ESSAYS IN INTERNATIONAL FINANCE

Dissertation Submitted for the Degree of
Doctorate of Philosophy
at the University of Leicester

by

TAMAT SARMIDI

Department of Economic
University of Leicester

March 2008

ProQuest LLC
789 East Eisenhower Parkway
P.O. Box 1346
Ann Arbor, MI 48106-1346

ACKNOWLEDGEMENT

Along this very long and challenging learning process, I inevitably have indebted to countless number of people. First and foremost, I am extremely grateful to my two supervisors Professor Panicos Demetriades and Dr. Sebastiano Manzan. Their support, guidance and wisdom improved this dissertation immeasurably. I am very thankful to Ministry of Higher Education of Malaysia and Universiti Kebangsaan Malaysia (UKM) who jointly sponsored my study. I would like to thank my colleagues at the Department of Economics in particular Halim Alwi, Dr Law Siong Hock, Dr S. and M. Fethi, Dr Rebbecca Munoz-Torres, Dr. Agelos Delis, Baserit Nasa, Tom Allen, Hakem Mobalagi, Angela Yang, Kavita Sirichand, Dr Sebastian Torres, Dr Gregory James, Dr Simeon Coleman, Dr Azlina Aziz and Dr. Nurhani Aba Ibrahim, your presence have been a constant source of inspiration. I am greatly thankful to my family for their unconditional support and encouragement. Last but not least, I am extremely indebted to my beloved wife and children who have accompanied me with love and patience through the highs and the lows of graduate study and whose constant support makes everything possible.

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Tamat Sarmidi

Abstract

This dissertation comprises of three empirical studies on the equity and foreign exchange markets of emerging economies. The motivations for these three studies evolve around the issue of financial liberalization in emerging markets. Specifically, the first empirical study examines the impact of financial liberalization on the volatility of equity returns in the emerging markets. Building on different GARCH models, the chapter shows that volatility could decrease, increase or be unchanged post financial liberalization depending on the level of domestic institutional quality and market characteristics. The analysis shows that volatility is prone to increase (decrease) for a country with low (high) quality of institution and market characteristics. The second study investigates the Uncovered Interest Rate Parity Hypothesis (UIP) for emerging countries. Considering economies that adopt relatively open capital account and free floating exchange rate regime, both dynamic time series and panel analysis suggest that the coefficient of interest rate differential on the UIP regression is positive and close to unity at longer horizons. The evidence is robust for different base countries (US, Germany or Japan). The third empirical study examines the hypothesis that claims that the exchange rate movements are may be predicted in the economies that are fundamentally unstable such as emerging economies. Employing the Vector Error Correction Model (VEC) under the bootstrap techniques proposed by Killian (1999), the findings provide evidence of exchange market predictability in emerging economies.

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CHAPTER ONE

INTRODUCTION

1.1 Background of the Study[1]

Financial liberalization programmes in emerging economies that started in the 1980s were designed to attract capital inflows which could be channelled into economic activity to promote growth and development.[2] These programmes were motivated by the shortage of capital to finance investment (and spur growth) in these economies. As a direct result of financial liberalization policies, emerging countries experienced massive capital inflows over the last few decades. As shown in Table 1.1, in the period 2000 to 2005, total capital inflows to emerging economies, including foreign direct investment (FDI), portfolio investment and other private investment, amounted to $3,815 billion US dollar, more than ten times the figure of $242 billion US dollar achieved in the period 1970 to 1974.

By the end of last decade, however, financial liberalization had become the single most controversial policy prescription. Financial liberalization programmes have been under attack from different sides. The criticisms are mostly focused on the free mobility of short term capital. One view is that financial liberalization has led to excessive appreciation/depreciation of the emerging market's currency, current account deterioration, and more generally, increased instability of the economies. The liquidity and volume effects of large uncontrolled foreign capital entering and

[1] Throughout this dissertation, I use very broad definition of an emerging market that has been used by the International Finance Corporation (IFC) of the World Bank. The IFC defines an emerging market as a market economy with low-to-middle per capita income or one where the ratio of investable market capitalization to GNP is low.

[2] A comprehensive discussion of the liberalization process can be found in Pill and Pradhan (1995 and 1997) for Africa and Asia, Aizenman (2005) for Latin America, and Rodlauer (1995) and Wyplosz (2001) for Eastern Europe and Williamson and Mahar (1998), and Beim and Calomiris (2001) for emerging markets in general.

leaving on the domestic equity, bond and foreign exchange markets have been highly destabilising. These have led to recurrent economic crises in emerging economies (Mexico 1994, East Asia 1997, Russia 1998, Brazil 1999, Turkey 2000 and Argentina 2001, to mention just a few) over the last two decades, which was partly caused by financial liberalization policies (Eichengreen and Bordo 2002; Stiglitz 2002; and Eichengreen 2003). As pointed by Granger *et al.* (2000), short term capital outflows caused the market crises, to a large extent.

Table 1.1: Capital Inflows to Emerging Markets (billions of US dollars) 1970 to 2005

Year	70-74	75-79	80-84	85-89	90-94	95-99	00-05
Foreign direct investment, net	14.36	33.46	54.18	72.35	248.25	758.72	1,116.46
GNP per capita (average)	687.55	795.00	867.70	934.84	979.52	1098.79	1546.32
Net flows on debt, total long-term	60.72	194.93	300.33	184.66	270.08	400.05	181.83
Net flows on debt, total	70.63	257.13	362.49	212.49	390.96	454.88	320.25
Official net resource flows	39.83	97.27	174.06	191.36	266.53	209.90	180.68
Official net transfers	34.08	82.39	139.45	122.59	160.95	70.99	29.93
Portfolio equity flows	-0.007	-0.004	0.12	4.56	78.96	93.07	150.41
Private net resource flows	51.27	168.23	237.94	153.17	481.89	1,180.83	1,503.68
Private net transfers	32.44	92.49	-0.50	-102.3	226.95	657.34	513.98
Short-term debt net flows	8.61	56.61	32.16	41.65	113.27	20.71	168.97

Sources: World Development Indicator, World Bank and the Triennial Central Bank Survey, 2004 of the Bank of International Settlement (BIS).

In some ways these economic crises have created an ambiguity on the advantages or benefits of financial liberalization. Some economists, like Singh (1997), Frazer and Power (1997) and Stiglitz (2002), blame the act of pressuring emerging economies to relax controls on capital mobility during the 1990s as being highly irresponsible. The moves could only induce more instability in the financial markets and bring no significant contribution to economic development. Krugman (1993), for example, argues that there is no significant evidence that capital inflows have any strong positive effect on the growth of developing countries. Further in an analysis in the connections between capital account liberalization and growth, Rodrik (1998) concludes that there is no empirical evidence that countries without capital controls have grown faster, invested more, or experienced lower inflation. Even the International Monetary Fund (IMF) has criticized the uncontrolled capital mobility and has provided (at least some) support for capital controls, particularly in the aftermath of the 1997 Asian currency crisis.[3]

In contrast, the proponents of liberalization strongly believe that financial liberalization could bring more benefit rather than harm to the economy (see Shaw, 1973; McKinnon, 1973; Fry, 1997; De Santis and Imrohoroglu, 1997; Domowitz *et al.* 1998; and Wyplosz (2002) among others). They argue that apart from increasing accessibility of foreign capital to domestic residents, (which could induce more efficient resource allocation) financial liberalization is expected to reduce price instability through greater participation of foreign investors. New investors could broaden the market, which in turn reduces the shock on prices. Furthermore, foreign participators may also make prices more efficient by increasing the precision of public information regarding fundamental values. The advantages of financial

[3] Refer to the article "IMF chief happy to gamble on debt-laden Argentina" written by Alan Beattie in Financial Times, September 15th, 2003; page 16.

liberalization have also been confirmed by a number of econometric analyses such as Bekaert *et al.* (2005 and 2006) who show that financial liberalization could significantly increase annual real economic growth. Galindo *et al.* (2002) find empirical evidence that the liberalization of domestic financial sectors has significantly increased economic growth of countries with intensive external funding. Wyplosz (2002) added that if financial liberalization is not doing much good to the economy, it is not found to do any harm either, at least in the long run. This view has been confirmed by Kaminsky and Schmukler (2003) who argue that huge capital inflows as a result of financial liberalization could only induce a boom-bust cycle in the short run. In the long run, financial liberalization policy will bring more economic prosperity and stability. Obstfeld (1998) claims that the most obvious gain of financial liberalization are the availability of foreign capital to domestic residents.

