future dividends, monetary policy could also have an influence on stock returns premium. This is by affecting the degree of uncertainty market participants face.

While stock prices are expected to decrease following an increase in interest rates, there have been several empirical studies providing evidence in favor of a positive rather than a negative relationship between these two variables of interest. A changing risk premium is one possible explanation. That is, a very low level of interest rates could be an outcome of increased risk and/or precautionary saving. In particular, the decrease in interest rates might be attributable to the fact that market participants move away from riskier assets (stocks for example) towards less risky ones (bonds or real estate for example).[4] In addition, such a positive relationship between interest rates and stock prices could be explained by the fact that changes in interest rates could carry information about some future fundamentals changes (dividends for example).[5]

As a review of recent empirical literature confirming strong effects of monetary policy shocks on the equity market, Lastrapes (1998) examines the response of stock prices to these shocks in industrialized countries, using solely long-run (neutrality) restrictions to identify monetary policy shocks. Specifically, it is assumed that interest rates, real output, real stock prices and real money balances do not permanently respond to money supply shocks. For most of the countries examined, it is found that real stock prices are significantly affected by unexpected changes in nominal money supply, with a varying magnitude of effects across different countries. Similar to Lastrapes (1998), Rapach (2001) also uses a long-run restriction to achieve the identification of money supply

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[4]See Barsky (1989).
[5]See Shiller and Beltratti (1992).

and other macro shocks. Conforming to the standard dividend discount model, the evidence shows that expansionary monetary policy shocks have a positive effect on real US stock prices.

Neri (2004) uses the structural vector autoregression methodology to examine the effects of monetary policy shocks on stock market indices in the G7 countries and Spain. The results show that contractionary monetary policy shocks have a negative effect on the equity market index. Nevertheless, such an effect is small and transitory. In addition, the persistence, magnitudes and timing of these effects of monetary policy are different across countries.

2.2 Exchange Rates and Stock Prices

A causal relationship running from exchange rates to stock prices is suggested in a vast amount of literature. Among others, Granger *et al.* (2000), in order to determine the appropriate Granger causality relationships between equity prices and exchange rates, apply unit root and cointegration models, using data from the Asian financial crisis. With respect to the impulse response functions reported, for South Korea, they find that exchange rates lead stock prices. This is in line with the traditional approach discussed above. For the Philippines, on the other hand, equity prices negatively lead exchange rates. This is expected under the portfolio approach. Strong feedback relationships are found for Hong Kong, Malaysia, Singapore, Taiwan and Thailand. For Indonesia and Japan, no recognizable pattern could be revealed.

In line with the goods market hypothesis and traditional approach, changes in the exchange rate could affect the competitiveness of multinational firms. Consequently, their earnings and stock prices are affected. On the one hand, domestic currency depreciation (appreciation) results in cheaper (more expensive) exporting goods and leads to an improvement (decrease) in competitiveness and foreign demand. The value of an exporting firm would then increase (decrease), following its domestic currency depreciation (appreciation). On the other hand, such a relationship is the opposite for importing firms. Furthermore, this also applies if lots of imported inputs are used in their production. Because of currency depreciation, their production costs rise and both sales and profits might decline. Hence, a fall in their stock prices could occur. In addition, the transaction exposure of a firm is affected by fluctuations in the exchange rate. In particular, movements of exchange rates have an effect on a firm's future payables (or receivables) that are denominated in foreign currency. That is, domestic currency appreciation would decrease the profits of an exporting firm and depreciation of the local currency would increase profits. Besides multinational firms, domestic firms with insignificant international activities could also be exposed to exchange rate risk. This is when fluctuations in currency have an effect on their prices of input and output and on demand for their goods. Equity prices could also be affected by movements of exchange rates inducing equity flows.

3 Econometric Methodology

We divide this section into three subsections. These are (i) the structural vector autoregression model, (ii) the sign restrictions approach and (iii) the identification and implementation of sign restrictions.

3.1 Structural Vector Autoregression Model

We consider a reduced-form VAR of order p

$$Y_t = B(L)Y_{t-1} + u_t, \tag{1}$$

where Y_t is an $n \times 1$ vector of endogenous variables, $B(L)$ is a matrix polynomial in the lag operator L, of order p, u_t is an $n \times 1$ vector of reduced-form residuals, with a variance-covariance matrix $E[u_t u_t'] = \sum$ and $t = 1, ...T$. A constant, a time trend and exogenous variables might also be allowed for in this reduced-form representation.

Ordinary least squares could be performed to estimate consistently the above reduced-form VAR of order p. Nevertheless, identifying a structural representation of the VAR is demanding. To accomplish the identification of the VAR, it is necessary to impose enough restrictions to decompose u_t and to obtain economically meaningful structural shocks. In particular, we require a matrix A such that $Av_t = u_t$. We rewrite the reduced form above as

$$A^{-1}Y_t = A^{-1}B(L)Y_{t-1} + A^{-1}u_t = A^{-1}B(L)Y_{t-1} + v_t \qquad (2)$$

where uncorrelated, orthogonal structural shocks are represented by the $n \times 1$ vector of structural innovations $v_t = A^{-1}u_t$. This has the identity matrix as a variance-covariance matrix. That is, $E[v_t v_t'] = I_n$. A mapping between the structural representation and the reduced-form VAR of order p is specified by the identifying matrix A. In addition, with this identifying matrix A, the contemporaneous impact of structural shocks on the n endogenous variables could also be computed. In particular, an impulse vector and the contemporaneous impact of the ith structural shock of one standard deviation in size on each of the n endogenous variables in the system is represented by the ith column of this identifying matrix A, a_i.

To attain the identifying matrix A, we need at least $\frac{n(n-1)}{2}$ identifying restrictions to be imposed on A to achieve a unique solution. This is in addition to the property we might use that $\sum = E[u_t u_t'] = AE[v_t v_t']A' = AA'$. In particular, this property is the only restriction on A so far before imposing those $\frac{n(n-1)}{2}$ identifying restrictions. Nevertheless, such a property is not sufficient to achieve a unique solution, identification, for the identifying matrix A.

Common identification methods are (i) the use of Cholesky decomposition to orthogonalize the reduced-form disturbances or a recursive ordering of endogenous variables that restricts A to be lower triangular, (ii) the imposition of

contemporaneous restrictions on the error terms, (iii) the imposition of restrictions on the long-run dynamics of the model and (iv) the decomposition into temporary and permanent components.[6]

3.2   Sign Restrictions Approach

In our empirical study, in order to achieve identification of the above VAR model, we pursue an alternative identification approach stemming from the methodology developed inter alia by Canova and de Nicolo (2002), Uhlig (2005) and Mountford and Uhlig (2008). In particular, we impose sign restrictions on the impulse responses of a set of variables. Essentially, underlying structural shocks could be identified by verifying whether the signs of the corresponding impulse responses are accepted by priori consensual considerations.

According to a number of properties outlined by Uhlig (2005) and Mountford and Uhlig (2008), if there exists an n-dimensional vector q of unit length such that $a_i = \tilde{A}q$, where $\tilde{A}\tilde{A}' = \sum$ and $\tilde{A}$ is any arbitrary decomposition of $\sum$ (Cholesky decomposition for example), then the impulse vectors $a_i \in \Re^n$ could be retrieved, even if the true matrix A is not identified.

Based on this property, Monte Carlo simulations are performed, that is, given the estimated reduced-form VAR of order p, drawing random vectors q of unit length, computing the associated a_i vectors, calculating the candidate impulse responses and verifying whether the signs of the corresponding impulse responses

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[6]For the fourth identification method, see Blanchard and Quah (1989).

are accepted by priori consensual considerations over a number of horizons k. If all of the impulse responses satisfy the sign restrictions, the draw is kept.[7]

According to Mountford and Uhlig (2008), this sign restrictions approach could be generally applied to achieve the identification of multiple s underlying structural shocks, where $s \leq n$. For our empirical study, we would like to identify two shocks, contractionary monetary policy shocks and exchange rate deprecia-tion shocks. In particular, we could retrieve an impulse matrix $\left[a^{(1)}, a^{(2)}\right]$. Due to the fact that the covariance between the underlying structural shocks $v_t^{(1)}$ and $v_t^{(2)}$ corresponding to impulse vectors $a^{(1)}$ and $a^{(2)}$ is zero by construction, we could impose economically meaningful sign restrictions on the impulse re-sponses in order to characterize such an impulse matrix. Again, sign restrictions are required in addition to restrictions that provide orthogonality of these two underlying structural shocks.

Formally, instead of drawing a single n-dimensional vector q of unit length, we draw an $n \times s$ matrix $Q = \left[q^{(1)}, ..., q^{(s)}\right]$. This matrix contains s orthonormal vectors $q^{(s)}$, i.e. orthogonal vectors of unit length $(QQ' = I_s)$. This allows us to calculate the associated matrix $\tilde{A}Q = \left[a^{(1)}, ..., a^{(s)}\right]$. This also has a dimension $n \times s$ and contains s candidate impulse vectors.

Practically, using the relevant sign restrictions, we identify $a^{(1)} = \tilde{A}q^{(1)}$ and $a^{(2)} = \tilde{A}q^{(2)}$ to identify an impulse matrix $\left[a^{(1)}, a^{(2)}\right]$. Orthogonality conditions are jointly imposed. A joint draw from the posterior of the normal-Wishart

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[7]Provided with a structural impulse vector a_i, the matrix of impulse responses at horizon m, r_m, could be computed. That is, $r_m = [I - B(L)]^{-1} a_i$.

family for $(B\left(L\right),\sum)$ as well as a draw from the unit sphere are taken to attain candidate q vectors. This is described in detail in Uhlig (2005). Furthermore, Cholesky decomposition is calculated using the draw from the posterior.[8] Consequently, impulse responses for the impulse vector a could be computed.[9] If all of the impulse responses satisfy all the sign restrictions imposed for each of the s shocks, the joint draw is kept in such cases. If the sign restrictions are not met, each of such q draws is discarded.

3.3 Identification and Implementation of Sign Restrictions

The vector autoregression model comprises monthly data of (i) the log of the real industrial production index, (ii) the consumer price inflation, (iii) the interest rate, (iv) the log of the real exchange rates, (v) the log of the financial sector real stock prices index and (vi) the log of the real stock prices index. The industrial production index and stock prices are deflated by the consumer price index, so that they are measured in real terms. All the variables are in logarithms, except the consumer price inflation and the interest rate. The variables are specified in levels, which is consistent with most other related studies on sign restrictions

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[8]Note that, in the sign restrictions approach, the Cholesky decomposition calculated is not used for the purpose of identifying underlying structural shocks. In fact, it is only used as a useful computational tool. Mountford and Uhlig (2008) provide formal proof that similar results are achieved using any other factorization.

[9]See Mountford and Uhlig (2008), Appendix A. For the impulse vector $a^{(s)}$, the n-dimensional impulse response at horizon m, r_{am}, is computed as $r_{am} = \sum_{i=1}^{n} q_i r_{im}$ where r_{im} is the impulse response to the i-th column of $\tilde{A}$ at the same horizon and q_i is the i-th entry of $q = q^{(s)}$.

(see Uhlig (2005), Bracke and Fidora (2008) and Granville and Mallick (2009)).[10]

Formally, we estimate a VAR in the following endogenous variables

$$y_t = [GDP_t \ \ P_t \ \ i_t \ \ e_t \ \ f_t \ \ s_t] \tag{3}$$

where GDP represents the real GDP in log levels, P is the consumer price inflation, i corresponds to interest rates, e is the real exchange rates in log levels and f and s denotes financial sector and market index real stock prices in log levels, respectively.

The reduced-form VAR model is written as

$$Y_t = \Gamma_0 + \sum_{i=1}^{n} B_i Y_{t-i} + u_t, \tag{4}$$

where Y_t is a 6x1 vector of variables, that is, $Y_t = [GDP, P, i, e, f, s]'$. B_i is a coefficient matrix of size 6x6 and u_t is the one-step ahead prediction error with a variance-covariance matrix $\sum$. Γ_0 is the intercept.

This specification includes a set of variables that would allow us to identify simultaneously and uniquely monetary policy shocks and exchange rate shocks. We adopt the scheme of sign restrictions demonstrated in Table 1. Specifically, we assume a contractionary monetary policy shock to result in an increase in

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[10]Estimating the vector autoregression model which has the variables in its level specification is the now common practice. In particular, differencing to transform the model to the form that is stationary might not be necessary or appropriate. See Sims *et al.* (1990). In addition, specifying the variables in levels is also due to the fact that the sign restrictions approach is robust to nonstationarity. See Granville and Mallick (2009).

Table 1: Identifying Sign Restrictions

Shocks	GDP	P	i	e	f	s
Contractionary MP Shocks	-	-	+	app		
ER Depreciation Shocks	+	+	+	dep		

interest rates and a decrease in consumer price inflation, the real GDP and the real exchange rates (appreciation). Furthermore, a positive shock to the real exchange rates, i.e. a shock causing a depreciation in the real exchange rates, similarly increases the interest rates but also has a positive effect on the consumer price inflation and real GDP (in addition to the positive effect on the real exchange rates).

With the aim to identify uniquely and estimate the impact of (i) monetary policy shocks and (ii) exchange rate shocks on (i) financial sector index stock prices and (ii) market index stock prices, our identification scheme meets the following three purposes.

Firstly, the restrictions are in line with standard macroeconomic theory and have also been used in the empirical study using sign restrictions to identify these underlying structural shocks for individual economies. We argue that our sign restrictions on the dynamic responses to the underlying structural shocks examined are consistent with what would be suggested by dynamic stochastic general equilibrium (DSGE) models of both a new Keynesian model type and a real business cycle model type.

For contractionary monetary policy shocks, these are

1. A contractionary monetary policy shock does not lead to an increase (≤ 0) in consumer prices, following Uhlig (2005);

2. A contractionary monetary policy shock does not lead to an increase (≤ 0) in real GDP.

We are aware that these two restrictions could be controversial. For the non-positive change in consumer prices, there is some empirical evidence suggesting that inflation increases initially following contractionary monetary policy shocks. This is commonly referred to as the 'price puzzle'. See Eichenbaum (1992). For the non-positive change in real GDP, this has created research controversy (see, e.g., Uhlig (2005)). Nevertheless, the sign restrictions imposed here are the same as those in the empirical literature.

3. Following a negative relation between the domestic interest rate and the exchange rate, an increase in the domestic interest rate leads to a decrease (≤ 0) in the exchange rate - equivalently, to an appreciation of the domestic currency.

For exchange rate depreciation shocks, these are

1. Consumer prices (and domestic inflation) would not decrease (≥ 0) facing an exchange rate depreciation due to an increase in import prices;

2. The real GDP would not decrease (≥ 0) in response to exchange rate depreciation shocks. As exchange rates depreciate, imported goods become more expensive, while exported goods become less expensive, so demands for domestic goods increase, and outputs increase. Furthermore, we argue that, in the short run, changes in the price level are higher than changes in the cost of production.

Following the law of supply, the quantity of goods or services offered by suppliers temporarily increases accordingly.[11,12]

3. Interest rates would not decrease (≥ 0) in response to the exchange rate depreciation shocks. This is due to the fact that monetary policies would be exercised in such a way as to back up exchange rate depreciation. In addition, in the context of a monetary policy reaction function, such an increase in domestic short-term interest rates is also required due to an increase in consumer prices, import prices for example.

Secondly, the restrictions uniquely identify contractionary monetary policy and exchange rate depreciation shocks. The restrictions discriminate the two different shocks, in the sense that the set of sign restrictions imposed is mutually exclusive. The restrictions also aim to discriminate these two shocks from other potential underlying structural shocks to the economy (labor supply shocks, technology shocks or fiscal policy shocks).

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[11]We focus on the channel through which international trade affects the domestic economy. This is because the distinguishing feature of industrialization in Asian developing countries is that they have opted for an export-oriented strategy (see James *et al.* (1989)). Nevertheless, we are aware that the positive sign restriction on real GDP for exchange rate depreciation shocks could be arguable. That is, exchange rate depreciation shocks could have a negative effect on the real GDP. This is due to the fact that an increase in interest rates following exchange rate depreciation shocks might lower such a variable. As our first alternative identification scheme, we leave the impact of exchange rate depreciation shocks on the real GDP unrestricted. In addition, while the literature imposes a positive sign restriction on interest rates for exchange rate depreciation shocks (see Fratzscher *et al.* (2008) for example), as our second alternative identification scheme, we leave the impact of exchange rate shocks on such a variable unrestricted to identify solely an impact of the disturbance in the foreign exchange market. The results for these two alternative identification schemes show that our core findings are still attained. They are not reported to conserve space, but available upon request.

[12]Paustian (2007) shows that the imposed restrictions need to be sufficiently numerous for the sign restrictions to be able to define uniquely the unconstrained impulse responses. Using the identification scheme set out in Table 1, in some sense, follows such an argument.

Thirdly, we leave the impact on financial sector and market index real stock prices unrestricted. This would allow us to meet the main purpose of our analysis. That is, to assess the impact of contractionary monetary policy shocks and exchange rate depreciation shocks on these two variables.

Before providing the results and an analysis in the next section, two issues pertaining to (i) an interpretation of exchange rate shocks and (ii) an independent monetary policy under a fixed exchange rate regime should be clarified.

For the first issue, while the interpretation of monetary policy shocks is generally discussed in the literature, trying to give a precise structural interpretation of exchange rate shocks is challenging. Following Kim (2002), exchange rate shocks could be thought of as the disturbance in the foreign exchange market, altering the equilibrium of such a market and affecting the exchange rates. This could be due to abrupt shifts in portfolio preference (between domestic and foreign assets). Changes in the way market participants form expectations of exchange rates and evaluate the relative risks of domestic and foreign assets are also attributed to the disturbance in the foreign exchange market.

For the second issue, the standard Mundell-Fleming model predicts that, under a fixed exchange rate regime, monetary policy would be ineffective since capital flows induced by interest rate changes would completely offset the initial changes in money supply. Therefore, a fixed exchange rate regime is usually considered as one kind of monetary policy. This is because the economy is pegged on a 'nominal anchor' and the monetary authority loses its monetary autonomy

completely. However, Reisen (1993) suggests that many south-east Asian countries are successful in achieving the 'impossible trinity', namely, fixed exchange rates, independent monetary policy and free capital movements, mainly due to the weak interest rate mechanism in the domestic economy.

4 Data and Results

We divide this section into three subsections. These are (i) data, computation and specification, (ii) impulse responses analysis and (iii) variance decomposition analysis.

4.1 Data, Computation and Specification

Our data set is from February 1989 to November 2008 (238 observations). CPI and exchange rate series are taken from the IMF International Financial Statistics. US CPI is used in the definition of the real exchange rates. The industrial production index and interest rate series are also from the IMF International Financial Statistics for Malaysia and South Korea and from the Bank of Thailand for Thailand. Stock prices are taken from the Stock Exchange of Thailand, Kuala Lumpur Stock Exchange and Korea Stock Exchange for the Thailand SET index, Malaysia KLCI composite index and South Korea SE KOSPI 200 index, respectively; financial sector stock price series are also from the same sources, except for the Straits Times for Malaysia.

For computation, the procedure outlined in Section 3 is intensively repeated to identify simultaneously and uniquely contractionary monetary policy shocks and exchange rate depreciation shocks in an integrated framework and to generate estimation. In particular, for our results provided in the next subsection, we take joint draws from the posterior of B and $\sum$ of the VAR and draws of orthonormal matrices Q (from the unit sphere for each draw for the VAR) to identify simultaneously those two underlying structural shocks. If the range of impulse responses satisfies the sign restrictions, we keep the draw; otherwise we discard it. Computationally, we repeat this procedure until 1,000 draws compatible with the sign restrictions are acquired. Based on the draws kept, we calculate the median impulse responses and probability bands. The number of impulse response function steps to compute is 60.

In order to determine the lag selection, the Akaike information criterion (AIC) is used. Such a criterion generally selects two to six lags. In our empirical exercise, six lags of each variable are to be included in the model for all three countries examined in this chapter.[13] In addition, the VAR we estimate includes a constant. A horizon constrained for the sign restrictions to be imposed over, k, is five, which is conventional. In particular, the impulse responses are required to display the anticipated sign both contemporaneously and over the next five months. Note that this corresponds to a half-year horizon. In the Robustness

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[13]While selecting other lag lengths does not have a qualitative impact on the results, the impulse responses of both market and financial sector index real stock prices to the underlying structural shocks examined become less clear.

section, while we still select the same lag lengths and include a constant, we use different horizons for the sign restrictions to be imposed, $k = 2$ and $k = 8$.

We analyze the results for Thailand, Malaysia and South Korea. Firstly, we analyze the impulse responses of the real stock prices index and financial sector real stock prices index to contractionary monetary policy shocks and to exchange rate depreciation shocks. Secondly, we analyze variance decompositions, which are assessments of the importance of monetary policy and of exchange rate in explaining movements of the real stock prices index.

4.2 Impulse Responses Analysis

Figure 1 to Figure 3 show the impulse responses to contractionary monetary policy shocks and exchange rate depreciation shocks for Thailand, Malaysia and South Korea, respectively, using the pure-sign restrictions approach. Shocks are normalized to the magnitude of one standard deviation in size. For contractionary monetary policy shocks, the responses of the consumer price inflation, real GDP and real exchange rates are restricted not to be positive and the responses of interest rates are restricted not to be negative for six months. For exchange rate depreciation shocks, the responses of the consumer price inflation, real GDP, interest rates and the real exchange rates are restricted not to be negative for six months. The solid line is the median of the posterior distribution and the dashed lines represent the 16 percent and 84 percent quantiles of the posterior distribution of impulse responses, corresponding to one standard

deviation under the assumption of normality.