After experiencing a decade of turbulence in financial markets, another view suggests that financial liberalization could bring more negative effects, induce more instability to financial markets and lack of significant contribution to economic development unless it is implemented with more caution, appropriate timing and proper management especially in developing countries with weak institutions (see among others, Arestis and Demetriades, 1999; Stiglitz, 2004; and Aizenman, 2005). Policy makers in emerging economies are advised to be more cautious and should implement a number of significant reforms to foster the development and stability of their markets. The reforms are mainly focused on minimizing the instability impact of capital flows and reducing the vulnerability to financial crisis. The discussion on the issue of reforming emerging financial markets can be found in the debate surrounding the proposals for a new international financial architecture. Excellent references on

the diverse proposals for reforming the international financial system are Rogoff (1999b), Eichengreen (1999), Stiglitz (1999), Fischer (2002) and Ghosh (2005).

Even though there is disagreement about the policies needed to minimize the incidence of crises, one of the common features that emerge from these proposals is to give priority to strengthen the international and national financial institutions involving governments, banks and corporations. It is argued that markets that have better institutional quality seem to suffer less from economic crises (e.g. Singapore, Australia and New Zealand in the case of the 1997 Asian crisis). It is believed that markets with weak financial supervision and poorly developed domestic capital markets will transfer large changes in capital mobility into macroeconomic volatility. The impact of volatility might thus be dependent on the level of institutional quality and market characteristics of the domestic markets. Markets with strong and high-quality institutional background are assumed to benefit from financial liberalization by experiencing less volatility and enjoying higher economic growth. For instance, Demirguc-Kunt and Detragiache, (1999) empirically show that banking and financial crisis are more likely to occur in liberalized financial system with weak institutions. The probability of banking sector would be adversely affected by financial liberalization is lower in the market with stronger institutional system.

Given that there are real benefits of financial liberalization to the economy but it could also bring instability if not properly managed, thus the main question that remains open is how and when to liberalize an emerging economy to reap the benefit and mitigate the risk associated with the process? Finding the answer to these questions might be helpful in facilitating financial liberalization programmes to achieve their noble goals. Therefore, as noted by Stiglitz (2002 and 2004) that the impact of financial liberalization programmes remains at the centre of the discussion

in emerging economies. Due to these important issues, the first chapter of this dissertation investigates the significant impact of the financial liberalization process on the volatility of equity markets. Further, motivated by the suggestion of Arestis and Demetriades, (1999), it is important to investigate thoroughly the role of institutions in determining the volatility during the process of financial liberalization. Therefore, in the first chapter, the study tries to analyse the link between financial liberalization, volatility of equity returns and institutional quality. The findings from this research are expected to reconcile what seems to be a contradictory view among economists on the real impact of financial liberalization on volatility.

Financial liberalization programmes in emerging markets are not only aimed at equity market but rather to facilitate the whole financial market efficiency and integration, including foreign exchange markets. Surprisingly, besides the importance of foreign exchange markets that are regarded as one of the most actively trading asset markets in emerging economies, only a few studies have systematically investigated the status of currency market efficiency and integration in the post financial liberalization era. Table 1.2 shows the composition of equity and foreign exchange markets in emerging economies. Turnover in the foreign exchange market could reach $0.35 trillion US dollar in a day compared to only $4.6 trillion US dollar for the annual equity market capitalization.

Table 1.2: Summary Measures of Selected Economic Indicators for World, Developed and Emerging Countries (US dollar)

	World		Developed Markets		Emerging Markets	
	1998	2005	1998	2005	1998	2005
Population	5.8 billion	6.4 billion	0.87 billion (15%)	1.0 billion (15.7%)	4.93 billion (85%)	5.4 billion (84.3%)
GNP per capita (constant 2000)	$5,016	$5,655	$25,016	$28,303	$1,126	$1,435
Equity market capitalization of listed companies	$26.9 trillion	$43.6 trillion	$25.6 trillion (95.2%)	$38.9 trillion (89.2%)	$1.3 trillion (4.8%)	$4.6 trillion (10.8%)
Equity market capitalization of listed company (% of GNP)	91.85	99.81	106.92	113.19	24.34	50.29
Foreign Exchange market (trading/day)[c]	$1.5 trillion	$1.8 trillion[d]	$1.28 trillion (85.7%)	$1.45 trillion (81%)[d]	$0.22 trillion (14.3%)	$0.35 trillion (19%)[d]

Sources: World Development Indicator, World Bank. [c] Data is from the Triennial Central Bank Survey, 2004. [d] Data for 2004.

A large number of studies that examined the impact of financial liberalization on the emerging markets focused only on the integration and efficiency of equity markets and implicitly assumes that the other financial markets (including foreign exchange market) are integrated with the equity market (Francis *et al.* 2002). Consequently, many studies (for example) have imposed the same price of world equity market risk on portfolio of equities and foreign currency markets (Dumas and Solnik 1995; De Santis and Gerard 1998). Furthermore, the equity market alone is not a good proxy to generalize financial markets integration. As noted by Frankel (1993), only interest parity tests that can be interpreted without ambiguity as a test of a

country's financial markets integration. In other words, the testing of financial market integration based on equity market without considering other parts of financial markets could be misleading and elusive (Francis *et al.* 2002). In addition, most of the literature on the emerging markets routinely assumes no interest parity exists even though its validity is still subject to discussion.[4] Due to the importance of exchange market integration for the integration of emerging financial market into the world capital markets, in the second chapter we investigate the validity of the Uncovered Interest Rate Parity Hypothesis (here after referred to as UIP) as suggested by Frankel (1993).

The importance of this study is further supported by the fact that the literature that investigates the UIP in emerging economies is relatively scarce. The probable reason is that prior to the financial liberalization period, excessive constraints were imposed by developing countries either on capital movements or exchange rate movements. This situation makes testing the UIP impossible and meaningless. However, the opening up of emerging markets in the late 1980s and early 1990s has created an excellent environment and new room for the study since it fulfils the theoretical assumption of the UIP set-up. Furthermore, the recent literature that tests the validity of the UIP in developed countries has shown significant support for the UIP when longer interest maturities are considered (see Fujii and Chinn, 2001; and Chinn and Meredith, 2004). On the other hand, most of the existing literature that tests the UIP in emerging economies uses a short-term forecast horizon (1- to 3-month). For instance, Flood and Rose (2001), Francis *et al.* (2002) and Frankel and Poonawala (2004) used a 1-month horizon, while Bansal and Dahlquist (2000) used both 1- and 3-month horizons. Motivated by the current developments in emerging

[4] For instance refer to Carvalho *et al.* (2004), and Singh and Banejee (2006) among others who assume that the UIP holds in emerging markets

markets and the need to extend the literature, it is particularly timely and relevant to test the UIP in emerging markets using longer exchange rate forecast horizons and interest rate maturities.

Further, in the third empirical chapter, this study explores the validity of a hypothesis proposed by some economists, like McNown and Wallace (1994) and Rogoff (1996 and 1999a), that the exchange rate movements of emerging countries' currencies should be more predictable (compared to developed countries' currencies).[5] This hypothesis is based on the argument that the correlation between exchange rates and monetary fundamentals should be stronger in countries with higher monetary instability. Since emerging countries are economically more volatile and considering that many emerging economies have adopted relatively floating exchange rate regimes and reduced the constraints on cross-border capital mobility, in the third empirical chapter, we test this hypothesis using various monetary models that are suited to the emerging market environment. Furthermore, to make the test more meaningful, we modify the standard method of calculating deviations from monetary fundamentals to match the requirement of the emerging economies as suggested by Chinn (1998). Specifically, the sticky price and relative price Balassa-Samuelson monetary models are used to account for the characteristics of emerging countries. These models are expected to perform better compared to the standard flexible monetary model, especially for countries that are still in the process of liberalization (see Crespo-Cuaresma *et al.* (2005); Candelon *et al.* 2007). An additional benefit of this study is that, given the fact that the issue of exchange rate

[5] One of the main unresolved issues in the area of exchange rate economics is the question why the monetary model of exchange rate determination cannot forecast much of the variation in exchange rates, the so-called exchange rate forecastability puzzle. The exchange rate forecastability puzzle suggests that macroeconomic fundamentals provide only very negligible predictive content.

predictability is one of the least addressed issues in the emerging economy,[6] investigating the forecastability puzzle in emerging currencies do not only provide an alternative ground to explain the movements of exchange rates, but the study also could be an interesting guide for the policy maker and business strategy.

1.2 Objectives of the Study

This dissertation consists of three empirical chapters that examine the impact of financial liberalization on the equity market's return, the validity of UIP and the predictability of the foreign exchange market in emerging economies. Specifically, the aim of this dissertation is to empirically answer the following four research questions:

1. To investigate the impact of financial liberalization on the volatility of returns in emerging equity markets.
2. To explain the differences of returns' volatility across different markets by considering the level of domestic institutions and market characteristics.
3. To re-investigate the macroeconomic empirical puzzle of the UIP hypothesis for emerging economies both at short and medium horizons.
4. To test the predictability of the exchange rate puzzle using various monetary models and different forecasting horizons for selected emerging economies.

1.3 Contributions

This dissertation attempts to contribute to the literature on emerging markets in three ways.

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[6] Among other issues of interest are the optimal exchange rate regime, (Hochreiter and Tavlas, 2004; Alfaro, 2005), exchange markets integration, (Francis *et al.* 2002; Cheung *et al.* 2006; Rogers, 2006; Tai, 2007), exchange markets and financial crisis, (Phengpis, 2006; Kan and Andreosso-O'Callaghan, 2007) and exchange rate determinants, (Civcir, 2004; Candelon *et al.* 2007).

Firstly, regarding the issue of the impact of financial liberalization on volatility, this study complements the existing literature by explaining the role of institutional quality and market characteristics. The existing literature finds that the volatility effect during financial liberalization is mixed and inconclusive. Some markets experienced increasing volatility while in some others it did not change and others had a decreasing volatility. Taking institutional quality and market characteristic into consideration, the study sheds light on the issue of why volatility differs across markets in the liberalization process. The evidence suggests that volatility increases in markets that have relatively low institutional quality. On the other hand, the markets that have relatively better institutional quality seem to experience either lower or unchanged volatility. This study also extends the number of equity markets and the time span included in the analysis compared to the existing literature. It considers 30 emerging equity markets with monthly data from 1984 to 2005. The existing literature, for instance Bekaert and Harvey (1997) and Jayasuriya (2005), uses only 17 equity markets with a dataset of around 5 years before and after the official liberalization periods. Even though Jayasuriya (2005) has considered institutional quality in her analysis, this study differs from the existing literature in a way that it employs Chinn and Ito (2002 and 2005) KAOPEN Liberalization Index. This index is considered to be more comprehensive compared to the official liberalization date or the Edison and Warnock (2003) Liberalization Intensity Index.[7] Furthermore, the analysis of volatility effect using the Bekaert and Harvey (2005) official liberalization date is also included as a robustness check of the study. In addition to uncovering the importance of institutions and market characteristics in explaining the dynamic of volatility in the liberalization process, the study also sheds

[7] Detailed discussion will follow in the Chapter 2.

light on the sequencing process for the implementation of liberalization policies. It suggests that liberalization should be implemented only after necessary reforms to the domestic institutions are implemented in order to promote the noble goals of globalization.

Second, this study attempts to contribute to the existing literature by analysing the current status of UIP using longer maturity interest rate and exchange rate forecast horizons for emerging countries that adopt relatively flexible exchange rate regimes. The findings could be used to clarify the issue of why empirical testing fails to support the UIP. In fact, this study extends the works of Bansal and Dalhquist (2000), Flood and Rose (2001), Francis *et al.* (2002) and Frankel and Poonawala (2004) that used short horizon variables. The empirical evidence shows that at the long, but not the short horizons, the slope coefficients of β (the coefficient of interest rate differentials) for both time series and panel regressions are positive and getting closer to unity as predicted by the theory. The findings are in favour to the earlier results from the developed economies that show the UIP holds only at long-horizons (Fujii and Chinn, 2001; and Chinn and Meredith, 2004). In addition, the study shows that the success or failure in testing the UIP is sensitive to the selection of the prediction time horizon, k, and therefore it could be used to reconcile the theoretical-empirical puzzle of the UIP testing.

Third, numerous studies have been conducted to examine the predictability of exchange rates in developed markets using various exchange rate determination models, yet none of the results are affirmative. Therefore this dissertation (in the third chapter) attempts to contribute to the literature on this issue by conducting an empirical investigation on the emerging countries' exchange markets using various monetary models (flexible price, sticky price and relative price models) under

different forecasting horizons (1 quarter to 16 quarters). This is because, so far, only a handful of studies have been conducted to investigate the predictability of exchange rates in the emerging markets. The third chapter also reveals strong support for the proposition put forward by economists like Rogoff (1999a) that exchange rate movements are easier to forecast in countries with unstable macroeconomic fundamentals. Specifically, this study finds support for long-horizon predictability for countries that are unstable, such as high inflation economies. Furthermore, the use of a Vector Error Correction Model (VECM) under the bootstrap procedure to test the forecastability of the exchange rates in the emerging markets as proposed by Kilian (1999) has increased the power and reliability of long horizon forecasting accuracy. Therefore, this study yields important lessons to provide an improved understanding of why macroeconomic fundamentals have continuously failed to predict exchange rate movements.

1.4 Outline of the Dissertation

This dissertation is intended to shed new light on three unresolved issues concerning the equity and foreign exchange markets by taking advantage of the financial liberalization process in the emerging economies. These three issues are empirically studied in three different chapters, Chapter 2, 3 and 4. The organization of the dissertation is as follows:

Chapter 2 presents the empirical study on the effect of volatility of returns during the process of liberalization in 30 emerging equity markets. The Generalized Autoregressive Conditional Heteroskedasticity (GARCH) family, including ARCH, GARCH, EGARCH and TGARCH have been employed to capture the effects of volatility. A number of macroeconomic fundamental variables including foreign

stock returns, inflation, interest rates, real exchange rates and the political stability index are considered in the conditional mean equation to show the economic factors that determine the returns. Two liberalization variables, the KAOPEN Liberalization Intensity Index by Chinn and Ito (2002 and 2005) and the Official Liberalization Date by Bekaert *et al.* (2005) have been used in the conditional variance equation to explain the effect of liberalization on the volatility. Furthermore, the question of why different volatility effects are observed across markets is critically analysed by linking them to the domestic institutional quality and market characteristics. The variables included in the institutional quality are corruption, law and order, and bureaucratic quality. While the market characteristics variables are market transparency, investment profile and market exit openness.

Chapter 3 deals with the empirical investigation of the validity of the UIP hypothesis for the emerging economies utilising a longer exchange rate forecast horizon and interest rate maturity period. Considering the nonstationarity and persistency properties of the series, both dynamic panel and time series techniques are used to unveil the importance of the exchange rate prediction horizons in determining the status of the UIP.

Chapter 4 attempts to explain the exchange market forecastability puzzle using a dataset from relatively more volatile economies i.e. emerging countries. The analysis is restricted to markets that adopt a floating exchange rate regime and impose less control on capital mobility. The empirical testing is carried out using three monetary models, flexible price, sticky price and relative price for 1-, 8-, 12- and 16-quarter horizons. The fundamental value is constructed using Mark's (1995) method after considering emerging market characteristics as suggested by Chinn

(1998). The Kilian (1999) vector error correction (VEC) bootstrap method is used to test the hypothesis of no predictability.

The last chapter, Chapter 5, is the conclusion. It provides a summary and discussion of the overall findings and policy implications as well as suggesting new avenues for further research.

CHAPTER TWO

WHY THE VOLATILITY OF RETURNS DIFFERS ACROSS MARKETS AFTER LIBERALIZATION?

2.1 Introduction

One of the main questions which has been least addressed by researchers in the financial liberalization literature is the issue of why volatility differs across markets during the liberalization process in emerging markets? To be more specific, why are some markets more volatile after liberalization compared to others?[8] Finding the correct answer is vitally important for the proper functioning of the capital market and the stability of an emerging country's economy. This is because volatility might not only have a financial effect, but a real destabilizing effect on the economies, such as by increasing the cost of capital, the misallocation of resources and also the reluctance of risk averse firms to raise funds or even seek a listing on the stock market. On the contrary, if liberalization does not cause excessive volatility of returns, then opening-up emerging markets to the international investors should produce the sought-after results postulated by its advocates.