For Thailand, the figure indicates that the monetary policy effect on real stock prices has a marginally significant impact over the long run. This is strongly persistent. A one standard deviation contractionary monetary policy shock - roughly equal to a ten basis points increase in interest rates - results in a decrease of market index real stock prices.[14] This is statistically significant a month following the shocks. The course reversal is found at around twenty-five months. For exchange rate depreciation shocks, the results seem to prove that they are statistically significant and positive only in the first four months immediately following the shocks, before reversing the course.[15] For Malaysia, while real equity prices show a statistically significantly negative response to real exchange rate shocks over the short run, we are unconfident in concluding that contractionary monetary policy shocks have a significant effect on the market index real stock prices.[16] Nevertheless, it is indicative that contractionary monetary policy shocks tend to result in a fall in real equity prices. In addition, we observe that falling market index real stock prices reverse the course in around five months. Interestingly, this corresponds to the course reversal of

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[14]We observe Dornbusch's (1976) well-known exchange rate overshooting hypothesis stating that an increase in the interest rates should cause the exchange rates to appreciate instantaneously, then to depreciate in line with the uncovered interest parity (UIP) condition.

[15]In the section pertaining to robustness, using horizons for the sign restrictions to be imposed (k) equal to eight, we can conclude more confidently that exchange rate depreciation shocks have a statistically significant effect on the market index real stock prices over the short run, for Thailand.

[16]According to the results reported in the Robustness section using $k = 8$, we can conclude more confidently that contractionary monetary policy shocks have a statistically significant effect on the market index real stock prices for Malaysia.

real exchange rates from depreciation to appreciation. For South Korea, both contractionary monetary policy shocks and exchange rate depreciation shocks have a statistically significantly negative effect on the stock market. Similar to the results for Malaysia, exchange rate shocks have a marginally significant impact over the short run and course reversal is observed in around ten months. Monetary policy shocks result in a strongly persistent effect on market index real stock prices from the fourth month to the forty-third month (forty months). We note that, for Thailand, our results are consistent with those of Granger *et al.* (2000). That is, in agreement with the traditional approach, exchange rates lead stock prices with positive correlation. For Malaysia and South Korea, the statistically significantly negative responses of real stock prices to exchange rate depreciation shocks might be attributed to the assembly industry in their countries. This needs to be explored by further research.

For all the countries examined, impulsed by both contractionary monetary policy shocks and by exchange rate depreciation shocks, financial sector index real stock prices react similarly to market index real stock prices, but with a greater magnitude.[17] To show this, we observe from the results for Thailand that, immediately following contractionary monetary policy shocks, the lowest point (-6 percent) of the financial sector is lower than that of the market (-4 percent). In addition, serving as one of our main findings, the figures indicate that contractionary monetary policy shocks result in a strongly persistent effect

[17]See, for example, Chamberlain *et al.* (1997) for the sensitivity of banking stock returns to movements of exchange rates.

on market index real stock prices, in an average of 37.5 months. Having a marginally significant impact only over the short run, the impact of exchange rate depreciation shocks is short-lived, with an average of 7 months. Those results are summarized in Table 2.

For the real GDP, it is worth mentioning that our results acquired are consistent with those of existing literature. That is, aggregate demand shocks have real effects and, empirically, contrationary monetary policy shocks lead to a persistent decline in that variable. See Christiano *et al.* (1996).

4.3   Variance Decomposition Analysis

We turn to a variance decomposition. In particular, we answer the question of how much of the variation in market and financial sector index real stock prices over the sample period is accounted for by monetary policy shocks as compared with exchange rate movements. Table 3 displays, for each country examined, the variance decomposition of market and financial sector index real stock prices with respect to the effect of contractionary monetary policy shocks and exchange rate depreciation shocks.

From the results demonstrated, we argue that, in general, both monetary policy and the exchange rate explain a large share (more than 25 percent on average) of the variation in the market index real stock prices. In detail, for all the countries examined, for twelve months following the shocks, the peak of the fraction of variance explained by exchange rate shocks is larger than that

accounted for by monetary policy shocks.

Served as our second main finding, Table 3 regarding the variance decomposition suggests that exchange rate shocks are as important as monetary policy shocks for explaining the dynamics of market and financial sector index real stock prices. More precisely, for all the countries examined, real exchange rate developments have been more important in the short run.

5   Historical Decomposition

As an alternative discussion on the importance of monetary policy shocks and exchange rate shocks for the Asian stock markets, we show a historical decomposition of stock prices into the contribution of the underlying structural shocks identified in section 3. In other words, the variable of interest, market index stock prices, could be demonstrated as the sum of a deterministic part (baseline) and the contribution of current and past shocks. Complementing the analysis using impulse responses and variance decomposition, historical decomposition provides evidence on (i) whether the underlying structural shocks modeled have actually occurred in reality and (ii) whether the actual developments of the variable of interest, stock prices, could be explained by them.

Sims (1980) developed historical decomposition and Burbidge and Harrison (1985) originally based their analysis upon it. This means that historical decomposition is not a new issue for the analysis using a vector autoregression model.

Nevertheless, the literature performs an analysis using historical decomposition less frequently than using impulse responses and variance decomposition. See Doan (2009) for a description of and, specifically, the appropriate methodology for finding the historical effect of a shock identified by the sign restrictions approach.

We compute the historical decomposition over the selected period, January 1995 to December 1999 and October 2001 to September 2007. This could be thought of as a case-study approach.

For the first period, a consideration for the selection of this period is that it covers the 1997 Asian financial crisis and that the very strong movements of the variables of interest, interest rates, exchange rates and market index stock prices, are evident. For market index stock prices, we provide the actual series, for each country examined, in the top panel of Figures 4 to 6. A significant worsening of the stock markets (and financial conditions) is apparent. The Thai stock market drops by 75 percent. The Seoul stock exchange falls by 4 percent on November 7, 1997. On the next day, it plunges by 7 percent, its biggest one-day drop to that date. Generally, we argue that both the improvement and deterioration of financial conditions are covered in this selected period. Therefore, it presents an interesting case study for the objective of this chapter, that is, to segregate interest rate shocks and exchange rate shocks as simultaneous determinants of stock prices.

The results are presented in Figures 4 to 6. The contribution of monetary

policy shocks and exchange rate shocks to stock prices as deviations from the baseline are shown, together with the actual time series of market index stock prices.

From these figures, our finding reached by analyzing the variance decomposition is strongly supported. While the exchange rate, generally, is as important as monetary policy for explaining the dynamics of market index real stock prices, real exchange rate developments have been more important at some periods of time, for all the countries examined. Therefore, we emphasize the contribution of the exchange rate in the following discussion.

For Malaysia, there is a negative contribution of exchange rate shocks to stock prices from the last quarter of 1997. This negative contribution could be seen continually until the last quarter of 1998, becoming more negative. During such a period, real exchange rate shocks appear to have a larger contribution in explaining the deviations from the baseline. A decrease in the contribution of exchange rate depreciation shocks to stock prices is observed afterwards. For South Korea, exchange rate shocks make a negative contribution to stock prices in 1997. This is even more pronounced in 1998 and larger than that of monetary policy shocks. For Thailand, a negative effect of exchange rate shocks is found from the second half of 1996. The figure shows a decrease in the contribution of both monetary policy shocks and exchange rate shocks from mid-1998 to mid-1999. Nevertheless, in line with our finding attained by analyzing the variance decomposition, the contribution of monetary policy shocks to stock prices is

smaller than that of exchange rate shocks, especially during the last months of the selected period in this historical decomposition.

For the period after the Asian financial crisis, we provide the historical decomposition from October 2001 to September 2007 on our model for South Korea as a case study, shown in Figure 7. The main reason for selecting South Korea is that it had the fastest recovery, compared with other Asian economies, and we believe that this could provide us with the normal stage for an analysis.[18]

The selected period from October 2001 to September 2007 covers the longest and sharpest stock price increase in South Korea since February 1989. Interest rates increased at the beginning of the selected period, which was the first rise since December 2000. Interest rates stayed at a high level before declining in 2003. Corresponding to that period of high interest rates, we observe a negative contribution of monetary policy shocks starting from the second half of 2002 to the first half of 2004. Such a negative contribution could be, apparently, realized again in 2006, corresponding to a continuing contractionary monetary policy implemented in South Korea from the last quarter of 2005. When interest rates decreased in 2003 through 2005, a small positive contribution of monetary policy shocks is observed. For a contribution of exchange rate shocks, it is obvious that a positive effect prevails throughout the selected period. This is reasonable, since the Korean won started to appreciate since October 2001, with a distinguished appreciation from October 2004. With the results provided, it is

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[18]See Koo and Kiser (2001).

reasonable to argue that, during the period of an increase in market index stock prices, e.g. October 2001-May 2002, May 2003-May 2004 and from the second quarter of 2005 to mid-2006, the positive contribution of real exchange rate shocks to stock prices appears to be a dominant explanation for the deviations from baseline.

6   Robustness

To make our analysis complete, in this section, we (i) justify our results using alternative horizons for the sign restrictions to hold, k, (ii) employ a common approach, that is, a recursive ordering of variables, as an alternative identification method and (iii) perform robustness checks.

In order to justify our results obtained using a six-month horizon, $k = 5$, for the sign restrictions to be imposed, an alternative k should also be examined. This is due to the fact that the selection of horizons for the sign restrictions to hold is ad hoc in the estimation. Using the same identification methodology, the literature has chosen different horizons for the sign restrictions to hold.[19] Nevertheless, we argue that we follow the convention of selecting $k = 5$ (a six-month horizon) in reaching our core findings in the previous section. In this section, while we still select the same lag lengths and include a constant as specification in Section 4, we use different horizons for the sign restrictions to

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[19]For example, Uhlig (2005) selects three-, six-, twelve- and twenty-four-month horizons to examine the impact of monetary policy on output.

be imposed, $k = 2$ and $k = 8$.

Figure 8 demonstrates, for all the countries examined, the response of market index real stock prices to contractionary monetary policy shocks and exchange rate depreciation shocks, using three- and nine-month horizons for the sign restrictions to hold. In general, while the results for $k = 8$ are qualitatively similar to those reported in the previous section using $k = 5$, we find that the response of real stock prices to two underlying structural shocks becomes more persistent and lasts for a longer period. Contractionary monetary policy shocks for Thailand and exchange rate depreciation shocks for South Korea show this. More importantly, compared with the results previously reported using $k = 5$, we are more confident in concluding that contractionary monetary policy shocks have a statistically significant effect on the market index real stock prices for Malaysia and exchange rate depreciation shocks over the short run for Thailand. On the other hand, while the results reported in the previous section following the convention of selecting $k = 5$ show a statistically significant response of market index real stock prices to monetary policy shocks for Thailand, such results turn out to be insignificant when the horizon for the sign restrictions to hold is three periods. Exchange rate shocks for Thailand and South Korea also demonstrate this.

In order to examine whether our core findings obtained from sign restrictions are consistent with ones implied by an alternative identification method, a common approach, that is, a recursive ordering of variables, is employed. In

particular, we estimate our vector autoregression model and use Cholesky decomposition in order to achieve identification. The ordering of the six variables corresponds to equation (3). Impulse responses obtained from this method, for each country examined, are not shown here due to space limitations, but available upon request. While counter-intuitive impulse responses of some endogenous variables are found, a comparison of the impulse responses of market index real stock prices to monetary policy shocks and exchange rate shocks with those acquired from sign restrictions shows consistency in general. One of our core findings is observed. That is, the effect of monetary policy on real stock prices is more persistent than that of the exchange rate and the impact of exchange rate shocks is short-lived over the short run. In addition, the variance decomposition obtained from the current method also suggests that, in the short run, the impact of the exchange rate in explaining the dynamics of real stock prices is more important than that of monetary policy. Nevertheless, for Thailand, the results show that real exchange rate developments negligibly explain the dynamics of real stock prices.[20]

In addition, we perform robustness checks, concerning three issues, in the following.

Firstly, frequent and dramatic regime change in Asian economies makes it hard to select a sample period for an empirical study. In addition, the study

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[20]For Thailand, monetary policy (exchange rate) explains 4.71 (0.69) of the variation in the market index real stock prices; 20.67 (21.48) for Malaysia; and 8.02 (10.05) for South Korea. These figures correspond to the peak during the first six months following the shocks.

needs to account in some way for the fact that the 1997-1998 Asian financial crisis has significantly affected the dynamics in those economies.[21] Taking into account the period of the Asian crisis, we add a dummy variable to isolate such a period, as commonly performed in the literature. In particular, the dummy variable takes a value of one for the period of the Asian financial crisis and zero otherwise. To be uniform across the three countries examined, the beginning and end of the crisis period are set to July 1997 and September 1998.[22]

Secondly, given the open economy nature of the countries examined, we add a variable measuring the economic conditions in the US and world interest rates. In particular, the Federal Funds rate is used as an exogenous variable and an endogenous variable.[23]

Thirdly, taking into account the fixed exchange rate period in Malaysia (September 1998 - June 2005) and Thailand (before July 1997), a dummy variable isolating such a period and real effective exchange rates are used to examine the robustness of our results. In addition, nominal exchange rates are also employed for the three countries examined.

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[21]To the best of our knowledge, empirical studies pursuing the sign restrictions approach use data over a long period of time and, hence, find it hard to avoid breaks in the series. This might be due to an intensive computation of that methodology requiring such long series of data.

[22]Using a shorter data set from January 1995 to November 2008 also allows us to reach similar conclusions. However, the analysis based on that shorter data set would yield a shortened horizon of impulse responses after the shock that could be appropriately calculated, 24 months ahead in our experiment. Nevertheless, as commonly expected, and so as to arrive at our conclusion, that there might be a lag in effect of monetary policy and that its effect would last for another couple of years, we prefer, reasonably, the analysis with a long data set that could provide us impulse responses for a longer horizon after the shock. Note also that Uhlig (2005) studies the effect of monetary policy for a horizon of up to 60 months after the shock.

[23]In order to preserve degrees of freedom, the exogenous variables enter the system only contemporaneously.

Overall, the final products of the VAR models including impulse responses and variance decomposition are robust to the inclusion of the Asian crisis dummy, the Federal Funds rate, the dummy isolating the fixed exchange rate period and the use of the real effective exchange rate and nominal exchange rate.[24] For the first issue, this is consistent with the existing literature. In particular, Granville and Mallick (2009) emphasize that Uhlig's methodology is robust to non-stationarity of series including breaks and, consequently, do not include any dummy variables in the system. Relating to the second issue, we could, consequently, argue that the countries examined do not adjust their interest rates systematically in response to the economic conditions in the US. The robust results might be explained by the fact that the events in the US economy relevant to the decisions of the monetary authority in each country examined are already reflected in the exchange rates between the domestic currency and the US dollar.

7   Conclusion

In this chapter, we employ a Bayesian structural vector autoregression model to examine the effect of monetary policy and of the exchange rate on stock price

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[24]Using the real effective exchange rate for Thailand shows that monetary policy explains a larger share of the variation in the market index real stock prices than the exchange rate, for twelve months following the shocks. Nevertheless, our main finding that real exchange rate developments have been more important in the short run still applies to financial sector index real stock prices, not to mention that exchange rate shocks are as important as monetary policy shocks for explaining the dynamics of real stock prices.

movements in Asia. Sign restrictions are used to identify simultaneously and uniquely contractionary monetary policy and exchange rate depreciation shocks in an integrated framework. Our findings are obtained by adapting standard VAR analyses to deal with these single shocks based upon sign restrictions. In particular, for each country examined, impulse responses, variance decomposition and historical decomposition are deliberately considered.

Two main findings emerge. Firstly, monetary policy shocks result in a strongly persistent effect on market index real stock prices whereas the impact of exchange rate shocks is short-lived over the short run. Secondly, with respect to the variance decomposition, the exchange rate is as important as monetary policy for explaining the dynamics of market and financial sector index real stock prices. More precisely, for all the countries examined, real exchange rate developments have been more important in the short run. In addition, the historical decomposition strongly supports our finding reached by analyzing the variance decomposition. While the exchange rate, generally, is as important as monetary policy for explaining the dynamics of market index real stock prices, real exchange rate shocks appear to have a larger contribution in explaining the deviations from the baseline in some periods of time, for all the countries examined. Based purely on our findings, two conclusions are reached. Firstly, because of the mistimed and/or persistent effect of monetary policy on both real economy and financial markets, we argue that one needs to be cautious in using monetary policy to constrain asset price misalignment. Secondly, due to

the evidence that exchange rates principally have a contemporaneous impact on equity prices, we suggest that, in the short run, such incorrectly aligned asset prices might potentially be corrected by focusing on exchange rate movements.

For the first conclusion, while we establish that, in general, monetary authorities could significantly affect equity market valuations by adjusting interest rates, the persistent effects of monetary policy on stock prices found might shed light on the effectiveness of such policy in developing countries in stabilizing the economy. This should be of concern to the central bank. That is, mistimed and/or too persistent effects of monetary policy would only make situations worse if stock price misalignment does not result in significant damage when it ends or if interest rates are high at the moment that a bubble bursts. Moreover, the impact of such high interest rates on the real economy would last for another couple of years and make the landing harder. This is what we have seen in Asian countries in the aftermath of the 1997 Asian financial crisis. Furthermore, the undesirable effect of such mistimed and/or too persistent monetary policy on financial markets is reflected by the fact that aggregate demand depends positively on the past level of asset prices via the investment balance sheet channel.[25]

From that point of view, we shift to the question of whether other instruments besides interest rates might be used to deal with asset price misalignment. Definitely, reasonable bank regulation and supervision should be thoroughly consid-

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[25]The investment balance sheet effects imply that there is a positive relationship between firms' ability to borrow and their net worth, which in turn relies on valuations of assets.

ered. Strong consensus that a well-structured prudential policy and regulatory system could make financial markets and financial systems less prone to troublesome situations is achieved. This is by helping to reduce the costs of stock price booms and bursts. With these alternatives, the need for contractionary monetary policy conducted by the central bank to burst a bubble is likely to be reduced. They also contribute to the stability of both output and inflation. Nevertheless, deciding what form such regulation and supervision should take is the more difficult issue. It has recently been argued that, in principle, banking regulation should change cyclically to rule out lending booms on the back of rises in asset prices. This should also be considered carefully.

For the second conclusion, confirming that exchange rates are an important determinant of stock prices, our results obtained using recent and growing methodology on the basis of sign restrictions are in line with those of existing literature employing the common method. That is, exchange rates principally have only a contemporaneous impact on equity prices, resulting in effects in the short run. The implication of this finding might be that, in the short run, asset price misalignment might potentially be corrected by smoothing excessive exchange rate fluctuations.

We have seen that, in emerging markets with large foreign-denominated debt in particular, a financial crisis could be triggered by sharp depreciation. Exchange rate fluctuations are of major concern to monetary authorities, even if they are targeting inflation. As suggested by Mishkin and Savastano (2001),

(inflation-targeting) central banks should not pursue a policy of benign neglect of exchange rates. In other words, monetary authorities in these countries may have to smooth excessive exchange rate fluctuations parted from fundamentals via, for example, foreign exchange market interventions. Nevertheless, this should be conducted without resisting market-determined movements in exchange rates over longer horizons. Such interventions lighten potentially destabilizing impacts of unexpected changes in exchange rates, which, at least in our context, have a contemporaneous impact on asset prices. Nevertheless, it is challenging for central banks, especially ones targeting inflation, to focus on exchange rate movements since this might obstruct them reaching the target rate of inflation.

With the information provided, the appropriate policy response to potential misalignments of stock prices would be calibrated by the central bank.

Figure 1: Impulse Responses to Contractionary Monetary Policy Shocks and to Exchange Rate Depreciation Shocks, 6-Variable VAR, Pure-Sign Approach, Thailand, Feb 89 to Nov 08, Stock Prices in Level

Monetary Policy Shocks and Exchange Rate Shocks, Thailand
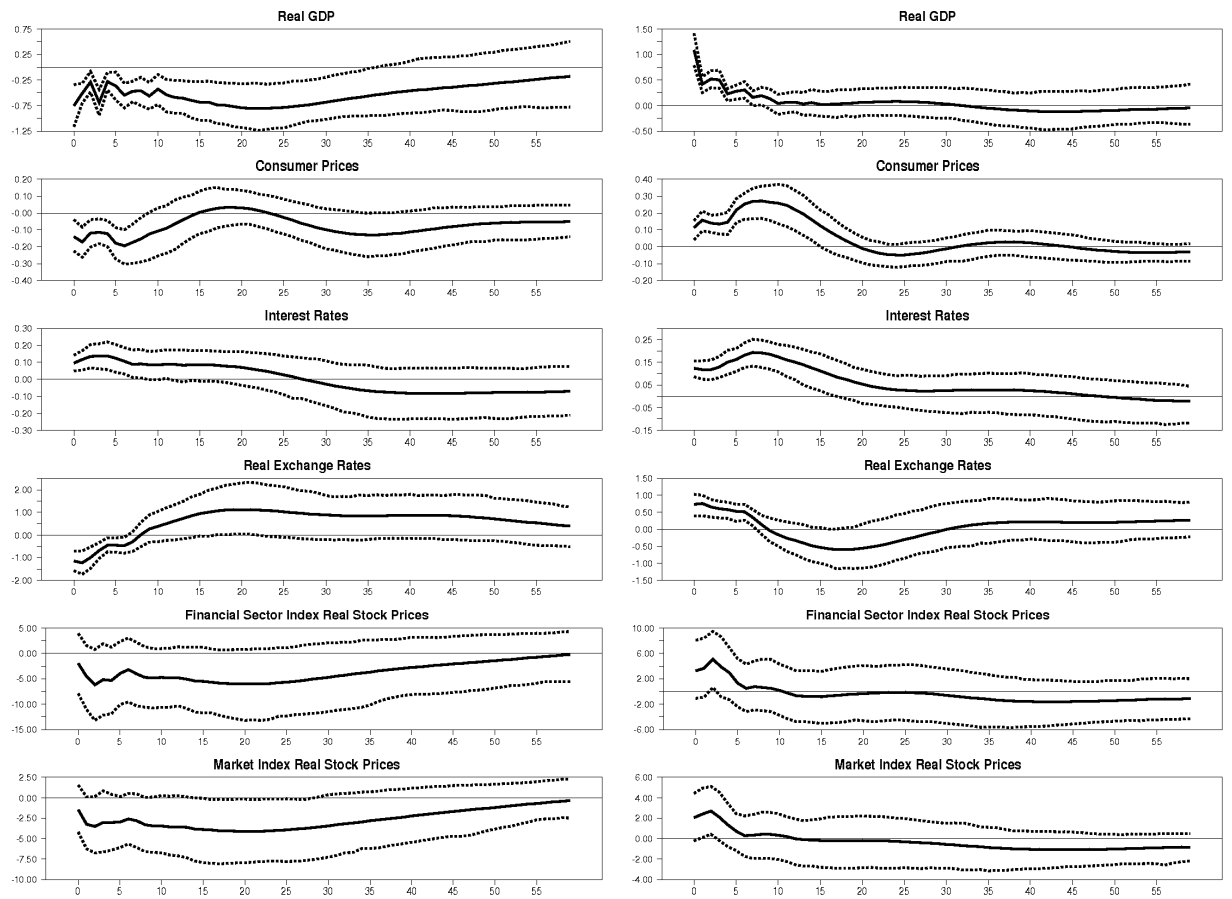
Figure 2: Impulse Responses to Contractionary Monetary Policy Shocks and to Exchange Rate Depreciation Shocks, 6-Variable VAR, Pure-Sign Approach, Malaysia, Feb 89 to Nov 08, Stock Prices in Level

Monetary Policy Shocks and Exchange Rate Shocks, Malaysia
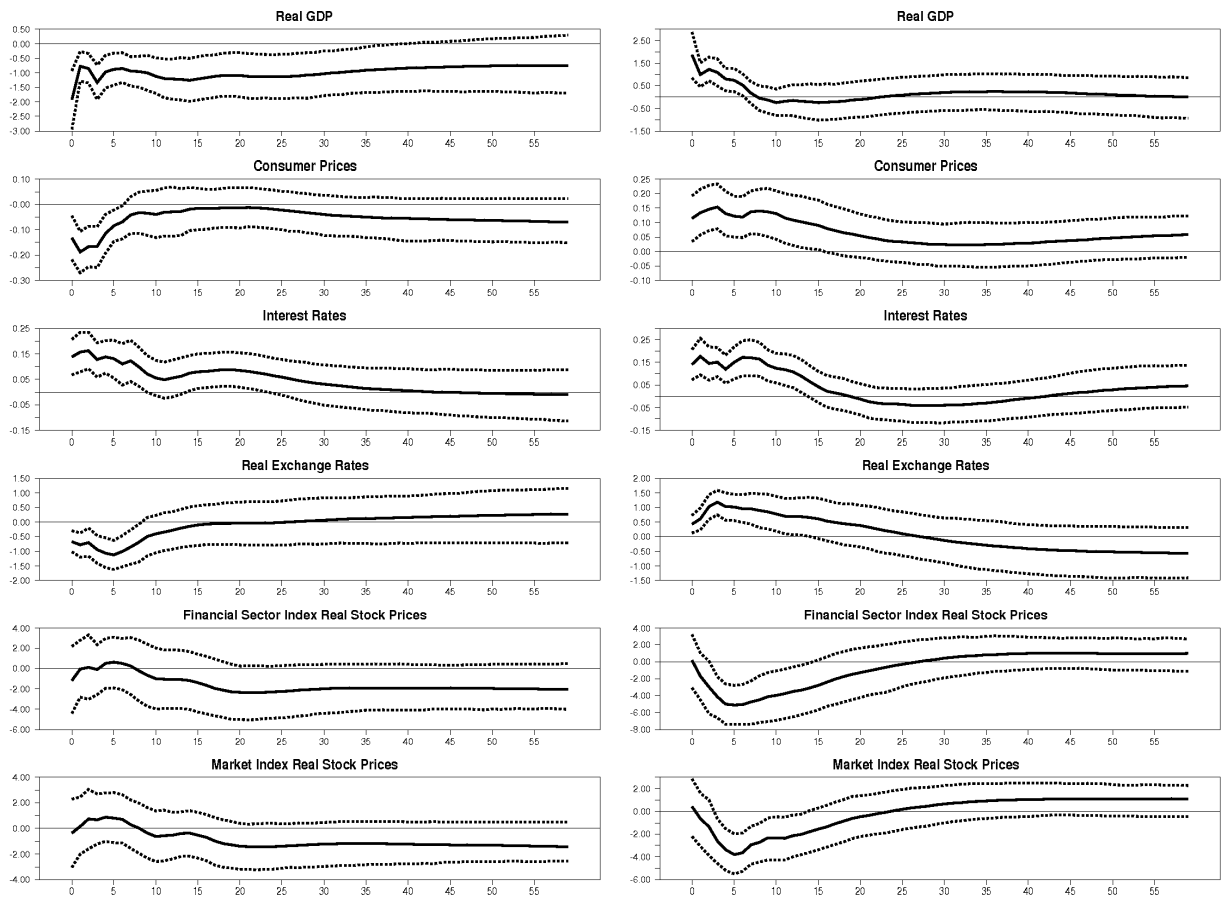
Figure 3: Impulse Responses to Contractionary Monetary Policy Shocks and to Exchange Rate Depreciation Shocks, 6-Variable VAR, Pure-Sign Approach, South Korea, Feb 89 to Nov 08, Stock Prices in Level

Monetary Policy Shocks and Exchange Rate Shocks, South Korea
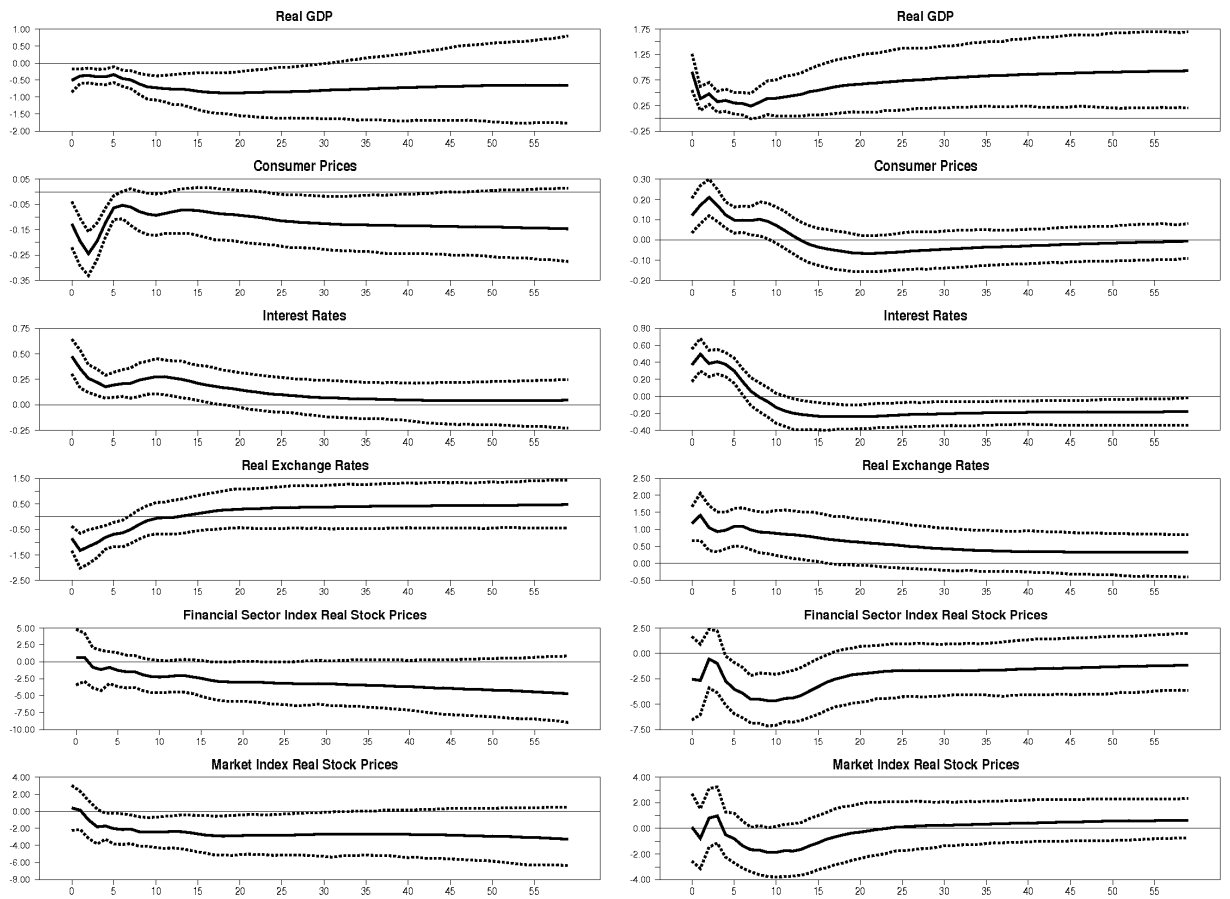
Table 2: Summary Results from Market Index Real Stock Prices Impulse Responses to Contractionary Monetary Policy Shocks and to Exchange Rate Depreciation Shocks

Contract MP Shocks		ER Depreciation Shocks	
Thailand			
Relationship:	Negative	Relationship:	Positive
Significant	1st mth-35th mth (35 mths)	Significant:	1st mth-4th mth (4 mths)
Persistent:	Yes	Persistent:	No
Magnitude:	f bigger than s	Magnitude:	f bigger than s
Course reversion:	Yes	Course reversion:	Yes
Malaysia			
Relationship:	Negative	Relationship:	Negative
Significant:	No	Significant:	3rd mth-15th mth (13 mths)
		Persistent:	No
		Magnitude:	f bigger than s
		Course reversion:	Yes
South Korea			
Relationship:	Negative	Relationship:	Negative
Significant:	4th mth-43rd mth (40 mths)	Significant:	7th mth-10th mth (4 mths)
Persistent:	Yes	Persistent:	No
Magnitude:	f bigger than s	Magnitude:	f bigger than s
Course reversion:	No	Course reversion:	Yes

Table 3: Variance Decomposition for Market and Financial Sector Index Real Stock Prices, Reported at Peak During 12 Months Following the Shocks

Variable	Contract MP	Depreciation	Both
Thailand			
Fin sector	10.71 (11 months)	12.70 (1 month)	23.29 (11 months)
Market	8.84 (12 months)	10.87 (12 months)	19.71 (12 months)
Malaysia			
Fin sector	16.19 (11 months)	17.53 (9 months)	33.69 (11 months)
Market	15.56 (12 months)	17.02 (12 months)	32.58 (12 months)
South Korea			
Fin sector	16.92 (12 months)	16.20 (12 months)	33.12 (12 months)
Market	15.23 (12 months)	15.68 (12 months)	30.91 (12 months)

Figure 4: Actual Market Index Stock Prices and the Contribution of Monetary Policy
Shocks and of Exchange Rate Shocks to Stock Prices, 6-Variable VAR, Thailand,
Jan 95 to Dec 99
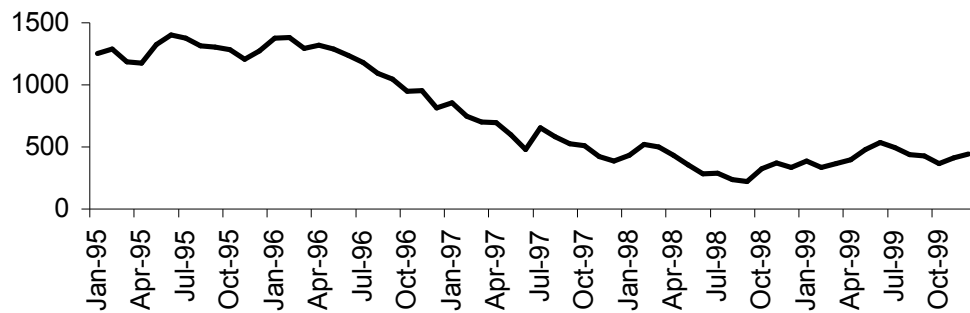
(*Actual Market Index Stock Prices*)
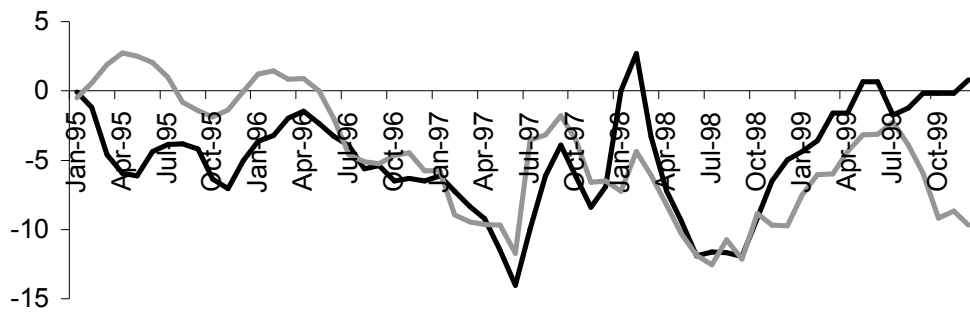
(*Contribution of Monetary Policy Shocks (Black) and of Exchange Rate Shocks (Grey) to Stock Prices*)

Figure 5: Actual Market Index Stock Prices and the Contribution of Monetary Policy
Shocks and of Exchange Rate Shocks to Stock Prices, 6-Variable VAR, Malaysia,
Jan 95 to Dec 99
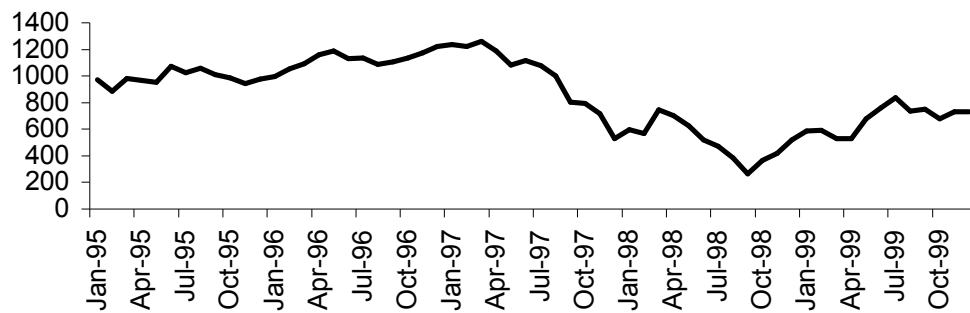
(*Actual Market Index Stock Prices*)
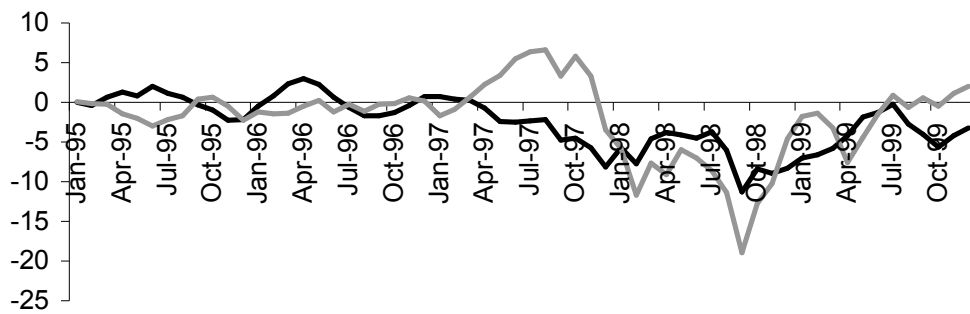
(*Contribution of Monetary Policy Shocks (Black) and of Exchange Rate Shocks (Grey) to Stock Prices*)

Figure 6: Actual Market Index Stock Prices and the Contribution of Monetary Policy Shocks and of Exchange Rate Shocks to Stock Prices, 6-Variable VAR, South Korea, Jan 95 to Dec 99
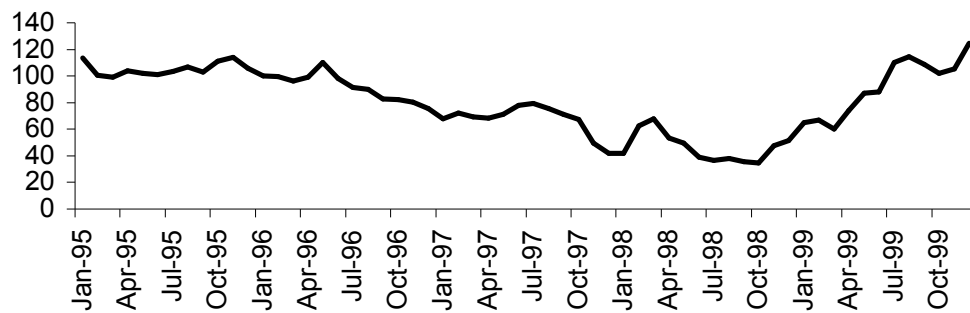
(*Actual Market Index Stock Prices*)
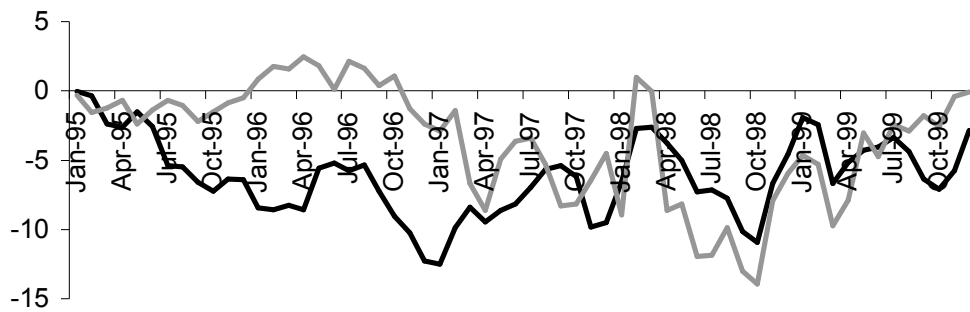
(*Contribution of Monetary Policy Shocks (Black) and of Exchange Rate Shocks (Grey) to Stock Prices*)

Figure 7: Actual Market Index Stock Prices, the Contribution of MP Shocks
and of ER Shocks to Stock Prices, Interest Rates and Exchange Rates, South Korea,
Oct 01 to Sep 07
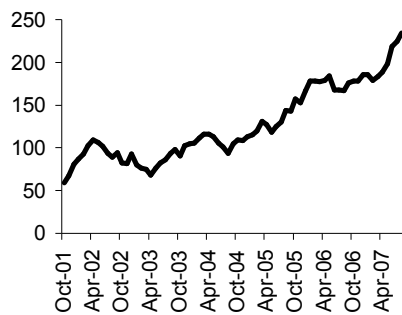
(*Actual Market Index Stock Prices*)
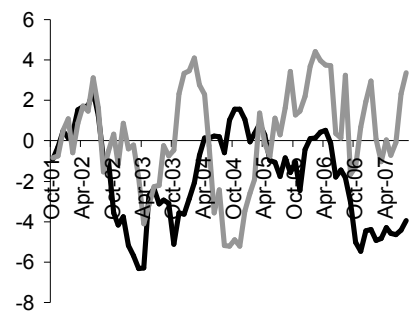
(*Contrib of MP (Black) and of ER (Grey) Shocks*)

(*Interest Rates*)

(*Exchange Rates*)

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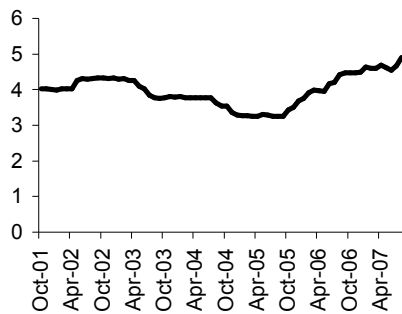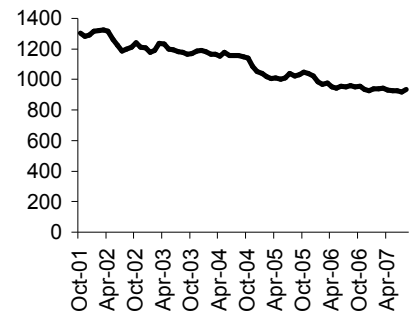
Figure 8: Impulse Responses of Market Index Real Stock Prices to Contractionary
Monetary Policy Shocks and to Exchange Rate Depreciation Shocks, Using Different
Horizons for the Sign Restrictions