The issue has attracted considerable attention from at least two different groups of economists. The first group, such as Grabel (1995) and Singh (1997), argue that financial liberalization could induce more instability in financial markets and bring no significant contribution to economic development. Singh (1993 and 1997) vilifies untimely financial liberalization in emerging markets for creating an increasingly noisy stock market environment, and even, in some circles, for impeding growth. He claims that emerging countries are not yet well equipped with

[8] Refer to Appendix 1 for details of the literature on the issue.

the necessary and essential infrastructures that are vital for well functioning markets such as regulatory and institutional structures. This is because the central problem with stock markets as a source of funds in open economies is the need for instant liquidity (Bhide 1994). Investors are not bound to any long-term commitment to the firms and by right could pump the funds in or out at any instance and for whatever reasons. Furthermore, with the current level of information technology assisting quicker transaction times, domestic markets may be perfectly and instantaneously exposed to uncertainties abroad (Fraser and Power 1997). Many empirical studies have found support for this view. Grabel (1995) who constructs the volatility indices based on the neo-classical and the Keynesian framework find volatility increases for most of the emerging stock markets after liberalization. Aitken (1996) tests weekly stock market return using variance ratio tests for 16 emerging markets finds evidence that volatility increases following liberalization. Experiences over the last decades seem to be in favour of this view since there was no single period where equity markets in emerging countries were free from turbulence (Eichengreen and Bordo 2002).

On the other hand, the advocates of financial liberalization, for instance Obstfeld (1994 and 1998), argue that in the presence of efficient financial markets, the financial deepening associated with financial market liberalization should decrease overall market instability. This is because financial market liberalization could increase the market size by attracting more high yield but risky investment in developing countries through risk diversification. Available capital could encourage more economic activities and increase profit opportunities, which in turn should result in higher growth levels and lower volatility. Bekaert and Harvey (2000) argue that even if volatility increases, this may not be damaging in the long run to the real

economy; as stated by Kaminsky and Schmukler (2003), 'short-run pain, long-run gain'. Empirical evidence to support this view can be found in Kim and Singal (2000) who find no significant evidence suggesting that volatility increases following liberalization for 20 emerging markets around 5 years pre- and post-liberalization using ARCH and GARCH models. Similar finding is also reported by De Santis and Imrohoroglu (1997) who do not find any systematic effect of market liberalization on stock return volatility.

The main objective of this paper is to shed light on the contradictory results found in the literature on the relationship between financial liberalization and stock market volatility in emerging economies. Even though a large empirical literature on developed markets exists, only a few studies have been conducted on emerging equity markets, including Kassimatis (2002) on six emerging markets, Kim and Singal (2000) on fourteen emerging markets, Levine and Zervos (1998) on sixteen emerging markets, Bekaert and Harvey (1997 and 2000) on seventeen emerging markets, Huang and Yang (2000) on ten emerging markets, and Jayasuriya (2005) on eighteen emerging markets. Except for Jayasuriya (2005) who investigates the role of market characteristics and institutional quality in explaining the impact of liberalization on volatility, the other papers are mainly interested in the impacts of volatility under the liberalization process.[9] Furthermore, the findings are also not consistent across countries and therefore it can be concluded that the volatility effect after the liberalization process is country-specific in nature.

Motivated by Arestis and Demetriades (1999) and by the debate surrounding the proposals for new international financial architecture (see for instance, Rogoff 1999b; Eichengreen 1999; Stiglitz 1999; and Fischer 2002), this study further

[9] Levin and Zervos (1998) do analyse the role of accounting standards and investor protection laws to the development of stock markets.

attempts to investigate the reasons why the volatility effect differs across countries by thoroughly looking into the various market characteristics and institutional quality for thirty emerging markets with a longer sample period. This is necessary because the real impact of financial liberalization is expected to vary between different time frames and across heterogeneous emerging markets as reported in most of the literature. (See for example Demirguc-Kunt and Detragiache (1999) for the effect of financial liberalization on the banking sectors, and Kaminsky and Schmukler (2003) for implication of financial liberalization on stock markets.)

Consistent with previous findings, our results indicate that liberalization has different effects on the volatility of returns depending on the country's institutional quality and market characteristics. Generally, economies characterized by low-level market characteristics and poor-quality institutions could experience higher volatility in the post liberalization period, while the opposite holds for countries with higher quality institutions and market characteristics. We are thus able to reconcile the different views on the real impact of liberalization on volatility by considering the relevant role played in this process by institutional and market characteristics.

From a policy perspective, the results provide guidance on the liberalization process for developing countries. It is of central importance to first upgrade the domestic financial system to a satisfactory level that could foster more efficient markets in order to avoid excess and undesired volatility effects on the economy. This view conforms to the existing literature in finance, liberalization and development (McKinnon, 1991; Arestis and Demetriades, 1999; Arestis *et al.*, 2002; Demetriades and Andrianova, 2003; Demetriades and Law 2006).

The rest of the paper is organized as follows. Section 2.2 briefly reviews the literature on financial liberalization, markets, institutional quality and volatility.

Sections 2.3 and 2.4 discuss the data and econometric models used in the empirical analysis, respectively. Section 2.5 presents the results and econometric specification test analysis. Finally, Section 2.6 concludes the paper.

2.2 Literature Review

The importance of good institutional quality and market characteristics in promoting sustainable and balanced growth in the liberalization process is not a controversial issue. Arestis and Demetriades (1999), and Demetriades and Andrianova (2003) have documented an extensive argument that urges the need for reliable markets and institutions to promote development in the process of liberalizing an economy. The empirical evidence suggests that an accumulation of factors, such as capital, human capital and technological change alone is not sufficient to explain differences in growth performance across countries. It is postulated that good quality institutions are expected to further deepen the financial markets and able to cope effectively with certain financial disequilibria (Fanelli 2007). Relevant research in this area is Easterly and Levine (1997), Pistor *et al.* (2000), and Demetriades and Law (2006). Generally, they find a significant contribution from better institutions to economic growth.

In contrast to the large body of work investigating the link between institutions and growth, relatively little work has examined the link between stock market volatility and institutions. Nelson and Sampat (2001) technically define institutions as 'social technologies' and provide an excellent explanation of their role in affecting economic performance and volatility. They postulate that when institutions are of low quality, due to frequent changes of rules, high levels of corruption, widespread nepotism and weak law enforcement, the markets will not be

functioning well and volatility and the allocation of resources will be severely affected. Accordingly, high-quality market characteristics play an equivalently important role in promoting an efficient and low-risk investment environment (La Porta *et al.* 1998; Johnson *et al.* 2000; Allegret *et al.* 2003).

Selected papers that try to explore the paradox of institutions, volatility and liberalization are Stiglitz (1999, 2000), Bekaert and Harvey (2000, 2003, 2005, and 2006), Caner and Onder (2005), and Jayasuriya (2005). Referring to the 1997 Asian financial crisis, Stiglitz (1999) insists on the need for sound institutions and transparent financial systems if economic instability and unsustainable development after liberalization are to be avoided. However, since the financial crisis spread out to countries with relatively well-developed financial institutions, Stiglitz (2000) launches a question on the reliability of full liberalization even within such good economic environments. However, Bekaert and Harvey (2000) argue that liberalization could not trigger volatility in whatever circumstances. Their study uses a pooled cross-section and time series regression on the conditional variance obtained from a variant of the GARCH models to investigate the impact of foreign speculators in the emerging markets on the expected returns and volatility. They could not detect any significant increase in volatility, despite the huge increase in capital inflows and outflows to the emerging countries after liberalization. Bekaert and Harvey (2003) further confirm the finding by comparing the return volatility of emerging markets with the situation in developed markets. They could not find any significant differences between volatility patterns in the two types of market from 1977 to 2002.

Caner and Onder (2005) try to dig deep into the sources of returns volatility in emerging markets in the post-liberalization era. They use a VAR model to

estimate the contribution of fundamental and additional factors that represent fiscal and monetary policy to the variation in stock returns for seventeen emerging markets. They found that the main sources of returns volatility in emerging markets are dividend yield and lagged return that account for 43% and 46%, respectively, while short-term, real interest rate, exchange rates and world markets play only a minor role. However, all of these studies fail to directly answer the question of why volatility increases in some countries and decreases in other countries after liberalization.

Estimating a variant of the GARCH model, Jayasuriya (2005) finds that institutional quality and market characteristics, such as market transparency, investor protection practice, level of corruption, rule of law, bureaucratic quality and market exit openness do contribute to the volatility of returns in emerging markets during the liberalization process. She found that countries with high levels of institutional quality and market characteristics experience decreased volatility. On the other hand, returns in the markets characterized by low institutional quality and poor market characteristics were found to be more volatile after liberalization. However, Jayasuriya (2005) considers only relatively short-term periods around liberalization. It is possible that in such a short period of time the economic agents may not fully respond to the new rules and regulations enacted in the economic system.

2.2.1 Measuring Institutional Quality

This study departs from much of the extant literature by considering market characteristics and institutional quality as a factor in the different impacts of financial market liberalization on volatility. We borrow the concept of 'social technologies' from Nelson and Sampat (2001) to represent institutions. Various types of proxies

have been used to measure the level of institutional quality and market characteristics in emerging markets. Bekaert *et al.* (2005, 2006) use the International Country Risk Guide's (ICRG) indices that accounts for political risk, corruption, law and order, and bureaucratic quality indices to proxy institutional quality in order to examine the impact of liberalization on the consumption growth variability. In addition, they used creditor rights and accounting standards taken from La Porta *et al.* (1998) to strengthen and deepen market characteristics variables. Keefer and Knack (1997), and Demetriades and Law (2004) also employ five institutional quality indices published by Political Risk Services (PRS) to measure overall institutional quality in emerging markets. These indices are corruption, rule of law, bureaucratic quality, government repudiation of contracts and risk of expropriation.

In this study, we follow Bekaert *et al.* (2005) and use three ICRG indices to measure the overall institutional quality. First, we use *Corruption*, which is a component of the ICRG index for the quality of institutions. It measures corruption within the political system that arises from excessive patronage, nepotism, job reservation, 'favour-for-favours', secret party funding, and suspicious ties between politics and business. A high level of corruption distorts the economic and financial environment and reduces the efficiency of the government and businesses by enabling people to assume positions of power through patronage rather than ability. Such corruption would potentially lead to popular discontent, unrealistic and inefficient controls of the state economy and encourage the development of the black market. The index ranges between zero and six and the higher the corruption, the lower the index.

The second index that we consider is *Law and Order*. The *Law* sub-component is an assessment of the strength and impartiality of the legal system,

while the *Order* sub-component is an assessment of popular observance of the law. However, both measures are assessed separately. The index ranges between zero and three with zero as a low rating and three as a high rating. A country can enjoy a high rating (3.0) in term of its judicial system, but a low rating (1.0) for order if the law is ignored for political reasons.

The third index used in our analysis is *Bureaucratic Quality*. It is a sub-component of ICRG's quality of institutions. The Bureaucratic Quality Index measures the strength of institutions to serve as a shock absorber that tends to minimizes revision of policy when governments change. The scale ranges from zero to four. Therefore, high points are given to countries where the bureaucracy has the strength and expertise to govern without drastic changes in policy or interruptions in government services. Countries that lack the cushioning effect of a strong bureaucracy receive low points because a change in government tends to be traumatic in terms of policy formulation and day-to-day administrative functions.

2.2.2 Measuring Market Characteristics

Market characteristics play a decisive role for capital inflow or outflow in emerging markets. This is because institutional investors in developed markets prefer to invest in emerging markets that have strong market characteristics. Broadly speaking, markets are characterized as good and strong if they are transparent, have a low risk of expropriation, high levels of contract viability and unimpeded movement of financial capital. Good market characteristics will serve as a capital magnet for emerging economies.

In this paper, we use three broad groups of market characteristics variables. First is market transparency. Transparency is important since it is one of the

theoretical conditions required for a market to be efficient and might contribute to the reduction of returns volatility (Johnson *et al.* 2000; Tomioka, 2001; Rodan, 2002). Following La Porta (1998) and Jayasuriya (2005), we use an accounting standards variable to proxy market transparency because excellence in accounting standards guarantees that investors have all the necessary and relevant information about companies. La Porta's (1998) accounting standards index is created by examining and rating companies' annual reports on their inclusion or omission of 90 items. These items fall into seven categories (general information, income statement, balance sheets, funds flow statement, accounting standards, stock data, and special items). The higher the rating, the more transparent is the market. Jayasuriya's accounting standards index is constructed from various issues of the *Emerging Stock Markets Factbook.* The index ranks from one (poor, requires reform), to two (adequate), to three (good, of internationally acceptable standards).

The second variable that we consider is the investment profile, which is a sub-component of ICRG's political risk index. It measures the government attitude to inward investment. The investment profile is determined by PRS's assessment of three sub-components: (i) risk of expropriation or contract viability; (ii) payment delays and (iii) repatriation of profits. Each sub-component is scored on a scale from zero (very high risk) to four (very low risk).

The third variable is market exit openness. It measures separately the restrictions imposed by the domestic market government on the repatriation of foreign income and capital. The index is range from one (closed, which means no repatriation of income or capital) to five (free repatriation of income or capital). Accounting standards and market exit openness variables are based on the index developed by Jayasuriya (2005) with some extension to 2001 to capture more recent

developments in the markets. Appendix (A) of Jayasuriya (2005) provides a detailed explanation of the index construction. These two indices are constructed based on information from various issues of the *Emerging Stock Markets Factbook.* Theoretically, a market with a high level of transparency, good investor protection laws and less restriction on exit or entry to the markets will attract more foreign institutional investors to the domestic market.

2.3 Data

The data for this study consists of three major variables: stock returns, liberalization indicators and macroeconomic fundamentals for thirty emerging markets.

2.3.1 Stock Returns

The stock return used in this study is defined as the first difference of the logarithm of monthly average stock indices ($y_t = ln(p_t/p_{t-1})$), where y_t is the stock return and p_t is the stock price. All stock market indices are retrieved from the Standard and Poor's/International Finance Corporation (IFC) Emerging Markets Database or respective domestic stock exchange index provided by Datastream. The countries considered are: 7 Latin American emerging markets, Argentina, Brazil, Chile, Colombia, Mexico, Peru and Venezuela; 10 Asian emerging markets, Bangladesh, India, Indonesia, Korea, Malaysia, Pakistan, Philippines, Singapore, Sri Lanka and Thailand; 7 African and Middle-Eastern emerging markets, Israel, Jordan, Kenya, Nigeria, Morocco, South Africa and Zimbabwe; and 6 European emerging markets, Austria, Belgium, Hungary, Poland, Spain and Turkey. Three developed stock markets; the U.S. (S&P500 index), the U.K. (FTSE100 index) and Japan