(*k=2, Thailand*)   (*k=8, Thailand*)

(*k=2, Malaysia*)   (*k=8, Malaysia*)

(*k=2, South Korea*)   (*k=8, South Korea*)
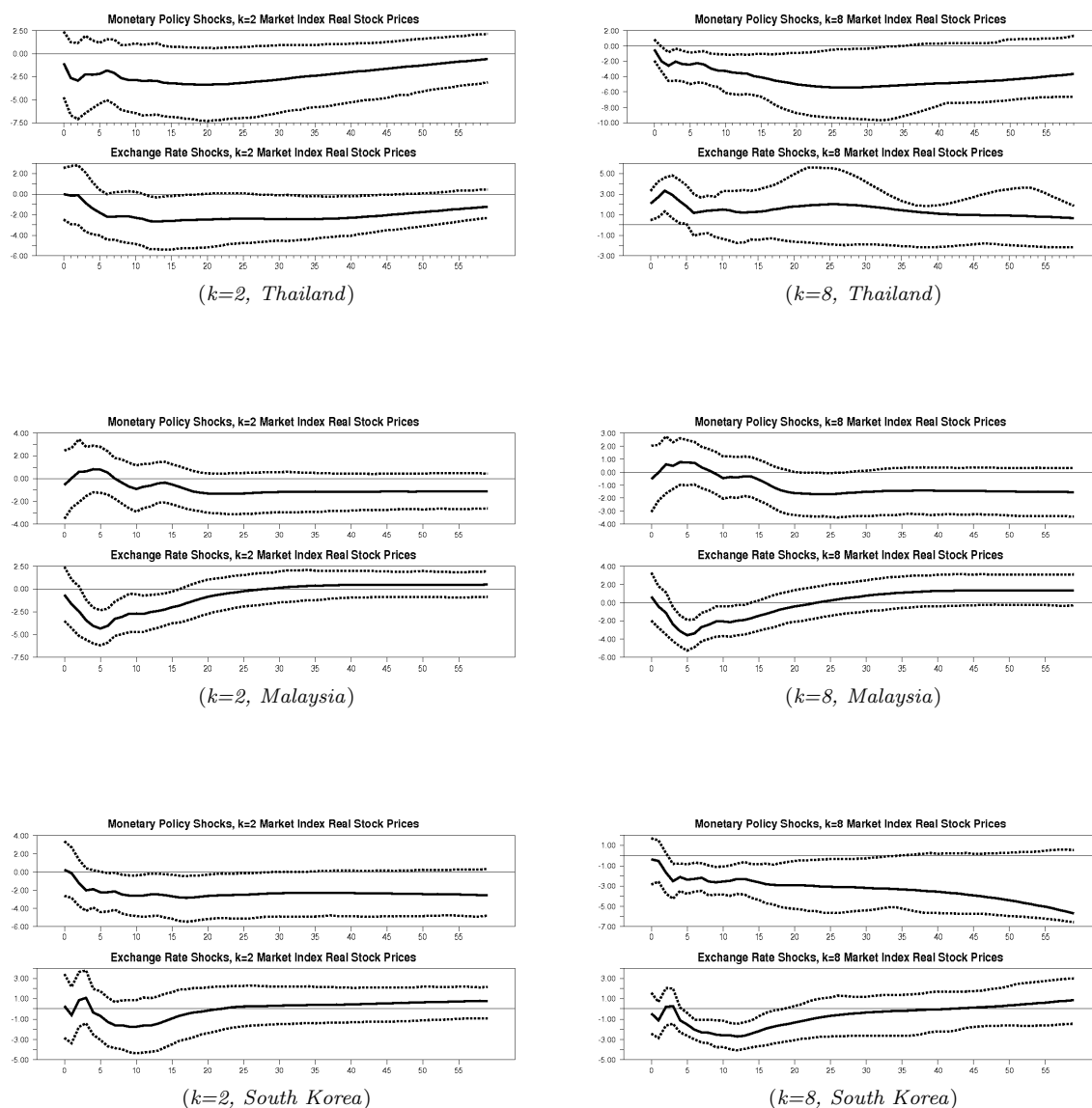
References

Ahmed, M. K., Md.Akhtaruzzaman, and Barua, S. (2006). Effects of monetary policy on price formation of financial assets: A test for Bangladesh. *Policy Analysis Unit, Research Department, Bangladesh Bank Working Paper Series*, (WP 0703).

An, L. (2006). Exchange rate pass-through: Evidence based on vector autoregression with sign restrictions. *Munich Personal RePEc Archive*, (527).

Assenmacher-Wesche, K. and Gerlach, S. (2008). Monetary policy, asset prices and macroeconomic conditions: A panel-VAR study. *National Bank of Belgium Research Series*, (149).

Baharumshah, A. Z., Masih, A. M. M., and Azali, M. (2002). The stock market and the ringgit exchange rate: A note. *Japan and the World Economy*, 14(4):471–486.

Barsky, R. B. (1989). Why don't the prices of stocks and bonds move together? *American Economic Review*, 79(5):1132–1145.

Bjornland, H. C. and Halvorsen, J. I. (2008). How does monetary policy respond to exchange rate movements? New international evidence. *Research Department, Norges Bank Working Paper Series*, (2008/15).

Bjornland, H. C. and Leitemo, K. (2009). Identifying the interdependence between US monetary policy and the stock market. *Journal of Monetary Economics*, 56(2):275–282.

Blanchard, O. J. and Quah, D. (1989). The dynamic effects of aggregate demand and supply disturbances. *American Economic Review*, 79(4):655–673.

Bollard, A. (2004). Speech by Alan Bollard: Asset prices and monetary policy. *Reserve Bank of New Zealand Bullentin*, 67(1):27–34.

Bordo, M. D., Dueker, M. J., and Wheelock, D. C. (2009). Inflation, monetary policy and stock market conditions: Quantitative evidence from a hybrid latent-variable VAR. *Research Division, Federal Reserve Bank of St. Louis Working Paper Series*, (2008-012B).

Borio, C. and Lowe, P. (2002). Asset prices, financial and monetary stability: Exploring the nexus. *Bank of International Settlement Working Paper Series*, (114).

Bracke, T. and Fidora, M. (2008). Global liquidity glut or global savings glut? A structural VAR approach. *European Central Bank Working Paper Series*, (911).

Burbidge, J. and Harrison, A. (1985). An historical decomposition of the great depression to determine the role of money. *Journal of Monetary Economics*, 16(1):45–54.

Canova, F. and Nicolo, G. D. (2002). Monetary disturbances matter for business fluctuations in the G-7. *Journal of Monetary Economics*, 49(6):1131–1159.

Cecchetti, S. G., Genberg, H., Lipsky, J., and Wadhwani, S. B. (2000). *Asset prices and central bank policy)*. Centre for Economic Policy Research.

Chamberlain, S., Howe, J. S., and Popper, H. (1997). The exchange rate exposure of US and Japanese banking institutions. *Journal of Banking and Finance*, 21(6):871–892.

Christiano, L. J., Eichenbaum, M., and Evans, C. (1996). The effects of monetary policy shocks: Evidence from the flow of funds. *The Review of Economics and Statistics*, 78(1):16–34.

Doan, T. (2009). Understanding the historical decomposition. .

Eichenbaum, M. (1992). Interpreting macroeconomic time series facts: The effects of monetary policy: Comments. *European Economic Review*, 36(5):1001–1011.

Evans, C. L. (2009). Remarks by Charles L. Evans: The international financial crisis: Asset price exuberance and macroprudential regulation. .

Fratzscher, M., Juvenal, L., and Sarno, L. (2008). Asset Prices, Exchange Rates and the Current Account. *Research Division, Federal Reserve Bank of St. Louis Working Paper Series*, (2008-031B).

Goodhart, C. A. E. and Hofman, B. (2007). *House prices and the macroeconomy: Implications for banking and price stability)*. Oxford University Press, Oxford.

Granger, C. W. J., Huang, B.-N., and Yang, C.-W. (2000). A bivariate causality between stock prices and exchange rates: evidence from recent Asian flu. *The Quarterly Review of Economics and Finance*, 40(3):337–354.

Granville, B. and Mallick, S. (2009). Monetary and financial stability in the euro area: Pro-cyclicality versus trade-off. *Journal of International Financial Markets, Institutions and Money*.

James, W. E., Naya, S., and Meier, G. M. (1989). *Asian development: Economic success and policy lessons*. University of Wisconsin Press, Wisconsin Madison.

Kent, C. and Lowe, P. (1997). Asset-price bubbles and monetary policy. *Economic Research Department, Reserve Bank of Australia Research Discussion Paper Series*, (9709).

Kim, S. (2002). Exchange rate stabilization in the ERM: Identifying European monetary policy reactions. *Journal of International Money and Finance*, 21(3):413–434.

Koo, J. and Kiser, S. L. (2001). Recovery from a financial crisis: The case of South Korea. *Economic and Financial Policy Review*, (Q IV):24–36.

Lastrapes, W. D. (1998). International evidence on equity prices, interest rates and money. *Journal of International Money and Finance*, 17(3):377–406.

Mishkin, F. S. and Savastano, M. A. (2001). Monetary policy strategies for Latin America. *Journal of Development Economics*, 66(2):415–444.

Mountford, A. and Uhlig, H. (2008). What are the effects of fiscal policy shocks? *NBER Working Paper Series*, (14551).

Neri, S. (2004). Monetary policy and stock prices: theory and evidence. *Economic Research Department, Bank of Italy Working Paper Series*, (513).

Paustian, M. (2007). Assessing sign restrictions. *The B.E. Journal of Macroeconomics*, 7(1).

Phylaktis, K. and Ravazzolo, F. (2005). Stock prices and exchange rate dynamics. *Journal of International Money and Finance*, 24(7):1031–1053.

Rajan, R. S., Thangavelu, S., and Parinduri, R. A. (2008). *Exchange rate, monetary and financial issues and policies in Asia*. World Scientific Publishing Company.

Rapach, D. E. (2001). Macro shocks and real stock prices. *Journal of Economics and Business*, 53(1):5–26.

Reisen, H. (1993). The impossible trinity in South-East Asia. *International Economic Insights*.

Shiller, R. J. and Beltratti, A. E. (1992). Stock prices and bond yields: Can the comovements be explained in terms of present value models? *Journal of Monetary Economics*, 30(1):25–46.

Sims, C. (1980). Comparison of Interwar and Postwar Business Cycles: Monetarism Reconsidered. *American Economic Review*, 70(2):250–257.

Sims, C. A., Stock, J. H., and Watson, M. W. (1990). Inference in linear time series models with some unit roots. *Econometrica*, 58(1):113–144.

Uhlig, H. (2005). What are the effects of monetary policy on output? Results from an agnostic identification procedure. *Journal of Monetary Economics*, 52(2):381–419.

Chapter 3: Fear, Volatility and Uncovered Interest Parity

Mark P. Taylor[*] and Tim Leelahaphan[†]

Abstract

Within the context of a time-varying transition probabilities Markov-switching model of the uncovered interest parity (UIP) condition, we examine if variables measuring fear and volatility have an effect on the probability of switching between the regime where the UIP condition holds and the regime where it does not. The state transition probability depends nonlinearly upon the variables examined. These are the exchange rate volatility, the VIX equity option implied volatility index and the TED spread. Applying this to both US dollar exchange rates and cross (exchange) rates from January 4, 1990 to September 11, 2008, we find that those three variables increase the probability of remaining in the regime where the UIP condition holds. In addition, the probability of switching from the regime where the UIP condition does not hold to the regime where the exchange rate follows the UIP condition decreases (increases) as these variables measuring fear and volatility fall (rise), especially the VIX equity option implied volatility index. For JPYAUS, JPYNZD and USDJPY, the smoothed probabilities show that these exchange rates essentially do not follow the UIP condition except during periods in which the fear and risk variables are increasing, as in the recent global financial crisis in particular.

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[*]Barclays Global Investors and The Department of Economics, The University of Warwick, UK.

[†]Contact information: The Department of Economics, The University of Warwick, Coventry, CV4 7AL, United Kingdom; email T.Leelahaphan@warwick.ac.uk. Tim Leelahaphan is indebted to Tom Doan and Tom Maycock for their invaluable help on RATS programming. Tim Leelahaphan would like to honestly thank Professor Michael P. Clements for his comments and sincerely appreciates Professor Ian W. Marsh sharing his RATS codes and his guidance during the early stage of this chapter.

1 Introduction

In international macroeconomics and international finance models, the uncovered interest parity (UIP) condition is commonly assumed. This parity condition claims that, over some particular time horizon, the foreign exchange's expected gain from holding funds in one currency rather than the other should be compensated by the opportunity cost of holding that currency instead of another. In particular, an interest rate differential between the two currencies must offset expected changes in the exchange rate under the UIP condition. In addition, one might hypothesize that market participants would expect currencies with higher interest rates to fall in value.

However, Fama (1984), a highly influential paper, notes instead that such currencies with high interest rates tend to appreciate. This is inconsistent with the commonly assumed UIP condition. With this result, we could infer that there would be an inverse relationship between the forward premium and the future exchange rate changes.[1] Evidently, for many currencies and periods examined, this remark in line with Fama (1984) has been asserted in the literature.

Although more than two decades have passed since Fama (1984), that deviation from the UIP condition has not yet been explained, as reviewed in Wagner (2008). From an economic perspective, the existence of a risk premium or market inefficiency might be attributable to such a deviation from the UIP condition. Nevertheless, risk premia models-based research has been limited in its

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[1]This is explained in detail in Section 2.

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success in explaining the deviation from the UIP condition convincingly. This is especially true for plausible degrees of risk aversion. In addition, explanations such as learning, peso problems, bubbles, consumption-based asset pricing and term-structure models have been attempted to be able to explain the issue convincingly. Nevertheless, they have not been successful.[2] Essentially, the consensus explanation of such a deviation from the parity condition has not yet been reached. Macroeconomists are still attempting to solve this issue. In fact, a so-called 'forward bias puzzle' or 'failure of the UIP condition' is still one of the most important puzzles in the area of international macroeconomics and international finance.

The deviation from the UIP condition leads to the result that profit could be generated through 'currency carry trade'. This is a strategy in which a market participant takes a short position on a currency with low interest rates, which is called the funding currency, and takes a long position on a currency with high interest rates, which is called the target currency. This is due to the fact that such failure of the UIP conditions indicates that the target currency (with a high interest rate) does not depreciate against the funding currency (with a low interest rate) by a percentage that matches the interest rate differential between the two currencies. Therefore, if that failure of the UIP condition actually prevails, a positive return would then be made.

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[2]For issues relating to models of risk premia, see e.g. Cumby (1988), Hodrick (1989), and Bekaert *et al.* (1997). See Lewis (1995) for issues relating to learning, peso problems and bubbles, Backus *et al.* (1993) and Bekaert (1996) for consumption-based asset pricing models and Backus *et al.* (2001) for term-structure models.

Gyntelberg and Remolona (2007) link the currency carry trade to, essentially, a bet against the UIP condition. They study the profitability of foreign exchange carry trades that use the Japanese yen and Swiss franc as funding currencies due to their low interest rates. They reveal that the currency carry trade is used when the interest rate differential between the two currencies is attractively high enough. This is in order to compensate traders for an underlying risk in foreign exchange markets. They find evidence supporting the opinion that downside risk is an important feature of currency carry trade. In addition, using measures of downside risk (as opposed to the standard deviation), their evidence affirms the view that such downside risk decreases the Sharpe ratio (an excess return, or risk premium, per unit of risk).[3] Nevertheless, foreign exchange carry trades generate higher Sharpe ratios than those market participants could obtain from equity markets.

Suominen *et al.* (2008) point out that currency carry trading is a main strategy for hedge funds and use regression analysis to provide new empirical evidence on the effects of foreign exchange carry trades. According to their results, foreign exchange carry trade activity of hedge funds has had a statistically significant effect on both interest rates and exchange rates across the world. This might be due to the very large increase in assets under management (AUM) in

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[3]The Sharpe ratio, the ratio of expected return to the variability of returns, is a benchmark by which many portfolio managers measure their investment performance. It is defined as $SR = \frac{R-R_f}{\sigma} = \frac{E[R-R_f]}{\sqrt{var[R-R_f]}}$, where an asset return is denoted by R, R_f shows a risk free rate of return. Therefore, $E[R-R_f]$ provides an expected value of the asset excess return over the risk free rate of return (or risk premium). σ shows a standard deviation of the excess of the asset return over the risk free rate of return.

hedge funds in recent years. Due to a rise in capital dedicated to foreign exchange carry trade activity, the currencies with high interest rates, or carry trade long currencies (target currency), have appreciated and the currencies with low interest rates, or carry trade short currencies (funding currency), have depreciated over the past several years. According to their estimates, because of this, the expected returns to foreign exchange carry trades are highly positive until recently. Nevertheless, they have been decreasing over time.

Recently, one consensus relating to foreign exchange carry trades has been likely to be reached by many market participants and monetary authorities. That is, the failure of the UIP condition (or the appreciation of currencies with high interest rates) has been associated with the currency carry trade activities with a trend lower in volatility. Supporting this view, de Rato (2007), the managing director of the International Monetary Fund (IMF), mentions in his speech that the currency carry trade reflects environments with both low volatilities and large interest rate differentials. Such conditions support market participants acquiring high excess returns per unit of risk (or the Sharpe ratio) measures of such strategy. In addition, this has placed downward pressure on the currencies with low interest rates. The speech addresses rapid reversal movements of exchange rates caused by the unwinding of the foreign exchange carry trade positions.[4]

In contrast to the currency environment associated with carry trade, the crisis

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[4]See Koyama and Ichiue (2008).

in foreign exchange markets affected by the recent global financial crisis began in August 2007.[5] This is when subprime-related turmoil in other asset classes finally spilled over into foreign exchange markets. At the early phase of the crisis, a major currency carry trade sell-off was evident. In November 2007, in response to restrictions on credit, a major deleveraging in financial markets was observed. Consequently, many investment funds were forced to liquidate their positions. In March 2008, the crisis intensified in the wake of the near-failure of Bear Stearns, a large investment bank, Citigroup and AIG. At this stage in the crisis, the authority's rescue of these too-big-to-fail firms coupled with the orderly takeover of Bear Stearns by JP Morgan Chase appeared to relieve the financial markets. As a result, financial markets returned to some semblance of normality. In September 2008, associated with the failure of Lehman Brothers, the crisis reached its peak (at least, so far). By any measure, with the Lehman Brothers bankruptcy, we are faced in this period with an unprecedented, in its scale and depth, crisis. Specifically, in foreign exchange markets, the levels of currency volatility we experience are unlike those that have gone before. In addition, with unseen levels of counterparty risk, liquidity disappears. The cost of trading currencies soars and consequently, trading of any substantial size becomes very difficult.

Furthermore, during the recent global financial crisis, a significant unwinding of currency carry trade activity associated with a continuously vast capital out-

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[5]See Melvin and Taylor (2009) for an overview of the important events of the crisis.

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flow from the hedge funds was observed. This resulted in remarkable movements in exchange rates. That is, we observed a large appreciation of the funding currencies, that is, currencies with low interest rates, and a large depreciation of the target currencies, that is, currencies with high interest rates. Specifically, the former corresponded to the Japanese yen, US dollar and Swiss franc and the latter to the euro and British pound. See also Suominen *et al.* (2008).[6]

Due to the above information provided, there should be a close link between fear, volatility and the performance of the currency carry trade associated with the failure of the UIP condition. Higher fluctuations in the financial markets, including prices of not only exchange rates but also stocks and bonds, could worsen excess returns that could be generated through the currency carry trade activities. This is due to the fact that many investors analyze performance, not only by returns, but also by the variability of those returns. That is, they measure their investment performance by the Sharpe ratio in particular. An increase in volatility with no change in return would decrease the Sharpe ratio

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[6]The Japanese yen started appreciating against the US dollar when the global financial crisis started in August 2007. However, a period of Japanese yen depreciation was evident afterwards. Specifically, such a period was realized after the authority's rescue of too-big-to-fail firms, the orderly takeover of Bear Stearns by JP Morgan Chase and the financial markets' return to some semblance of normality. That period of depreciation terminated in September 2008. In the post-Lehman period, it could be argued that the Japanese yen benefited from three factors. These are (i) the unwinding of currency carry trade activities; (ii) the rise in interest rates in Japan which in turn reduces its interest rate disadvantage to other currencies; and (iii) a view that the Japanese yen is a safe-haven currency owing to higher global risk aversion and plunging asset prices. Hattori and Shin (2007), similar to Gyntelberg and Remolona (2007), find evidence pointing out that volumes of currency carry trade involving the Japanese yen are high when the interest rate differential against the Japanese yen is wide. In addition, Japanese banks were not much affected by the exposures to the US subprime crisis. This, on the contrary, was not the case for their competitors in Europe and the US. Finally, associated with a growing contraction of the Japanese economy, that safe-haven notion began to disappear afterwards in the early part of 2009.

and a lower Sharpe ratio is less favorable. Consequently, it might be difficult for the foreign exchange carry trade activities to sustain the type of performance it has had if the environment with low volatility does not persist.

Motivated by the above arguments, we examine if variables measuring fear and volatility have an effect on the probability of switching between the regime where the UIP condition holds and the regime where it does not. This is within the context of a time-varying transition probabilities Markov-switching model of the UIP condition. The state transition probability depends nonlinearly upon the variables examined. These are the exchange rate volatility, the VIX equity option implied volatility index and the TED spread.