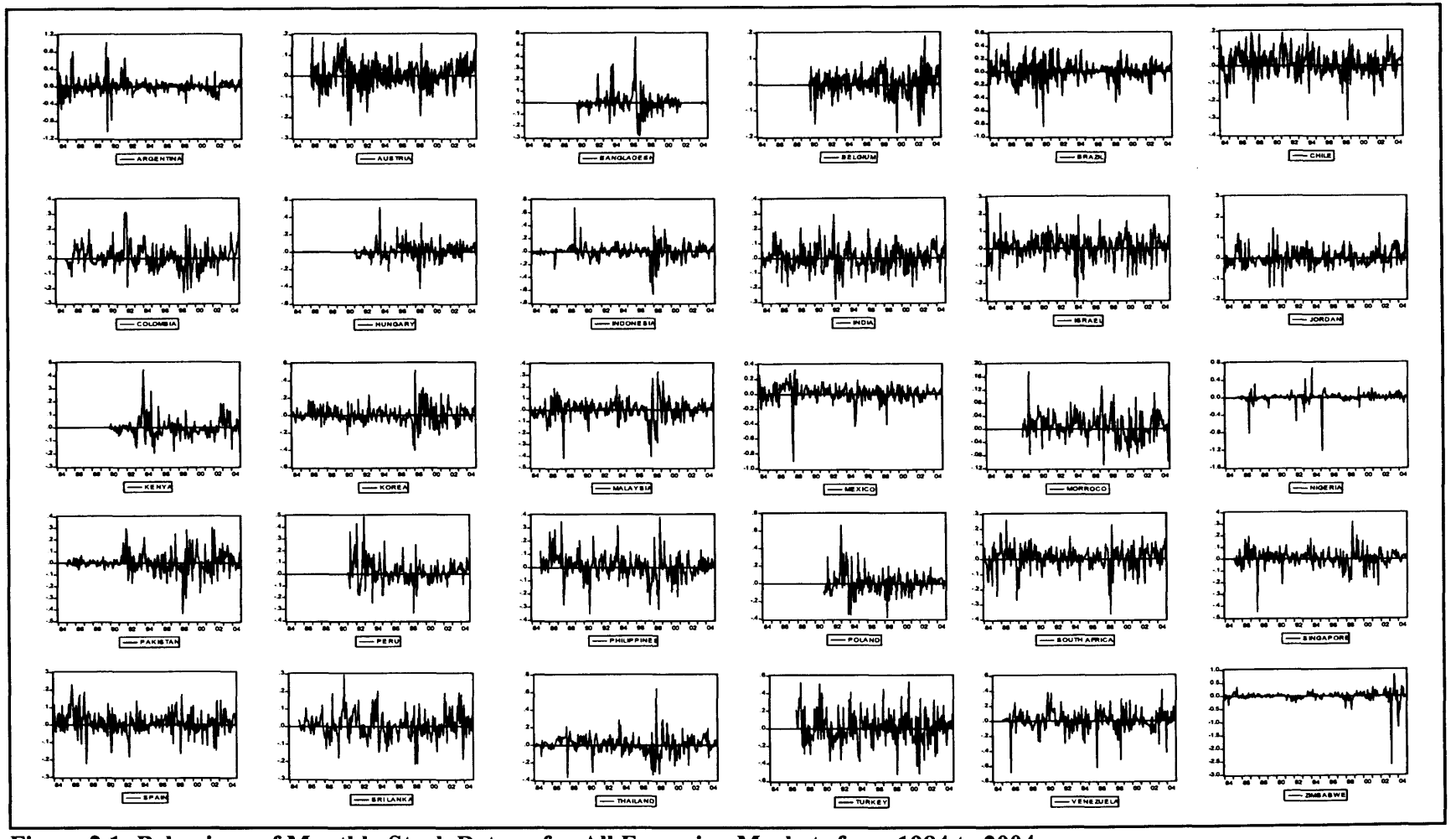


Figure 2.1: Behaviour of Monthly Stock Return for All Emerging Markets from 1984 to 2004.

(Nikkei225 index) are used to represent world market returns. Figure 2.1 graphically presents all the emerging equity markets' returns from January 1984 to December 2004, yielding a total of 252 monthly observations for most of the countries. The details of the sample periods and the name of stock exchange indices included in the estimation are in Appendix 2. From the Figure 2.1, it appears that there are periods where the volatility is relatively high and other periods where the volatility is relatively low. This indicates that there is evidence for time variation in the second moment of the return distribution that could show evidence of volatility clustering and has a strong autocorrelation in squared returns.

Figure 2.1 also clearly shows that all returns series are likely to be stationary in the mean although not necessarily in the variance and there are no obvious trends in the data. An Augmented Dickey Fuller (ADF) and Phillip Perron (PP) unit root test are carried out to statistically justify the absence of unit root in each return series. Table 2.1 confirms that the null hypothesis of a unit root is rejected for both the ADF and the PP test at the 1% critical level for all series, suggesting no obvious sources of non-stationarity in the return series.

Table 2.2 illustrates the descriptive statistics of the emerging market returns. On average, Asian market returns have a lower mean and lower unconditional variance (as measured by standard deviation) compared to the others. On the other hand, for countries in Latin America, returns generally show a considerably higher mean and standard deviation. In addition, the descriptive statistics show that returns are negatively skewed for most African, Middle Eastern and Latin American markets but positively skewed for Asian and European markets. This is suggesting that the market returns from Asia and Europe have a heavier tail of positive values relative to other regions in the sample. Furthermore, most of the returns show consistently

leptokurtic series, which indicates non-normality in the returns. The Jarque-Bera statistic confirms that a normal distribution hypothesis is rejected at a high level of significance for the majority of the return series. These returns characteristics conform to the majority literature concerning emerging markets (Bekaert and Harvey, 2003).

Table 2.1: Augmented Dickey-Fuller (ADF) and Philip-Perron (PP) Tests for All Monthly Market Returns from 1984:01 to 2003:12

Countries Return	ADF		PP	
	Intercept and no trend	Intercept and trend	Intercept and No trend	Intercept and trend
Argentina	-17.08	-17.05	-17.18	-17.16
Austria	-14.38	-14.27	-14.45	-14.41
Bangladesh	-10.64	-10.52	-10.84	-10.80
Belgium	-12.64	-12.51	-12.94	-12.93
Brazil	-15.78	-15.67	-15.97	-15.91
Chile	-12.59	-12.52	-12.84	-12.82
Colombia	-10.09	-10.01	-10.32	-10.30
Hungary	-12.74	-12.70	-12.86	-12.81
India	-13.28	-13.20	-12.42	-12.38
Indonesia	-14.20	-14.14	-14.32	-14.28
Israel	-15.02	-14.97	-15.24	-15.19
Jordan	-13.79	-13.72	-13.85	-13.79
Kenya	-9.60	-9.56	-9.71	-9.68
Korea	-14.68	-14.62	-14.72	-14.69
Malaysia	-12.98	-12.91	-13.11	-13.07
Mexico	-11.55	-11.50	-11.67	-11.61
Morocco	-12.82	-12.78	-12.93	-12.89
Nigeria	-15.29	-15.24	-15.35	-15.30
Pakistan	-14.74	-14.70	-14.87	-14.81
Peru	-11.57	-11.52	-11.68	-11.62
Philippines	-11.39	-11.33	-11.67	-11.62
Poland	-12.16	-12.10	-12.28	-12.23
S. Africa	-14.24	-14.18	-14.33	-14.28
Singapore	-13.30	-13.24	-13.47	-13.39
Spain	-13.68	-13.61	-13.79	-13.72
S. Lanka	-12.17	-12.05	-12.22	-12.20
Thailand	-13.97	-13.91	-14.11	-14.03
Turkey	-13.77	-13.72	-13.91	-13.85
Venezuela	-15.45	-15.39	-15.51	-15.46
Zimbabwe	-14.10	-14.06	-14.21	-14.17

Note: All statistics are significant at the 1% level. The lag length has been selected based on AIC to ensure white noise residual. The critical values provided by MacKinnon (1996).