Applying this to both US dollar exchange rates and cross (exchange) rates from January 4, 1990 to September 11, 2008, we find that those three variables increase the probability of remaining in the regime where the UIP condition holds. In addition, the probability of switching from the regime where the UIP condition does not hold to the regime where the exchange rate follows the UIP condition decreases (increases) as these variables measuring fear and volatility fall (rise), especially the VIX equity option implied volatility index. For JPYAUS, JPYNZD and USDJPY, the smoothed probabilities show that these exchange rates essentially do not follow the UIP condition except during periods in which the fear and risk variables are increasing, as in the recent global financial crisis in particular.

The remainder of the chapter is set out as follows. In Section 2, we describe

the methodology including linear Fama regression, the Markov-switching model of the UIP condition and the time-varying transition probabilities Markov-switching model of the UIP condition. In Section 3, we describe the data and report the estimation results including summary statistics, results from the linear Fama regression and results from the time-varying transition probabilities Markov-switching model of the UIP condition. This is with a discussion. We make some concluding comments in Section 4.

2 Methodology

We divide this section into three subsections. These are (i) the linear Fama regression, (ii) the Markov-switching model of the uncovered interest parity and (iii) the time-varying transition probabilities Markov-switching model of the uncovered interest parity.

2.1 Linear Fama Regression

Under the efficiency of foreign exchange market where information available to market participants should be fully reflected by prices and under risk neutrality, the UIP condition holds. The parity condition postulates that an expected change in exchange rate is equal to an interest rate differential between the two currencies over some particular time horizon. Formally, this could be written as

$$\Delta_k s^e_{t+k} = i_{t,k} - i^*_{t,k}, \tag{1}$$

where s_t shows the spot exchange rate (defined as the domestic price for one unit of the foreign currency) at time t in the natural logarithmic form. $i_{t,k}$ and $i^*_{t,k}$ denote the nominal interest rates achievable with similar domestic and foreign securities, respectively, with maturity associated with that particular horizon, that is, with k periods to maturity. $\Delta_k s_{t+k} \equiv s_{t+k} - s_t$ and superscript e indicates the expectation of market participants that is formed based upon information available at time t. Under the efficient speculative foreign exchange market, it should not be probable that a market participant would generate excess returns to speculation. Testing the UIP condition within the context of the equation (1) should also be realized as a framework whereby a joint hypothesis of rational expectations and risk neutrality, in an aggregate sense, of foreign exchange market participants could be examined.

In our empirical work, we test the UIP condition in the context of a relationship between spot exchange rates and forward exchange rates.[7] That is, we follow much previous literature on this issue. Essentially, this test we employ is on the reasonably consensual assumption that the covered interest parity (CIP) condition holds.[8] In fact, those empirical analyses of the UIP condition are owed

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[7]The UIP condition testing procedure applied in our empirical work is based on Sarno *et al.* (2006).

[8]When the returns on bonds (and on other debt instruments) are equal, the interest parity exists. More precisely, when the returns on bonds denominated in different currencies are equal, the CIP exists. This is when it is assumed that the forward markets are used to

to the extensive empirical evidence demonstrating and confirming that the CIP condition holds. Note that we would like to examine only the UIP condition. Therefore, for the purposes of our empirical work, it is assumed that the CIP condition exists.[9] If the UIP condition does not hold, then the currency carry trade possibility (which is occasionally, and incorrectly, referred to as uncovered interest arbitrage) would prevail. For the evidence pertaining to the CIP condition, see chapter 2 in Sarno and Taylor (2003) for a survey of evidence.

Following the literature on this issue, we assume that there are no arbitrage opportunities. As a result, the interest rate differential between the two currencies and the forward premium are equal. That is, under the CIP condition, $i_{t,k} - i_{t,k}^* = f_t^k - s_t$, where f_t^k denotes the k-period forward rate, that is, the rate established at this period for currencies being exchanged k periods ahead, in the natural logarithmic form.

Under the CIP condition, replacing the interest rate differential between similar domestic and foreign securities, $i_{t,k} - i_{t,k}^*$, with the forward premium (or the forward discount), $f_t^k - s_t$, equation (1) becomes

$$\Delta s_{t+1} = \alpha + \beta \left(f_t^1 - s_t \right) + v_{t+1}, \tag{2}$$

eliminate the excess rate of return associated with future currency exchanges (i.e. when the bonds mature).

[9]Assuming that the forward rate corresponds to the rate that is provided by the arbitrage inherent in the CIP condition could be another method. It should be noted that the forward rates attained from that calculation should be very accurate, compared with the forward rate data that we have. This is because the CIP condition is actually used for the computation of the forward exchange rate in the foreign exchange market.

where v_{t+1} denotes a disturbance term. Note that it is assumed for simplicity that $k = 1$. This form of regression is referred to as the 'Fama regression' following much previous literature. In addition, this is what a number of researchers have employed and estimated for testing the UIP condition.[10]

Under the UIP condition, the constant term, α, is zero and the estimate of the forward premium coefficient, β, is one. The disturbance term v_{t+1}, that is, the rational expectations forecast error under the null hypothesis, is not correlated with information available at time t. In other words, the UIP condition postulates that the expected change in the exchange rate is equal to the current forward premium. (See Fama (1984)). Note that, in this case, the risk-neutral efficient foreign exchange markets hypothesis holds.

For results from previous empirical analyses based on the Fama regression estimation, the UIP condition is generally rejected. These are examined for different currencies and time periods. In particular, with exchange rates against the US dollar, it constitutes an empirical stylized fact. That is, the forward premium coefficient estimates, β, are usually not statistically significantly different from zero. In addition, they are generally closer to the negative one than to the positive one, which is implied by the UIP condition. (See Froot and Thaler (1990)). As a main empirical finding associated with the forward bias puzzle, the stylized fact of a negative forward premium coefficient estimate in the Fama

[10]Note that the framework described in this subsection stands on the notion that spot exchange rates and forward exchange rates cointegrate. Consequently, the forward premium is stationary. See Brenner and Kroner (1995).

regression indicates that the more the foreign currency is at a premium in the forward market (the low interest rate currency), the less the home currency (the high interest rate currency) is predicted to depreciate. For the evidence pertaining to the UIP condition see, for example, Hodrick (1987), Lewis (1995), Taylor (1995) and Engel (1996), among others.[11]

In addition to the use of the Fama regression, the form presented in the following is also employed in the literature and in our empirical work for testing the UIP condition. That is, the forward premium is used as a predictor variable to derive the following form, which is an investigation of a predictability of the deviations of the UIP condition (or foreign exchange excess returns). Essentially, we reparameterize the Fama regression as in equation (2) in order to obtain the following form of regression in a linear model. This is

$$ER_{t+1} = \alpha + (\beta - 1)\left(f_t^1 - s_t\right) + v_{t+1} = \alpha + \beta^\tau \left(f_t^1 - s_t\right) + v_{t+1}, \qquad (3)$$

where the foreign exchange excess returns $ER_{t+1} \equiv \Delta s_{t+1} - \left(f_t^1 - s_t\right) \equiv s_{t+1} - f_t^1$. This form of regression is investigated in the literature. The estimate of the forward premium coefficient in the current form of regression, β^τ, should be zero under the UIP condition. This is to be consistent with the unity implied

[11]Note that the empirical evidence of the relevant literature is that, over the recent floating exchange rate system, significant time variation of the estimates of the forward premium coefficient in the Fama regression is shown. The range of such estimates across different sample periods is large. In particular, this displays both negative and positive values. See, for example, Baillie and Bollerslev (2000).

by the UIP condition of the estimate of the forward premium coefficient in the Fama regression. Empirical evidence, however, shows the deviation from the UIP condition. That is, on the basis of the lagged forward premium, the generated strong predictability of the deviations of the UIP condition, that is, foreign exchange excess returns, is evident. In particular, with this reparameterized form of regression, the empirical evidence shows that such an estimate of the forward premium coefficient is negative and statistically significantly different from zero. Obviously, owing to the fact that equation (3) is essentially derived from the reparameterization of the Fama regression in the form of the regression equation (2), the statistically significantly negative forward premium coefficient β^{τ} is consistent with a negative estimate of the β in the Fama regression. In other words, the predictability of foreign exchange excess returns that is found in equation (3) could be inferred from the forward bias puzzle which is evident in the Fama regression. For the evidence pertaining to the UIP condition acquired from this reparameterized form of regression, see, for example, Bilson (1981) and Fama (1984), among others.

2.2 Markov-Switching Model of the Uncovered Interest Parity Condition

The Markov-switching model was originally motivated by Goldfeld and Quandt (1973). Nevertheless, it has been popularized by Hamilton (1989). For our

empirical analysis, we would consider it as a switching regression written as (see chapter 22 in Hamilton (1994) for this particular form of the Markov-switching model)

$$y_t = \alpha_i + \beta_i x_t + \epsilon_t, \tag{4}$$

where y_t and x_t are the dependent variable and the exogenous regressor, respectively. They are directly observable. α_i and β_i are the state-depending constant term and coefficient, respectively, to be estimated. Their values depend on a discrete-valued unobserved state variable, S_t. That is, $i = 1$ if $S_t = 1$ and $i = 2$ if $S_t = 2$. ϵ_t is a Gaussian white noise. The state variable S_t is assumed to follow an ergodic first-order Markov process. The transition between states or regimes could be characterized using a transition probability matrix. In our context, the analysis has only two regimes and the transition probability matrix is

$$P = \begin{bmatrix} Pr\,(S_t = 1|S_{t-1} = 1) & Pr\,(S_t = 2|S_{t-1} = 1) \\ Pr\,(S_t = 1|S_{t-1} = 2) & Pr\,(S_t = 2|S_{t-1} = 2) \end{bmatrix} = \begin{bmatrix} p_{11} & p_{12} \\ p_{21} & p_{22} \end{bmatrix}, \tag{5}$$

where p_{ij} (i,j=1,2) indicates the transition probabilities of $S_t = j$ given that $S_{t-1} = i$ and $p_{i1} + p_{i2} = 1$.

For an estimation, there are different approaches to estimating the parameters of the Markov-switching model. These include, for example, maximum

likelihood estimation (MLE), the expectation maximization (EM) algorithm and the Gibbs sampling approach. See Hamilton (1989), Hamilton (1990) and Albert and Chib (1993) for each approach, respectively.[12]

Once the parameters of the model and the transition probability matrix have been estimated, the series of the smoothed probabilities of being in state i based on the knowledge of the complete series, $Pr\left(S_t = i | y_1, ... y_T\right)$, could be computed for each date. See Kim and Nelson (1999) for details of the algorithm. This is in contrast to the calculated series of the filtered probabilities of being in state i, which is based on the information up to date t, $Pr\left(S_t = i | y_1, ... y_t\right)$. It is straightforward to show that, for the last date, that is, $t = T$, the smoothed probability is equal to the filtered probability.

In our context, a Markov-switching model of the UIP condition might be written as

$$\Delta s_{t+1} = \alpha_i + \beta_i \left(f_t^1 - s_t\right) + v_{t+1}, \tag{6}$$

where $v_{t+1} \sim \mathrm{N}\left(0, \sigma_{S_t}^2\right)$. Depending on the realization of a discrete-valued unobserved state variable, S_t, where $S_t \in \{1, 2\}$, the parameter α_i and β_i ($i{=}1$ if $S_t = 1$ and $i{=}2$ if $S_t = 2$) each take on one of the two values. We assume that the state variable, S_t, evolves according to a two-state ergodic first-order Markov process. In addition, we set $\alpha_1 = 0$ and $\beta_1 = 1$. In other words, it is assumed further that the first state corresponds to the regime where the UIP

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[12]See also Krolzig (1997) and Kim and Nelson (1999).

condition holds, that is, the constant term, α, is equal to zero and the forward premium coefficient, β, is equal to one in the Fama regression as in equation (2). The second state corresponds to the Fama regression where the UIP condition might not exactly hold.

Note that we allow for shifts in the (weekly) volatility parameter across regimes. This is due to commonly found evidence of autoregressive conditional heteroskedasticity (ARCH) effects or fat tails in the distribution of innovations to nominal exchange rates at relatively high frequencies, say, the weekly frequency. Consequently, this characteristic should be modeled in a Markov-switching framework.[13]

2.3 Time-Varying Transition Probabilities Markov-Switching Model of the Uncovered Interest Parity Condition

The time-varying transition probabilities approach is in contrast to the constant transition probabilities approach described in the previous subsection. The latter means that the probability of switching from one regime to the other depends neither on time nor on the other variables indicating the state of the economy. The time-varying transition probabilities approach was originally developed by Filardo (1994). This is in the context of business cycle research. Specifically, Fi-

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[13]We perform formal tests for ARCH effects with our exchange rate data. The results suggest that we reject the null hypothesis of homoskedasticity at the standard significance level, either in terms of the changes in the exchange rate or in terms of the residuals from the estimated Fama regression, for each exchange rate examined.

lardo (1994) shows that the time-varying transition probabilities of the Markov process are significantly determined by business cycle indicator variables that help to infer the switching points between different stages.

Within the context of a time-varying transition probabilities Markov-switching model of the UIP condition, a key component of the approach taken here is that the probabilities of switching from one regime to another are modeled endogenously. In particular, we model the state transition probabilities as a function of (i) exchange rate volatility (VOL), (ii) the VIX equity option implied volatility index (VIX) and (iii) the TED spread (TED).

With the time-varying transition probabilities approach in our context, we could test if three such variables, believed to have an effect on the state transition probabilities, actually influence the probability of switching from one regime to the other. We begin by explaining each of these three variables.

Firstly, the exchange rate volatility is caused by the unpredictable fluctuations in the exchange rates. It could be thought of as a measure aiming to capture the uncertainty faced by market participants. Certainly, this is an unobservable variable. In literature, the use of Bollerslev's (1986) generalized autoregressive conditional heteroskedasticity (GARCH) model has been increasingly adopted for obtaining the exchange rate volatility. This is the generalization of the autoregressive conditional heteroskedasticity (ARCH) model proposed by Engle (1982). In our context, we follow this. Specifically, we use the measure derived from the GARCH(*1,1*) model as the measure of exchange rate volatility. That

is, the conditional variance of the first difference of the logarithm of the exchange rate.

For the GARCH model, suppose that the exchange rate return (Δs_t) is governed by the following autoregressive specification

$$\Delta s_t = \alpha_0 + \sum_{i=1}^{p} \beta_i \Delta s_{t-i} + u_t, \tag{7}$$

where α_0 shows a constant term, β_i's are coefficient terms and the error term, u_t, is normally distributed with mean zero and variance σ^2, $u_t|\Omega_{t-1} \sim N(0, \sigma^2)$, where Ω_{t-1} denotes the available information set at time t-1. Bollerslev (1986) allows for the variance to vary over time and for persistence in volatility with a relatively small number of parameters to be estimated. In particular, the change of variance over time is assumed to be characterized by the following GARCH(p,q) specification, becoming a function not only of the squared residuals, u_{t-i}^2, as in the ARCH model, but also of the lagged values of itself. That is,

$$\sigma_t^2 = \lambda_0 + \sum_{i=1}^{p} \phi_i u_{t-i}^2 + \sum_{i=1}^{q} \delta_i \sigma_{t-i}^2, \tag{8}$$

where σ_t^2 is the conditional variance of the logarithm of the exchange rate, u_{t-i}^2 represents the squared residual and σ_{t-i}^2 is the lagged value of itself. ϕ_i's and δ_i's are parameters to be estimated. All the estimates of those coefficients need to be positive. This is in order to assure that exchange rate volatility is positive.

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The most common form is the GARCH($1,1$). This could be represented as

$$\sigma_t^2 = \lambda_0 + \phi_1 u_{t-1}^2 + \delta_1 \sigma_{t-1}^2. \tag{9}$$

This is the specification we use to measure exchange rate volatility.[14]

Secondly, the VIX equity option implied volatility index, the ticker symbol for the Chicago Board Options Exchange (CBOE) volatility index, could be thought of as a measure of global market risk or risk aversion. In particular, it is a key measure of market expectations of near-term (30-day) volatility, that is, a forward looking volatility. The VIX equity option implied volatility index is constructed using the implied volatilities of S&P 500 index option prices, calculated from both calls and puts. The VIX equity option implied volatility index traded at the CBOE has been regarded by market participants as the world's premier measure of the sentiment of investors and volatility of financial markets since its introduction in 1993. Often, such a measure is referred to as the 'investor fear gauge'. Note that due to the fact that this investor fear gauge is calculated from equity options, it is not directly related to the exchange rates.

Finally, the TED spread is the difference between the LIBOR interbank market interest rate and the risk-free T-bill rate. The TED spread is referred to as indicating a perceived credit risk in the general economy. This is due to the fact that T-bills are considered risk-free, whereas LIBOR shows the credit

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[14]Note that estimation of the ARCH-class model would be appropriate if the mean equation exhibits evidence of ARCH effects.

risk that is associated with commercial bank lending. When an increase in the TED spread is realized, it would be interpreted that the default risk on interbank loans, which is also the counterparty risk, is perceived by the lenders to be increasing. Consequently, lenders in the interbank market would require a higher interest rate. In addition, in this situation, for a lower rate of return but safe investments, they might consider T-bills, for example. In the opposite situation, when it is perceived that the bank default risk is to decrease, the TED spread would also decrease.

We decided on the variables for which we would test for an effect on the transition probabilities. Following Diebold *et al.* (1994), Durland and McCurdy (1994) and Filardo (1994), a logistic function is employed in our empirical work. This is in order to ensure that the probabilities lie in the unit interval between 0 and 1. Thus, if we denote the transition probability of switching from regime i to regime i (remaining in regime i) at time t as p_t^{ii} for $i \in \{1, 2\}$, we could write the postulated function, the time-varying transition probabilities, as

$$
\begin{aligned}
p_t^{ii} &\equiv Pr\left[S_t = i | S_{t-1} = i, \ VOL_{t-1}, \ VIX_{t-1}, \ TED_{t-1}\right] \equiv \\
&\equiv \frac{exp\left[\beta_{i0} + \beta_{i1}VOL_{t-1} + \beta_{i2}VIX_{t-1} + \beta_{i3}TED_{t-1}\right]}{1 + exp\left[\beta_{i0} + \beta_{i1}VOL_{t-1} + \beta_{i2}VIX_{t-1} + \beta_{i3}TED_{t-1}\right]} \ ,
\end{aligned}
\tag{10}
$$

where β_{i0}, β_{i1}, β_{i2} and β_{i3} denote unknown parameters. Note that the exchange rate volatility, the VIX equity option implied volatility index and the TED

spread are lagged one time period in the time-varying transition probabilities. While this defines p_t^{11} and p_t^{22}, clearly we also have the implicit definitions $p_t^{12} \equiv 1 - p_t^{11}$ and $p_t^{21} \equiv 1 - p_t^{22}$. The transition between states or regimes at time t could be represented using a transition probability matrix at time t. In our context, the analysis has two regimes and the transition probability matrix at time t is

$$
P_t = \begin{bmatrix} \frac{exp[\beta_{10}+\beta_{11}VOL_{t-1}+\beta_{12}VIX_{t-1}+\beta_{13}TED_{t-1}]}{1+exp[\beta_{10}+\beta_{11}VOL_{t-1}+\beta_{12}VIX_{t-1}+\beta_{13}TED_{t-1}]} & 1 - \frac{exp[\beta_{10}+\beta_{11}VOL_{t-1}+\beta_{12}VIX_{t-1}+\beta_{13}TED_{t-1}]}{1+exp[\beta_{10}+\beta_{11}VOL_{t-1}+\beta_{12}VIX_{t-1}+\beta_{13}TED_{t-1}]} \\ \\ 1 - \frac{exp[\beta_{20}+\beta_{21}VOL_{t-1}+\beta_{22}VIX_{t-1}+\beta_{23}TED_{t-1}]}{1+exp[\beta_{20}+\beta_{21}VOL_{t-1}+\beta_{22}VIX_{t-1}+\beta_{23}TED_{t-1}]} & \frac{exp[\beta_{20}+\beta_{21}VOL_{t-1}+\beta_{22}VIX_{t-1}+\beta_{23}TED_{t-1}]}{1+exp[\beta_{20}+\beta_{21}VOL_{t-1}+\beta_{22}VIX_{t-1}+\beta_{23}TED_{t-1}]} \end{bmatrix}.
$$

$$(11)$$

Instead of constant probabilities, p_{11} and p_{22} (and the corresponding p_{12} and p_{21}), the current approach gives estimates of the coefficients β_{11}, β_{12}, β_{13}, β_{21}, β_{22} and β_{23}. From the assumed functional form of the time-varying transition probabilities given above, one could then infer the p_t^{11} and the p_t^{22} (and the corresponding p_t^{12} and p_t^{21}) series. In addition, one could use an algorithm developed by Kim (1994) to estimate the smoothed probability of being in each of the two regimes over time, using all the information in the sample, i.e. $Pr\left[S_t = i | I_T\right]$, where I_T is the information set that contains the sample histories of all the variables. This is after estimating the model and generating $Pr\left[S_t = i | I_t\right]$, which is based on the information up to date t.

3 Data and Results

We divide this section into four subsections. These are (i) data, (ii) summary statistics, (iii) results from the linear Fama regression and (iv) results from the time-varying transition probabilities Markov-switching model of the UIP condition.

3.1 Data

Our data set for the following empirical exercise is composed of observations, at weekly frequency, of spot (the domestic price for one unit of foreign currency) and four-week (or one-month) forward (foreign) US dollar exchange rates against (home) the Japanese yen, Australian dollar, New Zealand dollar and Swiss franc. It also comprises spot Japanese yen exchange rates against the Australian dollar and spot New Zealand dollar exchange rates against the Japanese yen.[15] The data are taken from *Datastream* (Federal Reserve Bank of New York, Reserve Bank of Australia, Reserve Bank of New Zealand and Barclays Bank Plc).

For VIX equity option implied volatility index historical data, price history data are taken from the Chicago Board Options Exchange (CBOE). Note that, after September 22, 2003, the volatility index prices using the new methodology would be stated as 'VIX' and the volatility index prices using the old methodology would be stated as 'VXO'.

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[15]Because of the very high interest rate spreads, these two currency pairs are the most popular pairs for the currency carry trade and, therefore, chosen.

For the TED spread, the weekly data for the three-month LIBOR interbank market interest rate and the three-month US risk-free T-bill interest rate are from Bloomberg. The TED spread is calculated as the difference between the three-month LIBOR and the three-month T-bill interest rates.

Because of an availability of VIX equity option implied volatility index data, the sample period for our empirical work spans from January 4, 1990 to September 11, 2008. This is for all the exchange rates examined. There are, therefore, 975 observations for each exchange rate examined.

Note that, in the following, a change from t to $t + 1$ corresponds to a four-week change in a variable for the notation to be kept simple. Consequently, f_t^j shows the forward exchange rate for a contract with a j-month maturity. From this data set, the time series of interest are constructed. These are the spot exchange rate and the one-month forward exchange rate in logarithmic form, s_t and f_t^1, respectively. Figure 1 shows the VIX equity option implied volatility index and the TED spread. All of these are at the weekly frequency.

3.2 Summary Statistics

We construct the depreciation rate, $s_{t+1} - s_t$, the forward premium, $f_t^1 - s_t$, and the return from currency speculation (excess return), $s_{t+1} - f_t^1$. This is from the weekly spot exchange rates and four-week (or one-month) forward exchange rates. Sample moments for the depreciation rate, the forward premium and the return from currency speculation (excess return) are reported in Table 1. Note

that the figures in parentheses are standard errors which are calculated by using an autocorrelation and heteroskedasticity consistent matrix of residuals, with three lags.[16]

From the summary statistics provided in Table 1, the stylized fact that the depreciation rate and the forward premium own a mean near zero and a large standard deviation (for the depreciation rate) is shown. See also Sarno *et al.* (2006) for similar summary statistics. The first-order autocorrelation coefficient of the depreciation rate is very small in size (between the range of 0.0003 for JPYNZD and 0.0642 for USDSWI) and generally statistically insignificantly different from zero (with p-values between the range of 0.1386 for USDSWI and 0.9962 for JPYNZD). For the forward premium, the first-order autocorrelation coefficient is, however, found to be larger (between the range of 0.0728 for JPYAUS and 0.6688 for USDNZD) and generally statistically significantly different from zero (with p-values in the range of 0 to 0.0168). These results are also in line with the stylized facts that the depreciation rate is not a persistent process, showing weak serial correlation, while the forward premium is, on the contrary, a highly persistent process. Due to the evidence that the depreciation rate shows weak serial correlation, this might imply a near random walk behavior of the exchange rate. From the summary statistics provided, it is confirmed that the mean and standard deviation of the return from currency speculation (excess return) are similar to those of the depreciation rate. The first-order au-

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[16]See Newey and West (1987).

tocorrelation coefficient of this return from currency speculation (excess return) is small (between the range of 0.0101 for JPYNZD and 0.0785 for USDSWI).

We also test for the unit root behavior of the spot exchange rate and the forward exchange rate time series examined by calculating several unit root test statistics. The results are reported in Table 2. For each exchange rate examined, we are unable to reject the unit root null hypothesis for both spot and forward exchange rates. However, differencing the spot exchange rate and the forward exchange rate time series does appear to induce stationarity, for each exchange rate examined. The forward premium is found to be stationary at conventional nominal significance levels. Therefore, the unit root tests clearly indicate that, for each exchange rate examined, the spot exchange rate and the forward exchange rate time series are a realization from a stochastic process integrated of order one, $I(1)$. The forward premium is, on the contrary, stationary. This is in line with the empirical evidence that spot exchange rates and forward exchange rates are cointegrated.[17]

3.3 Results from the Linear Fama Regression

Before performing the Markov-switching model of the UIP condition, we estimate the conventional (linear) Fama regression as in equation (2). This is for each exchange rate examined (four (foreign) US dollar exchange rates against (home) the Japanese yen, Australian dollar, New Zealand dollar and Swiss franc

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[17]See Brenner and Kroner (1995).

and two cross (exchange) rates, Japanese yen exchange rates against the Australian dollar and New Zealand dollar exchange rates against the Japanese yen). Table 3 reports the results. These results are consistent with the existence of the forward bias. In particular, for four US dollar exchange rates (USDJPY, USDAUS, USDNZD and USDSWI), the constant term, α, is very close to zero and often statistically insignificant. The forward premium coefficient, β, is estimated to be negative for USDJPY, USDNZD and USDSWI and is statistically insignificantly different from zero for USDNZD and USDSWI. A rejection of the UIP condition is implied by the negative forward premium coefficient usually reported in such regressions. This, combined with the negative constant term, suggests that currencies at a forward discount actually appreciate rather than depreciate as suggested by the UIP condition. Because a currency at a forward discount is the currency with high interest rates, the regression results reported appear to imply that market participants could expect the high interest rate currency to appreciate. This is contrary to the UIP condition hypothesis. Furthermore, the regression estimates seem to imply that the use of currency carry trade aimed at exploiting the forward bias would be profitable. This is because it pays to borrow in low interest rate currencies that appear to depreciate and, instead, invest in currencies with high interest rates that appear to appreciate. However, for one of the US dollar exchange rate regressions estimated, USDAUS is the exception. In particular, the estimate of the forward premium coefficient is positive (about 0.61) and statistically significant. Nevertheless, this estimate

is not exactly the theoretical value of unity that is implied by the UIP condition. In other words, this indicates that the currency with higher interest rates tends not to depreciate as much as suggested by the UIP condition.

To confirm a departure from market efficiency, we use the forward premium as a predictor variable to derive the predictability regression to be estimated. In particular, this linear model obtained from reparameterizing the Fama regression is the predictability of the deviations of the UIP condition (or foreign exchange excess returns). This is as in equation (3).

Corresponding with Bilson (1981) and Fama (1984), among others, we find strong evidence of deviations from the UIP condition for all the US dollar exchange rates examined. In other words, on the basis of the lagged forward premium, the predictability of foreign exchange excess returns is found. In particular, while β^τ should be zero under the UIP condition, the results shown in Table 3 are that the estimate of the forward premium coefficient in this reparameterized Fama regression, β^τ, is negative and statistically significantly different from zero.[18] This is consistent with a negative estimate of the forward premium coefficient, β, in the Fama regression. Hence, the departure from market efficiency (under which the estimate of the forward premium coefficient in this reparameterized Fama regression, β^τ, is zero) is confirmed. The forward premium, which is a component of the information set of the market participants, could be used for a prediction of foreign exchange excess returns.

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[18]This is at the 10 percent significance level for USDAUS and USDNZD.

For cross (exchange) rates (JPYAUS and JPYNZD), the estimates of the forward premium coefficient are 0.27 and -0.71 for JPYAUS and JPYNZD, respectively. They are not statistically significant. For JPYNZD, a rejection of the UIP condition is implied by that negative forward premium coefficient. For JPYAUS, the estimate of the forward premium coefficient, β, is much smaller than the theoretical value of unity that is suggested by the UIP condition. In addition, when examining the standard errors, that theoretical value implied by the UIP condition is not included in this estimate. This implies that the currency with a higher interest rate tends not to fall in value as much as predicted by the UIP condition.

3.4 Results from the Time-Varying Transition Probabilities Markov-Switching Model of the Uncovered Interest Parity Condition

According to equation (11), for the exchange rates examined, the UIP condition is endogenously captured by the estimates of the parameters β_{11} and β_{21}, β_{12} and β_{22} and β_{13} and β_{23}, which measure the effects of exchange rate volatility, the VIX equity option implied volatility index and the TED spread on the transition probabilities, respectively.[19] Figure 1 shows the VIX equity option implied volatility index and the TED spread and Figure 2 shows the exchange

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[19]The estimated GARCH equations are not reported since we only use the conditional variance acquired as a measure of exchange rate volatility.

rate volatility.

Prior to obtaining results from the time-varying transition probabilities Markov-switching model of the UIP condition, we conduct (i) a test of the null hypothesis of a linear specification versus a non-linear specification (see Garcia (1998)) and (ii) a simple likelihood ratio test for the time-varying transition probabilities model under the null hypothesis of no time variation.

For the former, we find that the non-linear (two-regime) model statistically significantly fits the data better than a linear model at the 5 percent significance level, for each exchange rate examined.[20] We do not examine a regime specification higher than two. This is because, in addition to challenging computation, it might be qualitatively difficult to describe such an additional regime. Also, among the existing literature on the Markov-switching model of the UIP condition, a regime specification higher than two is rare.

For the latter, that is, the test for the time-varying transition probabilities model under the null hypothesis of no time variation, we compute the likelihood ratio statistic associated with comparing the constant transition probability model as the restricted model and the time-varying transition probabilities model as the unrestricted model. We examine whether such a likelihood ratio

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[20]The values of the likelihood ratio test statistic are 217.276, 180.822, 165.532, 144.271, 224.434 and 33.422 for JPYAUS, JPYNZD, USDJPY, USDAUS, USDNZD and USDSWI, respectively. The critical values are 14.11 for the 5 percent significance level and 12.23 for the 10 percent significance level. Note that this test should be treated with caution, however, since the critical values of the test statistic derived by Garcia (1998) are for a model including no explanatory variables apart from a constant. Nevertheless, given the high value of the likelihood ratio test statistic, our conclusion might be made safely.

statistic is statistically significant, for each exchange rate examined. The likelihood ratio statistic approximately follows a chi-squared distribution and the degree of freedom is equal to the number of additional parameters in a transition probability matrix. We conclude that the time-varying transition probabilities model is superior to the constant transition probability model. This proves to be statistically significant at the 5 percent significance level, for all the exchange rates examined except JPYAUS.[21]

Estimation of the model is carried out using the maximum likelihood (ML) method. Parameter estimates for the Markov-switching model of the UIP condition and for the time-varying transition probabilities, estimated standard errors and significance are given in Table 4. Note that an estimate of the parameters β_{11} and β_{21}, β_{12} and β_{22} and β_{13} and β_{23} is subsequently set to zero if it is found to be insignificantly different from zero at the 10 percent significance level in the initial estimations. The statistically significant values of the estimates of the parameters β_{11} and β_{21}, β_{12} and β_{22} and β_{13} and β_{23} are explained in the following.

The estimate of the forward premium coefficient in the regime where the UIP condition does not hold, i.e. β_2, is negative or much smaller than the theoretical value of unity that is implied by the UIP condition.[22] Note that it is smaller

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[21]The significance levels for the likelihood ratio test for the time-varying transition probabilities model are 0.743, 0.033, 0.000, 0.009, 0.006 and 0.000, for JPYAUS, JPYNZD, USDJPY, USDAUS, USDNZD and USDSWI, respectively.

[22]This implies that the currency with a higher interest rate tends not to fall in value as much as suggested by the UIP condition.

than the one acquired in the linear Fama regression. However, for USDAUS, it is similar to the one in the linear Fama regression.

For the effects of exchange rate volatility on the transition probabilities, the negative and statistically significant value of the estimate of the parameter β_{21} has the effect that, with an increase in the exchange rate volatility, the probability of remaining in the regime where the UIP condition does not hold, p_t^{22}, that is, the constant term, α, is not equal to zero and the forward premium coefficient, β, is not equal to one in the Fama regression as in equation (2), is small. USDJPY and USDNZD show this, with the value of the coefficient -0.328 and -0.86, respectively. In addition, it is indicative that a rise in the exchange rate volatility tends to result in a fall in p_t^{22}. In other words, since $p_t^{21} \equiv 1 - p_t^{22}$, the probability of switching from the regime where the UIP condition does not hold to the regime where the exchange rate follows the UIP condition is large. It decreases as the exchange rate volatility falls. For USDAUS, USDNZD and USDSWI, the positive and statistically significant value of the estimate of the parameter β_{11} has the effect that an increase in the exchange rate volatility tends to result in a rise in the probability of remaining in the regime where the UIP condition holds. Therefore, we argue that low exchange rate volatility and the failure of the UIP condition are linked. This is due to the fact that the model is able to provide the evidence that, once the exchange rate embarks on a UIP condition path, it is likely to continue to follow the parity postulate when the exchange rate volatility is large. Note that the magnitude of the effect on

the probability of remaining in the first regime seems to be larger than that of remaining in the second regime. That is, the strong effects of volatility in the foreign exchange markets would be expected to be with respect to the probability of remaining in the regime where the UIP condition holds rather than remaining in the regime where the exchange rate does not follow the UIP condition.

For the effects of the VIX equity option implied volatility index on the transition probabilities, we look at the estimates of the parameters β_{12} and β_{22}. The positive and statistically significant value of the estimate of the parameter β_{12} indicates positive VIX equity option implied volatility index-UIP condition dependence. That is, the larger the VIX equity option implied volatility index is in the (financial) markets, the higher is the probability of its remaining in the regime where the UIP condition holds. USDJPY shows this, with the value of the coefficient 0.318. On the other hand, the negative and statistically significant value of the estimate of the parameter β_{22} has the effect that the probability of switching from the regime where the UIP condition does not hold to the regime where the exchange rate follows the UIP condition is large, with an increase in the VIX equity option implied volatility index. For the exchange rates examined, JPYAUS, USDJPY and USDSWI show this, with the value of the coefficient -0.506, -0.590 and -2.273, respectively. Therefore, we argue that the model is able to capture the notion that fear in the financial markets captured by the fear index (VIX) and the failure of the UIP condition are linked. Note that, for USDJPY, the effect of the VIX equity option implied volatility

index on the transition probabilities is stronger than that of the exchange rate volatility.

For the effects of the TED spread on the transition probabilities, we analyze the effects captured by the estimates of the parameters β_{13} and β_{23}. The point estimate for the parameter β_{13} for JPYNZD is positive, with the value of the coefficient 0.672, and proves to be statistically significant. This implies that the TED spread tends to decrease the probability of switching into the regime where the UIP condition does not hold when the exchange rate is in the regime where it follows the UIP condition. In other words, under certain conditions, the increase in the TED spread actually increases the probability of remaining in the regime where the UIP condition holds over the next period. Nevertheless, the statistically significant effect of the TED spread does not seem to be found in our context, for other exchange rates. Following Brunnermeier *et al.* (2008), it could be argued that an increase in the TED spread has similar, but less statistical, realized effects with an increase in the VIX equity option implied volatility index.

From these results, we argue that exchange rate volatility itself and the VIX equity option implied volatility index are important in statistically capturing the effect on the probability of remaining in and switching into the regime where the UIP condition holds. This is due to the fact that, with these two variables, the significant effect on the transition probabilities prevails in at least one of

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the two regimes, or in both regimes for USDJPY, USDNZD and USDSWI.[23] In addition, the probability of switching from the regime where the UIP condition does not hold to the regime where the exchange rate follows the UIP condition decreases (increases) as these variables measuring fear and volatility fall (rise), especially the VIX equity option implied volatility index.

We also look at the evidence of the estimate of the (weekly) volatility parameter of the regime where the UIP condition holds and the regime where the exchange rate does not follow the UIP condition. The average (weekly) volatility parameter is 0.04 and 0.02 for the first regime, σ_1, and for the second regime, σ_2, respectively. They are statistically significantly different from zero. We find that, for all the exchange rates examined, the volatility parameter of the regime where the UIP condition holds is higher than the volatility estimate of the regime where the exchange rate does not follow the UIP condition. This is in line with the existing literature that the movements of exchange rates tend to agree with the UIP condition during time with high exchange rate volatility.[24]

Corresponding to the estimates in Table 4, we calculate the weekly estimates for the smoothed probabilities of being in each of the two regimes over time. That is, the probabilities constructed using the whole sample data, $Pr(S_t = 1|I_T)$ and $Pr(S_t = 2|I_T)$, where I_T is the information set that contains the sample histories of all the variables, the depreciation rate, the forward premium, the exchange rate volatility, the VIX equity option implied volatility index and the TED

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[23]JPYNZD is an exception.
[24]Huisman and Mahieu (2007) report similar finding.

spread. Illustrating the ability of the model to identify long-term movement in the exchange rate, the smoothed probabilities provide us with more insight into the differences between the periods in which exchange rates follow the UIP condition and periods in which they do not. We show the estimated probabilities for the exchange rate being in the regime where the UIP condition holds and in the regime where it does not follow the UIP condition in Figures 3 to 8, for each exchange rate examined. For JPYAUS, JPYNZD and USDJPY, the smoothed probabilities show that these exchange rates essentially do not follow the UIP condition except during periods in which the fear and risk variables are increasing, as in the recent global financial crisis in particular.

4 Conclusion

This chapter is motivated by the fact that there should be a close link between fear, volatility and the performance of the currency carry trade associated with the failure of the UIP condition. Higher fluctuations in the financial markets, including prices of not only exchange rates but also stocks and bonds, could worsen excess returns that could be generated through the currency carry trade activities. This is due to the fact that many investors analyze performance, not only by returns, but also by the variability of those returns. That is, they measure their investment performance by the Sharpe ratio in particular. An increase in volatility with no change in return would decrease the Sharpe ratio

and a lower Sharpe ratio is less favorable. Consequently, it might be difficult for the foreign exchange carry trade activities to sustain the type of performance it has had if the environment with low volatility does not persist.

Within the context of a time-varying transition probabilities Markov-switching model of the UIP condition, we examine if variables measuring fear and volatility have an effect on the probability of switching between the regime where the UIP condition holds and the regime where it does not. The state transition probability depends nonlinearly upon the variables examined. These are the exchange rate volatility, the VIX equity option implied volatility index and the TED spread.

Applying this to both US dollar exchange rates and cross (exchange) rates from January 4, 1990 to September 11, 2008, we find that those three variables increase the probability of remaining in the regime where the UIP condition holds. In addition, the probability of switching from the regime where the UIP condition does not hold to the regime where the exchange rate follows the UIP condition decreases (increases) as these variables measuring fear and volatility fall (rise), especially the VIX equity option implied volatility index. For JPYAUS, JPYNZD and USDJPY, the smoothed probabilities show that these exchange rates essentially do not follow the UIP condition except during periods in which the fear and risk variables are increasing, as in the recent global financial crisis in particular.