Table 2.2: Descriptive Statistic of the Emerging Markets Returns

Country	Obs	Mean	Median	Max	Min	S.D	Skew[a]	Kurt[b]	Jarque-Bera[c]
				Latin America					
ARG	252	0.0050	0.0112	1.02	-1.05	0.1947	0.2526	11.4174	746.6*
BRA	252	0.0050	0.0109	0.45	-0.84	0.1665	-0.5946	5.8878	102.4*
CHI	252	0.0106	0.0084	0.19	-0.31	0.0762	-0.2873	3.8921	11.8*
COL	239	0.0091	0.0016	0.31	-0.23	0.0831	0.5741	4.8774	48.2*
MEX	252	0.0112	0.0198	0.32	-0.90	0.1248	-2.2178	15.9083	1956.1*
PER	168	0.0174	0.0083	0.48	-0.33	0.1088	0.8719	6.4215	103.2*
VEN	239	0.0020	0.0042	0.41	-0.69	0.1411	-0.9404	7.4457	232.0*
Average		0.0086	0.0092	0.45	-0.62	0.1279	-0.3345	7.9786	
				Asia					
BAN	180	-0.0021	-0.0008	0.57	-0.28	0.0924	1.5159	12.9372	809.5*
IND	252	0.0042	0.0040	0.30	-0.28	0.0871	0.1092	3.3805	2.0
INA	252	-0.0016	-0.0006	0.67	-0.67	0.1262	-0.4589	11.5787	781.5*
KOR	252	0.0044	-0.0061	0.51	-0.41	0.1087	0.2398	5.8751	89.2*
MAL	252	-0.0013	0.0013	0.33	-0.41	0.0927	-0.6276	6.6146	153.7*
PAK	240	0.0022	-0.0015	0.30	-0.43	0.0941	-0.0667	6.4647	120.2*
PHI	240	0.0072	0.0043	0.38	-0.35	0.1019	0.1086	5.2377	50.5*
SGP	239	0.0060	0.0093	0.32	-0.45	0.0784	-0.7728	8.6264	339.0*
SRI	239	0.0060	0.0046	0.29	-0.21	0.0750	0.2099	7.1324	214.5*
THA	252	0.0043	0.0071	0.64	-0.36	0.1093	0.2876	8.3652	305.7*
Average		0.0029	0.0022	0.43	-0.38	0.0966	0.0545	7.3212	
				Africa and Middle East					
JOR	252	0.0027	-0.0037	0.21	-0.14	0.0456	0.5876	5.1815	64.4*
ISR	252	0.0085	0.0132	0.26	-0.28	0.0720	-0.3933	4.6822	36.2*
KEN	180	-0.0025	-0.0097	0.44	-0.20	0.0792	1.5517	9.5118	390.2*
MOR	203	0.0086	0.0061	0.17	-0.11	0.0415	0.2523	4.2748	15.8*
NIG	240	0.0006	0.0102	0.68	-1.23	0.1379	-3.5052	36.8216	1930.4*
SAF	252	0.0070	0.0103	0.26	-0.36	0.0794	-0.8021	5.7371	105.6*
ZIM	252	-0.0032	0.0148	0.82	-2.61	0.2320	-6.1817	68.3377	6429.6*
Average		0.0031	0.0059	0.41	-0.70	0.0982	-1.2129	19.2210	
				Europe					
AUS	227	0.0057	0.0068	0.18	-0.23	0.0690	-0.1940	3.7334	6.5*
BEL	180	0.0030	0.0051	0.18	-0.18	0.0522	-0.4960	4.6880	28.7*
HUN	168	0.0074	0.0055	0.51	-0.42	0.1017	0.5131	8.2537	200.5*
POL	168	0.0133	0.0084	0.67	-0.38	0.1330	0.7346	6.9680	125.3*
SPA	252	0.0125	0.0086	0.23	-0.22	0.0666	-0.0931	4.0288	11.4*
TUR	216	0.0092	0.0011	0.53	-0.52	0.1838	0.1256	3.4611	2.4
Average		0.0085	0.0059	0.38	-0.32	0.1010	0.0983	5.1888	

[a] Skew is skewness. It measures the asymmetry of the distribution of the series around its mean. The skewness of a normal distribution is zero.

[b] Kurt is Kurtosis. It measures the peakedness or flatness of the distribution of the series. The kurtosis of the normal distribution is 3. If the kurtosis exceeds 3, the distribution is leptokurtic and if less than 3 platykurtic relative to the normal distribution.

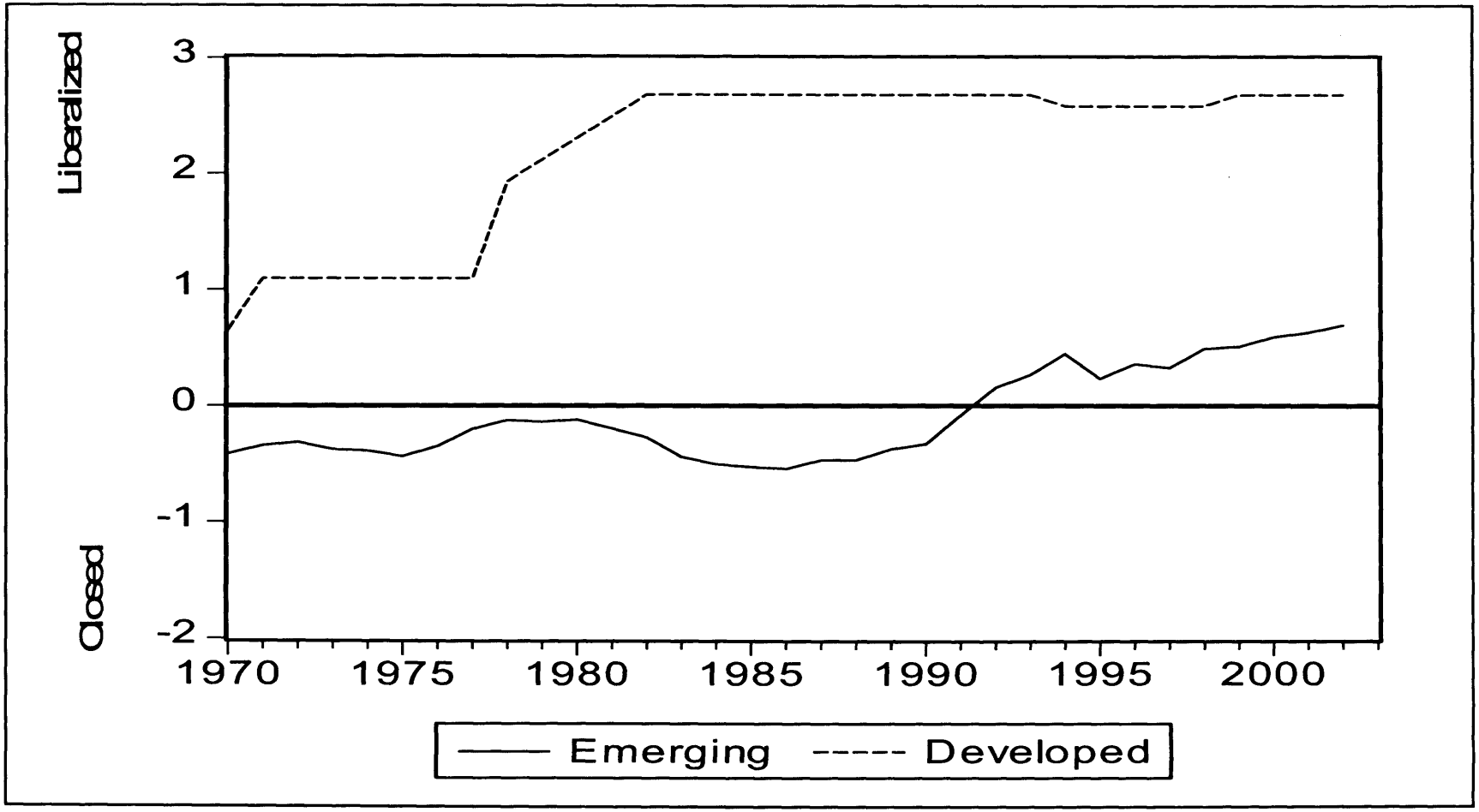
[c] The Jarque-Bera statistic summarizes the skewness and kurtosis and tests whether the series is normally distributed. * Indicates significant at the 5% level.

2.3.2 Liberalization Variable

In this research we utilised two liberalization variables, first the liberalization intensity index or the capital openness index KAOPEN developed by Chinn and Ito (2002) with an extension in Chinn and Ito (2005) and second the official liberalization dates compiled by Bekaert *et al.* (2005).[10]

The KAOPEN index is constructed based on four binary dummy variables that codify the tabulation of restrictions on cross-border financial transactions reported in the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER). The four major restriction categories on external accounts are: k_1, (a variable indicating the presence of multiple exchange rates), k_2, (a variable indicating restriction on current account transactions), k_3, (a variable indicating restrictions on capital account transactions) and k_4, (a variable indicating the requirement for the surrender of export proceeds). These four categories have been included in the construction of the KAOPEN index. In order to focus on the effect of financial liberalization rather than controls, they reverse the values of the binary variables of the AREAER series. Due to changes in the AREAER classification method after 1996, Chinn and Ito (2005) follow the extension of the four binary classifications developed by Mody and Murshid (2005). High positive values of the index indicate a higher degree of financial liberalization. Figure 2.2 shows the index of financial openness in emerging markets and developed markets from 1970 to 2003. On average, most of the emerging markets are relatively open after the 1990's compared to developed markets, which have been well opened since before the 1970's.

[10] Details discussion on the official liberalization dates data can be found in Appendix B of Bekaert and Harvey (2000) and Appendix A of Bekaert et al. (2005).



Note: An average sum of the liberalization index derived from Chinn and Ito (2005) KAOPEN index.