Figure 1: The VIX Equity Option Implied Volatility Index and the TED Spread
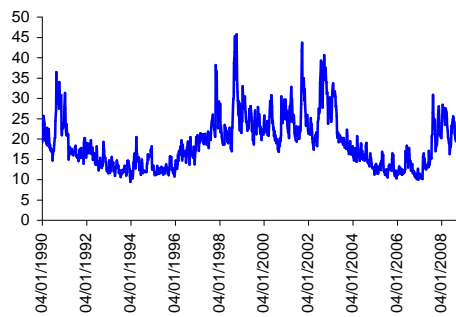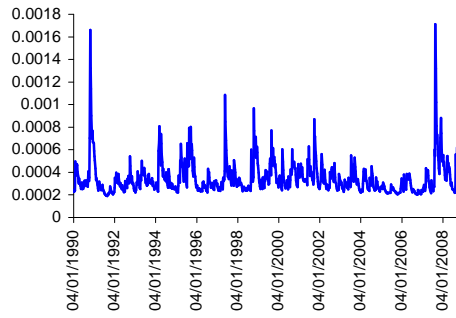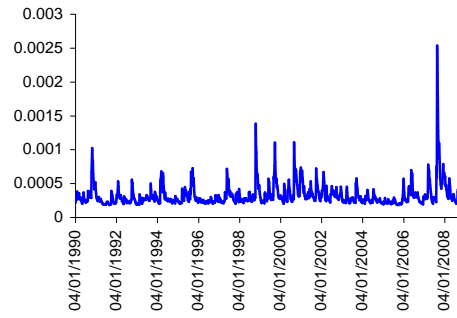

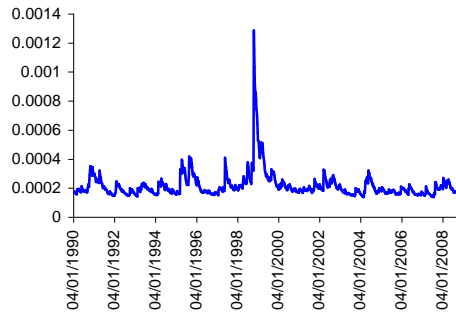(VIX Equity Option Implied Volatility Index)

(TED Spread)

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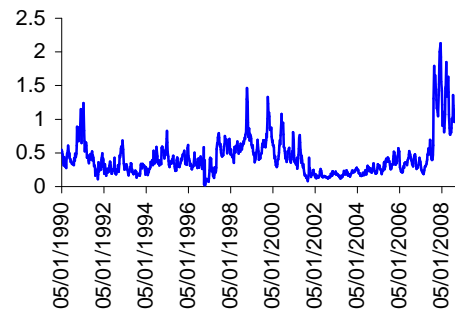
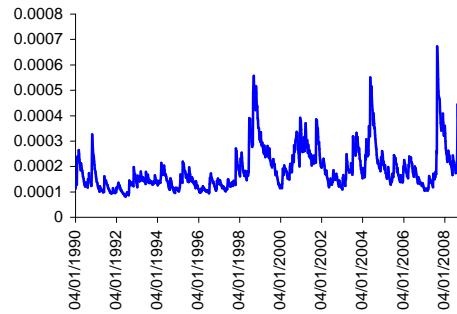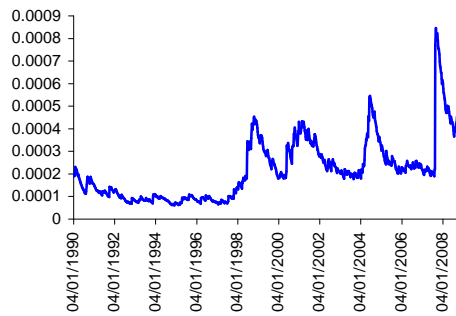## Figure 2: Exchange Rate Volatility

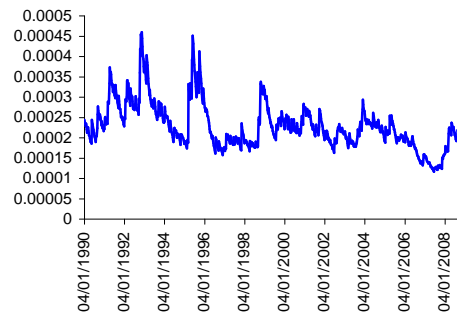(JPYAUS)

(JPYNZD)

(USDJPY)

(USDAUS)

(USDNZD)

(USDSWI)

Table 1: Sample Moments for the Depreciation Rate, the Forward Premium and the Return from Currency Speculation (Excess Return)

Currency	Sample moment				
	Depreciation rate, $s_{t+1} - s_t$				
	Mean	Standard deviation	AR(1)	Standard errors	P-value
JPYAUS	0.0008	0.0365	0.0474	(0.0574)	0.4087
JPYNZD	0.0006	0.0356	0.0003	(0.0546)	0.9962
USDJPY	-0.0012	0.0301	0.0562	(0.0564)	0.3190
USDAUS	-0.0004	0.0273	0.0161	(0.0527)	0.7600
USDNZD	-0.0006	0.0279	0.0324	(0.0557)	0.5604
USDSWI	-0.0013	0.0305	0.0642	(0.0433)	0.1386
	Forward premium, $f_t^1 - s_t$				
	Mean	Standard deviation	AR(1)	Standard errors	P-value
JPYAUS	0.0041	0.0053	0.0728	(0.0304)	0.0168
JPYNZD	0.0050	0.0026	0.3325	(0.0534)	0.0000
USDJPY	-0.0027	0.0028	0.4881	(0.0490)	0.0000
USDAUS	0.0014	0.0050	0.1200	(0.0323)	0.0002
USDNZD	0.0023	0.0017	0.6688	(0.0862)	0.0000
USDSWI	-0.0014	0.0028	0.5182	(0.0397)	0.0000
	Return from currency speculation (excess return), $s_{t+1} - f_t^1$				
	Mean	Standard deviation	AR(1)	Standard errors	P-value
JPYAUS	-0.0033	0.0367	0.0506	(0.0581)	0.3834
JPYNZD	-0.0044	0.0358	0.0101	(0.0548)	0.8543
USDJPY	0.0015	0.0305	0.0746	(0.0569)	0.1898
USDAUS	-0.0018	0.0272	0.0173	(0.0544)	0.7501
USDNZD	-0.0029	0.0280	0.0415	(0.0557)	0.4564
USDSWI	0.0001	0.0307	0.0785	(0.0435)	0.0710

Table 2: Unit Root Behavior Test

Currency	Spot	ΔSpot	Forward	ΔForward	Forward Premium
JPYAUS	0.2068	0.0000*	0.1918	0.0000*	0.0000*
JPYNZD	0.3178	0.0000*	0.3071	0.0000*	0.0337*
USDJPY	0.3679	0.0000*	0.3620	0.0000*	0.0000*
USDAUS	0.5676	0.0000*	0.5138	0.0000*	0.0000*
USDNZD	0.6322	0.0000*	0.6249	0.0000*	0.0000*
USDSWI	0.2034	0.0000*	0.2011	0.0000*	0.0152*

Note: * indicates statistical significance at 5 percent significance level.

Table 3: Results from the Linear Fama Regression

Currency	Parameter estimates							
	α	SE(α)	P-value(α)	β	SE(β)	P-value(β)	SD(v_{t+1})	T
JPYAUS	-0.000	0.002	0.894	0.265	0.233	0.256	0.036	971
	$\beta - 1$	β^τ	SE(β^τ)	t(β^τ)	P-value(β^τ)			
	-0.735	-0.735	0.233	-3.159	0.002			
	α	SE(α)	P-value(α)	β	SE(β)	P-value(β)	SD(v_{t+1})	T
JPYNZD	0.004	0.003	0.170	-0.714	0.491	0.146	0.036	971
	$\beta - 1$	β^τ	SE(β^τ)	t(β^τ)	P-value(β^τ)			
	-1.714	-1.714	0.491	-3.492	0.000			
	α	SE(α)	P-value(α)	β	SE(β)	P-value(β)	SD(v_{t+1})	T
USDJPY	-0.004	0.002	0.052	-0.953	0.460	0.039	0.030	971
	$\beta - 1$	β^τ	SE(β^τ)	t(β^τ)	P-value(β^τ)			
	-1.953	-1.953	0.460	-4.241	0.000			
	α	SE(α)	P-value(α)	β	SE(β)	P-value(β)	SD(v_{t+1})	T
USDAUS	-0.001	0.001	0.399	0.614	0.209	0.003	0.027	971
	$\beta - 1$	β^τ	SE(β^τ)	t(β^τ)	P-value(β^τ)			
	-0.386	-0.386	0.209	-1.848	0.065			
	α	SE(α)	P-value(α)	β	SE(β)	P-value(β)	SD(v_{t+1})	T
USDNZD	0.001	0.003	0.792	-0.552	0.846	0.514	0.028	971
	$\beta - 1$	β^τ	SE(β^τ)	t(β^τ)	P-value(β^τ)			
	-1.552	-1.552	0.846	-1.834	0.067			
	α	SE(α)	P-value(α)	β	SE(β)	P-value(β)	SD(v_{t+1})	T
USDSWI	-0.002	0.002	0.295	-0.479	0.515	0.353	0.030	971
	$\beta - 1$	β^τ	SE(β^τ)	t(β^τ)	P-value(β^τ)			
	-1.479	-1.479	0.515	-2.869	0.004			

Table 4: Parameter Estimates for the Time-Varying Transition Probabilities and for the Markov-Switching Model of the Uncovered Interest Parity, Estimated Standard Errors and Significance

Currency	Parameter Estimates							
	β_{10}	β_{11}	β_{12}	β_{13}	β_{20}	β_{21}	β_{22}	β_{23}
JPYAUS	5.069				4.306*		**-0.506***	
(Std Error)	5.798				0.367		0.195	
Signif	0.382				0.000		0.009	
	α_2	β_2	σ_1	σ_2				
	-0.009*	0.168	0.049*	0.022*				
(Std Error)	0.004	0.473	0.008	0.003				
Signif	0.013	0.723	0.000	0.000				
	β_{10}	β_{11}	β_{12}	β_{13}	β_{20}	β_{21}	β_{22}	β_{23}
JPYNZD	2.633*			**0.672****	2.828*			
(Std Error)	0.467			0.368	0.299			
Signif	0.000			0.068	0.000			
	α_2	β_2	σ_1	σ_2				
	-0.002	-1.061*	0.050*	0.023*				
(Std Error)	0.003	0.430	0.003	0.001				
Signif	0.445	0.014	0.000	0.000				
	β_{10}	β_{11}	β_{12}	β_{13}	β_{20}	β_{21}	β_{22}	β_{23}
USDJPY	0.895*		**0.318***		1.738*	**-0.328***	-0.590*	
(Std Error)	0.383		0.135		0.525	0.057	0.224	
Signif	0.019		0.018		0.001	0.000	0.009	
	α_2	β_2	σ_1	σ_2				
	0	-1.463*	0.045*	0.020*				
(Std Error)	0.001	0.280	0.003	0.001				
Signif	0.772	0.000	0.000	0.000				
	β_{10}	β_{11}	β_{12}	β_{13}	β_{20}	β_{21}	β_{22}	β_{23}
USDAUS	20.604*	**2.085***			2.566*			
(Std Error)	6.760	0.774			0.289			
Signif	0.002	0.007			0.000			
	α_2	β_2	σ_1	σ_2				
	-0.006*	0.603*	0.035*	0.015*				
(Std Error)	0.002	0.228	0.002	0.001				
Signif	0.001	0.008	0.000	0.000				
	β_{10}	β_{11}	β_{12}	β_{13}	β_{20}	β_{21}	β_{22}	β_{23}
USDNZD	9.867*	**0.870***			-4.849	**-0.860***		
(Std Error)	1.456	0.177			2.975	0.338		
Signif	0.000	0.000			0.103	0.011		
	α_2	β_2	σ_1	σ_2				
	-0.001	-0.835	0.038*	0.015*				
(Std Error)	0.002	0.759	0.002	0.001				
Signif	0.524	0.272	0.000	0.000				
	β_{10}	β_{11}	β_{12}	β_{13}	β_{20}	β_{21}	β_{22}	β_{23}
USDSWI	7.969*	**0.641***			8.404*		**-2.273***	
(Std Error)	0.180	0.021			3.210		1.074	
Signif	0.000	0.000			0.009		0.034	
	α_2	β_2	σ_1	σ_2				
	-0.010*	-4.336*	0.034*	0.018*				
(Std Error)	0.003	0.583	0.001	0.001				
Signif	0.001	0.000	0.000	0.000				

Note: * and ** indicate statistical significance at 5 and 10 percent significance level, respectively.

Figure 3: JPYAUS, Smoothed and Filtered Probabilities of Time-Varying
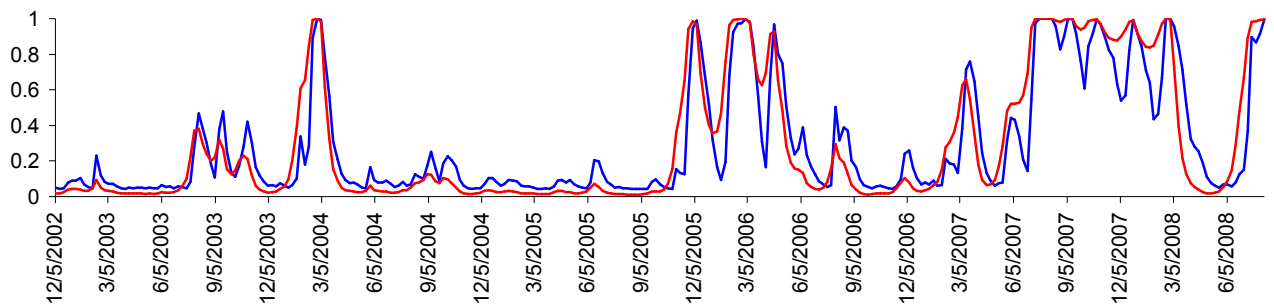Transition Probabilities Markov-Switching Model of the Uncovered Interest Parity
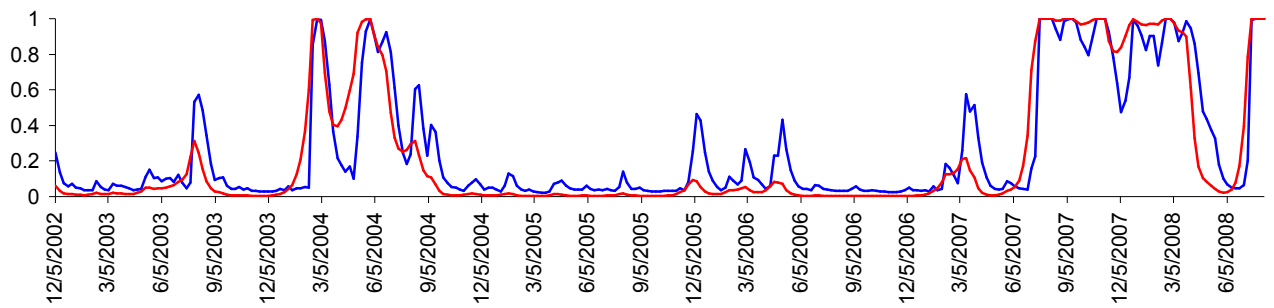
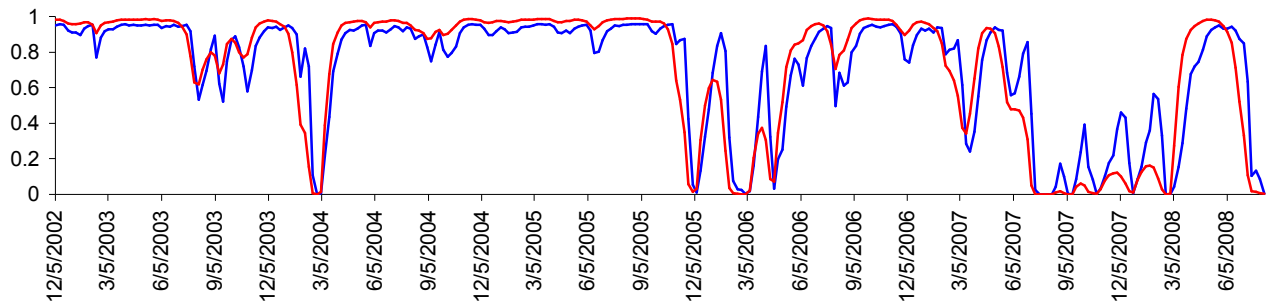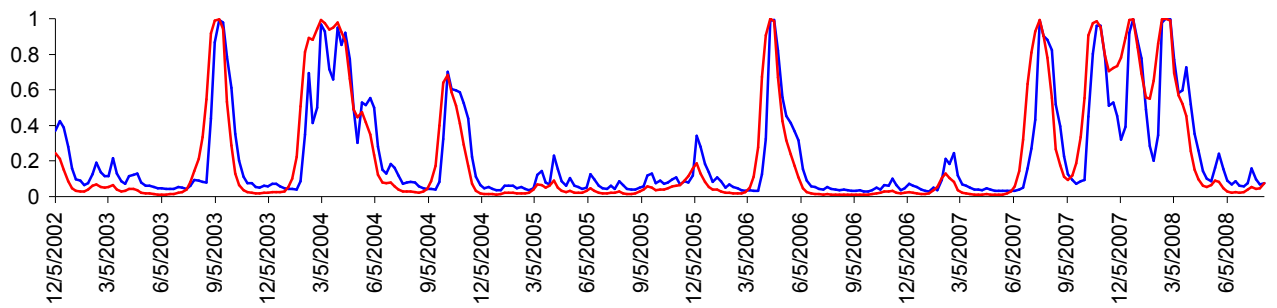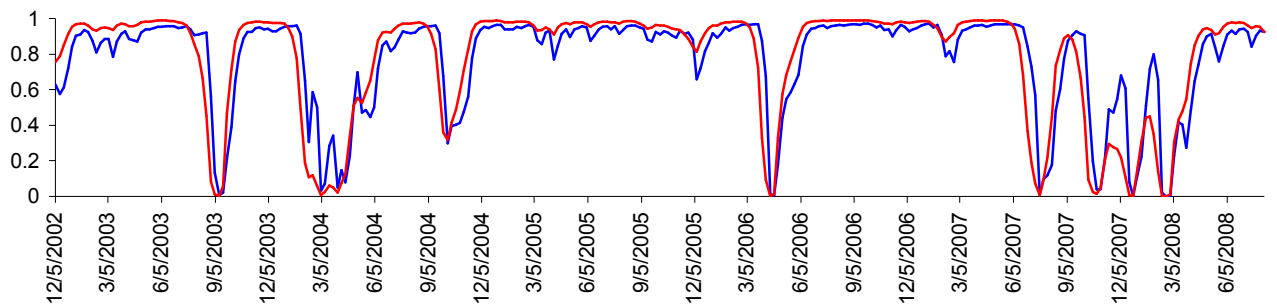(JPYAUS, after 2003, Probabilities of Regime 1, UIP holds)
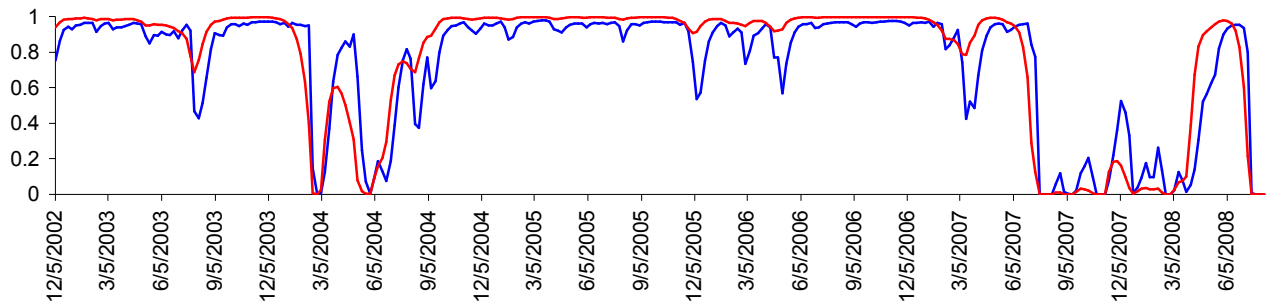
(JPYAUS, after 2003, Probabilities of Regime 2)

Figure 4: JPYNZD, Smoothed and Filtered Probabilities of Time-Varying
Transition Probabilities Markov-Switching Model of the Uncovered Interest Parity

(JPYNZD, after 2003, Probabilities of Regime 1, UIP holds)

(JPYNZD, after 2003, Probabilities of Regime 2)

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Figure 5: USDJPY, Smoothed and Filtered Probabilities of Time-Varying
Transition Probabilities Markov-Switching Model of the Uncovered Interest Parity
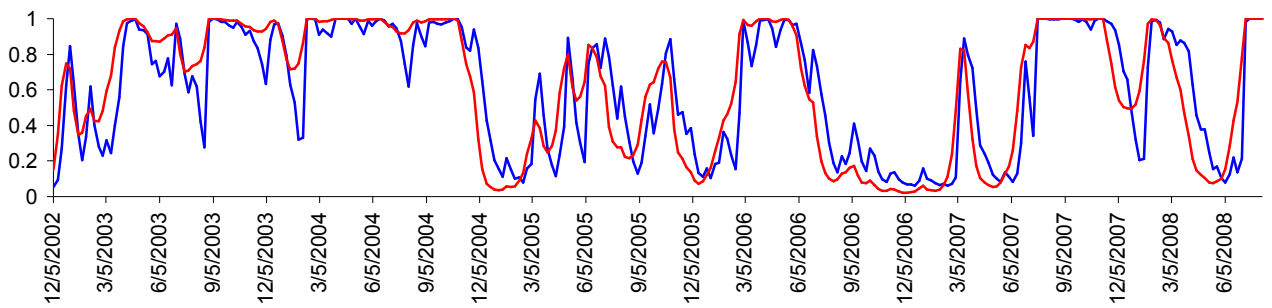

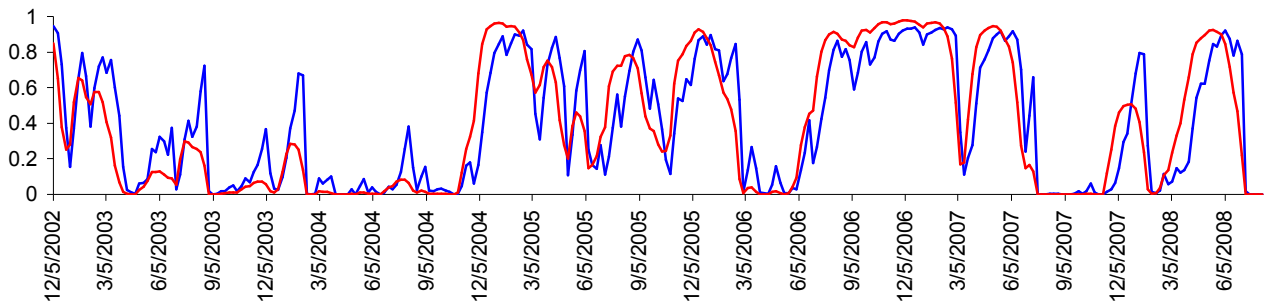(USDJPY, after 2003, Probabilities of Regime 1, UIP holds)

(USDJPY, after 2003, Probabilities of Regime 2)

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Figure 6: USDAUS, Smoothed and Filtered Probabilities of Time-Varying
Transition Probabilities Markov-Switching Model of the Uncovered Interest Parity
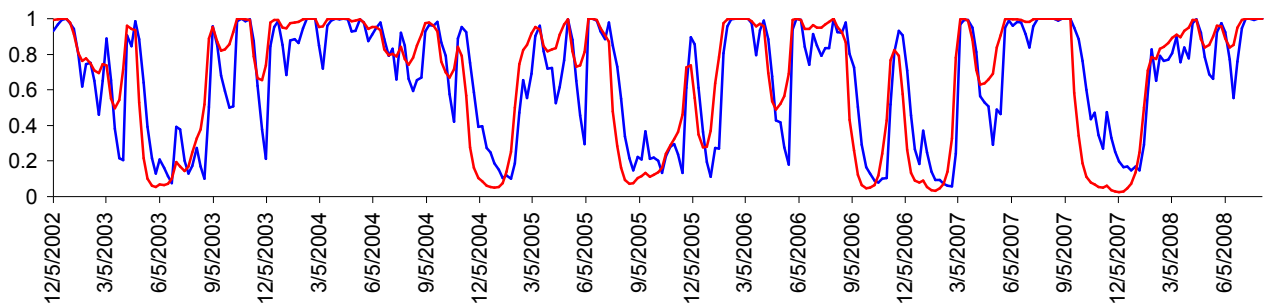

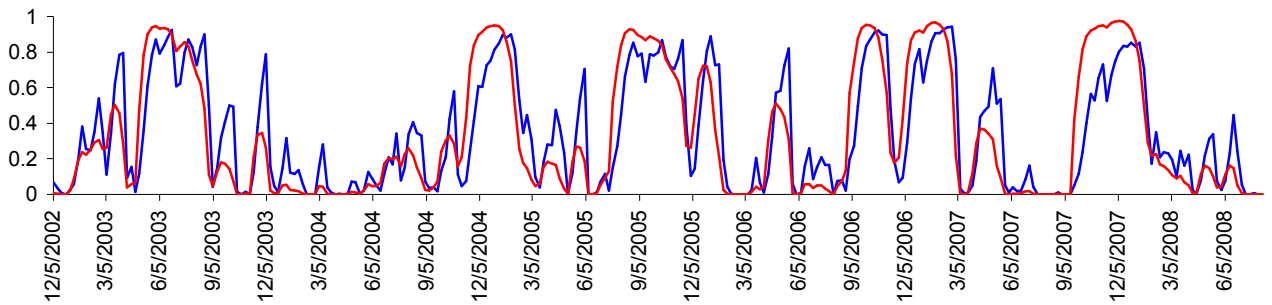(USDAUS, Probabilities of Regime 1, UIP holds)

(USDAUS, Probabilities of Regime 2)

Figure 7: USDNZD, Smoothed and Filtered Probabilities of Time-Varying
Transition Probabilities Markov-Switching Model of the Uncovered Interest Parity
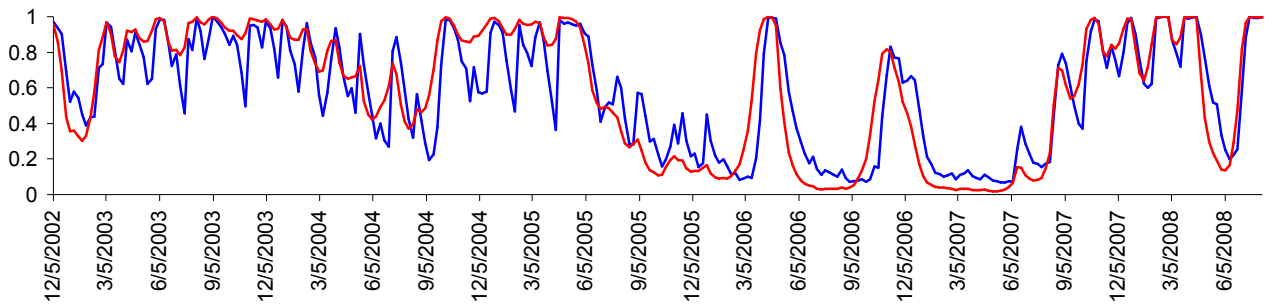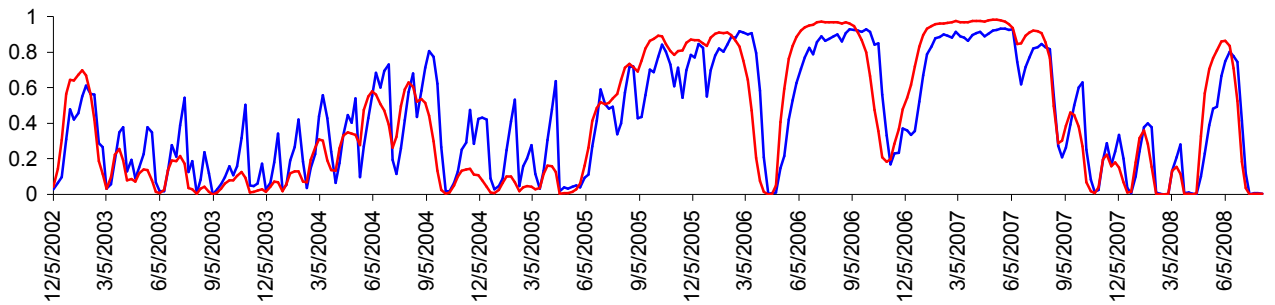

(USDNZD, Probabilities of Regime 1, UIP holds)

(USDNZD, Probabilities of Regime 2)

Figure 8: USDSWI, Smoothed and Filtered Probabilities of Time-Varying
Transition Probabilities Markov-Switching Model of the Uncovered Interest Parity

(USDSWI, Probabilities of Regime 1, UIP holds)

(USDSWI, Probabilities of Regime 2)

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References

Albert, J. and Chib, S. (1993). Bayesian analysis via Gibbs sampling of autoregressive time series subject to Markov mean and variance shifts. *Journal of Business and Economic Statistics*, 11:1–15.

Backus, D. K., Foresi, S., and Telmer, C. I. (2001). Affine term structure models and the forward premium anomaly. *Journal of Finance*, 56:279–304.

Backus, D. K., Gregory, A. W., and Telmer, C. I. (1993). Accounting for forward rates in markets for foreign currency. *Journal of Finance*, 48:1887–1908.

Baillie, R. T. and Bollerslev, T. (2000). The forward premium anomaly is not as bad as you think. *Journal of International Money and Finance*, 19:471–488.

Bekaert, G., Hodrick, R. J., and Marshall, D. (1997). The implications of first-order risk aversion for asset market risk premiums. *Journal of Monetary Economics*, 40:3–39.

Bilson, J. F. O. (1981). The "speculative efficiency" hypothesis. *Journal of Business*, 54:435–451.

Bollerslev, T. (1986). Generalized autoregressive conditional heteroskedasticity. *Journal of Econometrics*, 31:307–327.

Brenner, R. J. and Kroner, K. (1995). Arbitrage, cointegration, and testing the unbiasedness hypothesis in financial markets. *Journal of Financial and Quantitative Analysis*, 30:23–42.

Brunnermeier, M. K., Nagel, S., and Pedersen, L. H. (2008). Carry trades and currency crashes. *National Bureau of Economic Research Working Paper Series*, (14473).

Burnside, A. C., Eichenbaum, M., Kleshchelski, I., and Rebelo, S. (2006). The returns to currency speculation. *Centre for Economic Policy Research Discussion Paper Series*, (5883).

Cumby, R. E. (1988). Is it risk? Explaining deviations from uncovered interest parity. *Journal of Monetary Economics*, 22:279–299.

de Rato, R. (2007). Speech by Rodrigo de Rato. .

Diebold, F. X., Lee, J.-H., and Weinbach, G. C. (1994). ""Regime switching with time-varying transition probabilities". In Hargreaves, C., editor, *Non-stationary Time Series Analysis and Cointegration*. Oxford University Press, Oxford.

Durland, J. M. and McCurdy, T. H. (1994). Duration-dependent transitions in a Markov model of US GNP growth. *Journal of Business and Economic Statistics*, 12:279–288.

Engel, C. (1996). The forward discount anomaly and the risk premium: A survey of recent evidence. *Journal of Empirical Finance*, 3:123–192.

Engle, R. (1982). Autoregressive conditional heteroskedasticity with estimates of the variance of UK inflation. *Econometrica*, 50:987–1008.

Fama, E. F. (1984). Forward and spot exchange rates. *Journal of Monetary Economics*, 14:319–338.

Filardo, A. J. (1994). Business-cycle phases and their transitional dynamics. *Journal of Business and Economic Statistics*, 12:299–308.

Frankel, J. A. and Levich, R. M. (1975). Covered interest arbitrage: Unexploited profits? *Journal of Political Economy*, 83:325–338.

Frenkel, J. A. and Levich, R. M. (1977). Transaction costs and interest arbitrage: Tranquil versus turbulent periods. *Journal of Political Economy*, 85:1209–1226.

Froot, K. A. and Thaler, R. H. (1990). Foreign exchange. *Journal of Economic Perspectives*, 4.

Garcia, R. (1998). Asymptotic null distribution of the likelihood ratio test in Markov switching models. *International Economic Review*, 39(3):763–788.

Goldfeld, S. M. and Quandt, R. E. (1973). A Markov model for switching regressions. *Journal of Econometrics*, 1:3–15.

Gyntelberg, J. and Remolona, E. M. (2007). Risk in carry trades: A look at target currencies in Asia and the Pacific. *BIS Quarterly Review*, pages 73–82.

Hamilton, J. D. (1989). A new approach to the economic analysis of nonstationary time series and the business cycle. *Econometrica*, 57:357–384.

Hamilton, J. D. (1990). Analysis of time series subject to changes in regime. *Journal of Econometrics*, 45:30–70.

Hamilton, J. D. (1994). *Time Series Analysis*. Princeton University Press, Princeton.

Hattori, M. and Shin, H. S. (2007). The broad yen carry trade. *Institute for Monetary and Economics Studies, Bank of Japan Discussion Paper Series*, (07-E-19).

Hodrick, R. J. (1987). *The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets.* Harwood, London.

Hodrick, R. J. (1989). Risk, uncertainty, and exchange rates. *Journal of Monetary Economics*, 23:433–459.

Huisman, R. and Mahieu, R. (2007). Revisiting uncovered interest rate parity: Switching between UIP and the random walk. *Erasmus Research Institute of Management Report Series Research in Management*, (ERS-2007-001-F&A).

Kim, C.-J. (1994). Dynamic linear models with Markov-switching. *Journal of Econometrics*, 60:1–22.

Kim, C. J. and Nelson, C. R. (1999). *State Space Models with Regime Switching, Classical and Gibbs Sampling Approaches with Applications.* MIT Press, Cambridge, MA.

Koyama, K. and Ichiue, H. (2008). Regime switches in exchange rate volatility and uncovered interest parity. *Bank of Japan Working Paper Series*, (07-E-22).

Krolzig, H. M. (1997). *Markov-Switching Vector Autoregressions: Modeling, Statistical Inference, and Application to Business Cycle Analysis.* Springer-Verlag, Berlin.

Lewis, K. K. (1995). ""Puzzles in international financial markets". In Grossman, G. M. and Rogoff, K., editors, *The Handbook of International Economics*, volume 3, pages 1913–1971. North-Holland, Amsterdam.

Melvin, M. and Taylor, M. P. (2009). The crisis in the foreign exchange market. *CESifo Working Paper Series*, (2707).

Newey, W. K. and West, K. D. (1987). A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix. *Econometrica*, 55:703–708.

Sarno, L. (2005). Towards a solution to the puzzles in exchange rate economics: Where do we stand? *Canadian Journal of Economics*, 38:673–708.

Sarno, L. and Taylor, M. P. (2003). *The Economics of Exchange Rates.* Cambridge University Press, Cambridge and New York.

Sarno, L., Valente, G., and Leon, H. (2006). Nonlinearity in deviations from uncovered interest parity: An explanation of the forward bias puzzle. *Review of Finance*, 10:443–482.

Sharpe, W. F. (1994). The Sharpe Ratio. *Journal of Portfolio Management*, 21:49–58.

Suominen, M., Jylha, P., and Lyytinen, J.-P. (2008). Arbitrage capital and currency carry trade returns. *Helsinki School of Economics Working Paper Series*, (W-459).

Taylor, M. P. (1987). Covered interest parity: A high-frequency, high quality data study. *Economica*, 54:429–438.

Taylor, M. P. (1995). The economics of exchange rates. *Journal of Economic Literature*, 33:13–47.

Wagner, C. (2008). Risk-premia, carry-trade dynamics, and speculative efficiency of currency markets. *Oesterreichische Nationalbank Working Paper Series*, (143).

Conclusion

The thesis consists of three self-contained empirical studies. These are (i) J-Curve, Oil Price, House Price and US-Canada Imbalance, (ii) Monetary Policy, Exchange Rates and Asian Stock Markets and (iii) Fear, Volatility and Uncovered Interest Parity. Here, conclusions and implications of each chapter and conclusions of the thesis are addressed.

In chapter 1, we provide answers to questions relating to the imbalance in US-Canada bilateral trade, real exchange rate, real oil price and real asset price. We find that real exchange rate, real oil price and real new housing price index have statistically significant effects on US real trade balance with Canada in the long run. We acquire evidence of short-run J-curve effect with a percentage change in real trade balance equal to -0.45 and -0.54, following a 1 percent depreciation within linear and non-linear framework, respectively. Results from both linear and non-linear models show that short-run dynamic effects of real oil price are not so fearful, with statistically insignificant effects on real trade balance following an increase in real oil price. House price could be argued as being strongly relevant for settlement and adjustment of US trade balance in the

long run through wealth effects, with a distinguished coefficient of US real new housing price index. However, the immediate (next-quarter) effect of a change in housing wealth is insignificant, consistent with existing literature. With the transition probability matrix showing that moving to a regime presenting persistent correction is more likely than the opposite, we believe, reasonably, that a (small) chance to correct US-Canada imbalance prevails.

From the multi-step ahead forecasting exercise, we conclude that US real trade balance with Canada forecasts from our non-linear VAR model outperform ones from the linear VAR in first difference (DVAR) model and ones from the random walk model. Furthermore, our results indicate that the long-term out-of-sample forecastability is not much improved by the oil price and house price variables, which nonetheless actively explain in-sample movement of US real trade balance with Canada in the long run.

From our findings, we support the argument that J-curve effects from the most recent dollar decline, higher oil prices and asset prices could produce a temporary increase in the nominal and real trade balance deficits. We also believe reasonably that, once J-curve effects disappear, a country's trade balance would improve. Nevertheless, besides exchange rate being manipulated so as to gain balance, policies relating to our empirical analysis should be addressed. Since it is less likely that oil-consuming nations could have an influence on world oil prices, the US could increase refining capacity and provide an effective means to restrain the demand for oil in the medium term. Improving conservation and

a shift towards higher energy efficiency are suggested. This could also be good for the environment. With regard to asset price, private savings associated with wealth effects should increase in general. This is in addition to fiscal consolidation and could be assisted mainly by US tax system reforms. Our empirical analysis argues that current asset (housing) price bubble reflation would not help improve US trade deficit (with Canada in particular) over the long run. With appropriate actions, US trade balance deficit (if it still exists) might, then, not be so bad, stimulating its demand and global, in particular Asian, exports.

In chapter 2, we employ a Bayesian structural vector autoregression model to examine the effect of monetary policy and of the exchange rate on stock price movements in Asia. Sign restrictions are used to identify simultaneously and uniquely contractionary monetary policy and exchange rate depreciation shocks in an integrated framework. Our findings are obtained by adapting standard VAR analyses to deal with these single shocks based upon sign restrictions. In particular, for each country examined, impulse responses, variance decomposition and historical decomposition are deliberately considered.

Two main findings emerge. Firstly, monetary policy shocks result in a strongly persistent effect on market index real stock prices whereas the impact of exchange rate shocks is short-lived over the short run. Secondly, with respect to the variance decomposition, the exchange rate is as important as monetary policy for explaining the dynamics of market and financial sector index real stock prices. More precisely, for all the countries examined, real exchange

rate developments have been more important in the short run. In addition, the historical decomposition strongly supports our finding reached by analyzing the variance decomposition. While the exchange rate, generally, is as important as monetary policy for explaining the dynamics of market index real stock prices, real exchange rate shocks appear to have a larger contribution in explaining the deviations from the baseline in some periods of time, for all the countries examined. Based purely on our findings, two conclusions are reached. Firstly, because of the mistimed and/or persistent effect of monetary policy on both real economy and financial markets, we argue that one needs to be cautious in using monetary policy to constrain asset price misalignment. Secondly, due to the evidence that exchange rates principally have a contemporaneous impact on equity prices, we suggest that, in the short run, such incorrectly aligned asset prices might potentially be corrected by focusing on exchange rate movements.

For the first conclusion, while we establish that, in general, monetary authorities could significantly affect equity market valuations by adjusting interest rates, the persistent effects of monetary policy on stock prices found might shed light on the effectiveness of such policy in developing countries in stabilizing the economy. This should be of concern to the central bank. That is, mistimed and/or too persistent effects of monetary policy would only make situations worse if stock price misalignment does not result in significant damage when it ends or if interest rates are high at the moment that a bubble bursts. Moreover, the impact of such high interest rates on the real economy would last for another

couple of years and make the landing harder. This is what we have seen in Asian countries in the aftermath of the 1997 Asian financial crisis. Furthermore, the undesirable effect of such mistimed and/or too persistent monetary policy on financial markets is reflected by the fact that aggregate demand depends positively on the past level of asset prices via the investment balance sheet channel.

From that point of view, we shift to the question of whether other instruments besides interest rates might be used to deal with asset price misalignment. Definitely, reasonable bank regulation and supervision should be thoroughly considered. Strong consensus that a well-structured prudential policy and regulatory system could make financial markets and financial systems less prone to troublesome situations is achieved. This is by helping to reduce the costs of stock price booms and bursts. With these alternatives, the need for contractionary monetary policy conducted by the central bank to burst a bubble is likely to be reduced. They also contribute to the stability of both output and inflation. Nevertheless, deciding what form such regulation and supervision should take is the more difficult issue. It has recently been argued that, in principle, banking regulation should change cyclically to rule out lending booms on the back of rises in asset prices. This should also be considered carefully.

For the second conclusion, confirming that exchange rates are an important determinant of stock prices, our results obtained using recent and growing methodology on the basis of sign restrictions are in line with those of existing literature employing the common method. That is, exchange rates principally

have only a contemporaneous impact on equity prices, resulting in effects in the short run. The implication of this finding might be that, in the short run, asset price misalignment might potentially be corrected by smoothing excessive exchange rate fluctuations.

We have seen that, in emerging markets with large foreign-denominated debt in particular, a financial crisis could be triggered by sharp depreciation. Exchange rate fluctuations are of major concern to monetary authorities, even if they are targeting inflation. As mentioned in the chapter, (inflation-targeting) central banks should not pursue a policy of benign neglect of exchange rates. In other words, monetary authorities in these countries may have to smooth excessive exchange rate fluctuations parted from fundamentals via, for example, foreign exchange market interventions. Nevertheless, this should be conducted without resisting market-determined movements in exchange rates over longer horizons. Such interventions lighten potentially destabilizing impacts of unexpected changes in exchange rates, which, at least in our context, have a contemporaneous impact on asset prices. Nevertheless, it is challenging for central banks, especially ones targeting inflation, to focus on exchange rate movements since this might obstruct them reaching the target rate of inflation.

In chapter 3, we are motivated by the fact that there should be a close link between fear, volatility and the performance of the currency carry trade associated with the failure of the UIP condition. Higher fluctuations in the financial markets, including prices of not only exchange rates but also stocks

and bonds, could worsen excess returns that could be generated through the currency carry trade activities. This is due to the fact that many investors analyze performance, not only by returns, but also by the variability of those returns. That is, they measure their investment performance by the Sharpe ratio in particular. An increase in volatility with no change in return would decrease the Sharpe ratio and a lower Sharpe ratio is less favorable. Consequently, it might be difficult for the foreign exchange carry trade activities to sustain the type of performance it has had if the environment with low volatility does not persist.

Within the context of a time-varying transition probabilities Markov-switching model of the UIP condition, we examine if variables measuring fear and volatility have an effect on the probability of switching between the regime where the UIP condition holds and the regime where it does not. The state transition probability depends nonlinearly upon the variables examined. These are the exchange rate volatility, the VIX equity option implied volatility index and the TED spread.

Applying this to both US dollar exchange rates and cross (exchange) rates from January 4, 1990 to September 11, 2008, we find that those three variables increase the probability of remaining in the regime where the UIP condition holds. In addition, the probability of switching from the regime where the UIP condition does not hold to the regime where the exchange rate follows the UIP condition decreases (increases) as these variables measuring fear and

volatility fall (rise), especially the VIX equity option implied volatility index. For JPYAUS, JPYNZD and USDJPY, the smoothed probabilities show that these exchange rates essentially do not follow the UIP condition except during periods in which the fear and risk variables are increasing, as in the recent global financial crisis in particular.

To conclude, the thesis has considered three important issues of interest to economists and market participants. These are (i) the global imbalance, (ii) monetary policy and asset prices and (iii) the recent global financial crisis. These three issues certainly have implications for the recovery of the global economy. Recently, we have seen a greater international cooperation and attempt to correct the problem that has had an adverse global impact on financial matters and beyond. Nevertheless, it might not be wrong to argue that a solution to the global economic problem ultimately lies with domestic authorities in calibrating the appropriate policy response to the current global economic situation, in addition to international cooperation. Provided with useful economic research results, the domestic authority has to decide whether to take the current difficult situation in the global economy as a lesson or to reproduce the problem.