Figure 2.2: Aggregate of KAOPEN Liberalization Index for Developed and Emerging Markets from 1970 to 2003

KAOPEN is a *de jure* measure of financial openness because it attempts to comprehensively proxy the extent and nature of the regulatory restrictions on external account transactions. Although the KAOPEN index does not specifically measure stock market liberalization intensity, it represents a good proxy as it measures the restrictions on capital account and current account transactions using its k_2 and k_3 components, respectively. Due to these advantages we decided to use KAOPEN instead of the Edison and Warnock (2003) capital account liberalization intensity index.

2.3.3 Economic Fundamentals

We considered four economic fundamentals variables: the domestic interest rate, inflation, the real exchange rate and a political stability index. The political stability index is obtained from the ICRG database to represent a country's domestic political stability. The index ranges from zero to one hundred, where zero indicates the highest risk and a hundred the lowest. A country that has a higher political stability index is expected to attract more investment, which may lead to higher returns and lower volatility. The interaction sign with returns is expected to be positive.

The real exchange rate is defined as the nominal exchange rate adjusted by the ratio of the U.S. prices to the domestic price level with 2000 as a base year.[11] A decrease in real exchange rate, therefore, is interpreted as a depreciation of the real exchange rate. Various domestic interest rates are also included in the estimation of mean equation to provide an alternative investment to stocks (the details of type of

[11] Real Exchange Rate (R) = $e(\frac{P_f}{P_d})$, where e is nominal exchange rate, P is price level, and subscript f and d refer to foreign and domestic respectively.

interest rate used are in Appendix 2). Higher domestic interest rates will decrease the demand for stocks and hence stock returns. We expect a negative coefficient for the interest rate and real exchange rate, and a positive coefficient for inflation. The growth rate percentage for each of the economic fundamentals variables is calculated and extracted from the International Financial Statistics (IFS) database.

2.4 Methodology

In this study, to analyse the relationship between stock returns (y_t), volatility of stock returns (h_t) and financial liberalization (X_t), we use models from the Generalized Autoregressive Conditional Heteroskedasticity (GARCH) family including ARCH, GARCH, EGARCH and TGARCH. The advantages of a GARCH specification are that it allows capturing volatility clustering, leptokurtosis and skewness along with the leverage effect in stock returns. This approach has been widely employed in the area of conditional volatility modelling in emerging markets by, among others, Koot and Padmanabhan (1993), De Santis and Imrohoroglu (1997), Kwan and Reyes (1997), Kim and Singal (2000) and Jayasuriya (2005). Variations of the methodology including the Semi Parametric Autoregressive Conditional Heteroskedasticity (SPARCH) model of Bekaert and Harvey (1997), and Shin (2005) could be traced in the literatures of volatility and after-liberalization effects.

First, we start to model the conditional volatility of stock returns and liberalization by specifying the conditional mean equation:

$$y_t = u_t + \varepsilon_t \qquad \text{Equation 2.1}$$

$$u_t = \alpha + \beta_i Z_t + \sum_{i=1}^{m} \varphi_i y_{t-i} + \sum_{j=i}^{n} \pi_j \varepsilon_{t-j} \qquad \text{Equation 2.2}$$

where y_t is the stock return of emerging markets with conditional mean μ_t. The conditional distribution of ε_t is assumed to be $\sqrt{h_t}v_t$ and has the property of $\varepsilon_t \sim N(0,h_t)$ with is h_t conditional variance. The v_t is an *i.i.d.* sequence with zero mean and unit variance. The vector Z_t, of Equation 2.2 includes foreign stock returns (S&P500 index, FTSE100 index and Nikkei225 index), growth rates of domestic economic fundamentals (inflation, interest rate, real exchange rate and the political stability index) and the time trend. The mean equation estimation also captures the long-term trend (if any) by adding a time trend variable. The model tries to correct for ARMA terms in the errors by augmenting the model with an ARMA component in the mean equation. The BIC criterion is used to select of the ARMA Lags. A zero-one dummy crisis variable has been considered in the model for countries that have been affected by economic crisis within the sample period. Those countries are Argentina, Brazil, Malaysia, Mexico, Thailand, Philippines, and Indonesia.

Second, we consider the best fitted GARCH-type model including ARCH(q), GARCH(p,q), EGARCH(p,q) or TGARCH(p,q) in the selection process of the appropriate models. The best model is chosen based on the best information given by the Schwarz Information Criteria (BIC). Bolleslev and Wooldrige (1992) quasi-maximum likelihood (QML) covariances and robust standard errors are used in the estimation processes. A heteroskedasticity consistent covariance methodology is used to capture the problem of non-normality of the standardized residuals which results in the parameter estimates still being consistent.

$$h_t = \omega + \sum_{i=1}^{p} \zeta_i h_{t-i} + \sum_{j=1}^{q} \eta_j \varepsilon_{t-j}^2 + \kappa \varepsilon_{t-1}^2 d_{t-1} + \lambda X \qquad \text{Equation 2.3}$$

where $d_{t-1} = 1$ if $\varepsilon_{t-1} < 0$ and otherwise.

Equation 2.3 specifies the general TGARCH(p,q) model. Equation 2.3 can also be an ARCH(q) model if both p and κ are set to zero or a GARCH(p,q) model if κ is set to zero. Good news, $\varepsilon_{t-j} > 0$, has an impact of (η_j), while bad news, $\varepsilon_{t-j} < 0$, has an impact of ($\eta_j + \kappa_j$). If κ is positive then bad news will increases volatility, which implies that there is a leverage effect for the j-th order. If κ is significantly different from 0 then the news impacts are asymmetric. Equation 2.4 meets the specification of the general EGARCH(p,q) model where the leverage effect is exponential, rather than quadratic.

$$log(h_t) = \omega + \sum_{i=1}^{p} \zeta_i \, log(h_{t-i}) + \sum_{j=i}^{q} \left[\eta_j \left| \upsilon_{t-j} \right| + \kappa_j (\upsilon_{t-j}) \right] + \lambda X$$

Equation 2.4

The left-hand side of Equation 2.4 ($\log(h_t)$) is the log of the conditional variance. The presence of leverage effects can be tested by the hypothesis that ($\kappa_j < 0$) and the impact is asymmetric if ($\kappa_j \neq 0$).

In order to examine the effect of market liberalization on the volatility of returns, what we are interested from the above conditional variance models (Equation 2.3 and 2.4) is the λ (estimated coefficient for liberalization variable X). Positive two (+2) has been added to the KAOPEN index to ensure positivity. If the coefficient λ, is significantly positive (negative), it is an indication of increase (decrease) in returns volatility during or following liberalization. For clarity and robustness checking, we considered two liberalization indices, the Chinn and Ito (2002) KAOPEN liberalization intensity index and the Bekaert *et al.* (2005) official liberalization date dummy. We name them as Model A for estimation that uses KAOPEN, Chinn and Ito (2002) index and Model B for estimation that uses the Bekaert *et al.* (2005) official liberalization dates. The Chinn and Ito (2002)

KAOPEN index ranges from January 1984 to December 2003, whilst the Bekaert *et al.* (2005) official liberalization date dummy is set equal to zero from the beginning of the sample to T, (and to one otherwise), where T is the official liberalization date.

2.5 Results

Our primary interest is on the coefficient of the liberalization variable, λ, in the conditional variance equations (Equations 2.3 and 2.4). Therefore, we are less concerned about other estimates including persistence and asymmetry parameters. However, the conditional mean estimates (Equation 2.2) show significant evidence that inflation, interest rate, real exchange rate, political stability and stock market returns in developed markets play a significant role in explaining returns in emerging countries. The results of the mean equation from various GARCH estimations for both Model A and B are not presented here. They are available in Appendix 3. This finding is consistent with Fama, (1981); Gultekin, (1983); and Muradoglu *et al.* (2001); who find a significant relationship between macroeconomic fundamental variables, the general political environment and the influence of developed markets on emerging market returns.

2.5.1 Model A (Chinn and Ito 2002, KAOPEN Liberalization Intensity Index)

As to the selection of the appropriate model, we experimented with the GARCH-type specification, including ARCH, GARCH, TGARCH and EGARCH models, based on the BIC criterion that fits the data best. The results suggest that (in both Models A and B) it is sufficient to use either ARCH(1), GARCH(1,1), EGARCH(1,1) and TGARCH(1,1). Apart from being commonly used in the existing literature, this specification has the desirable features of interpretability and good fit

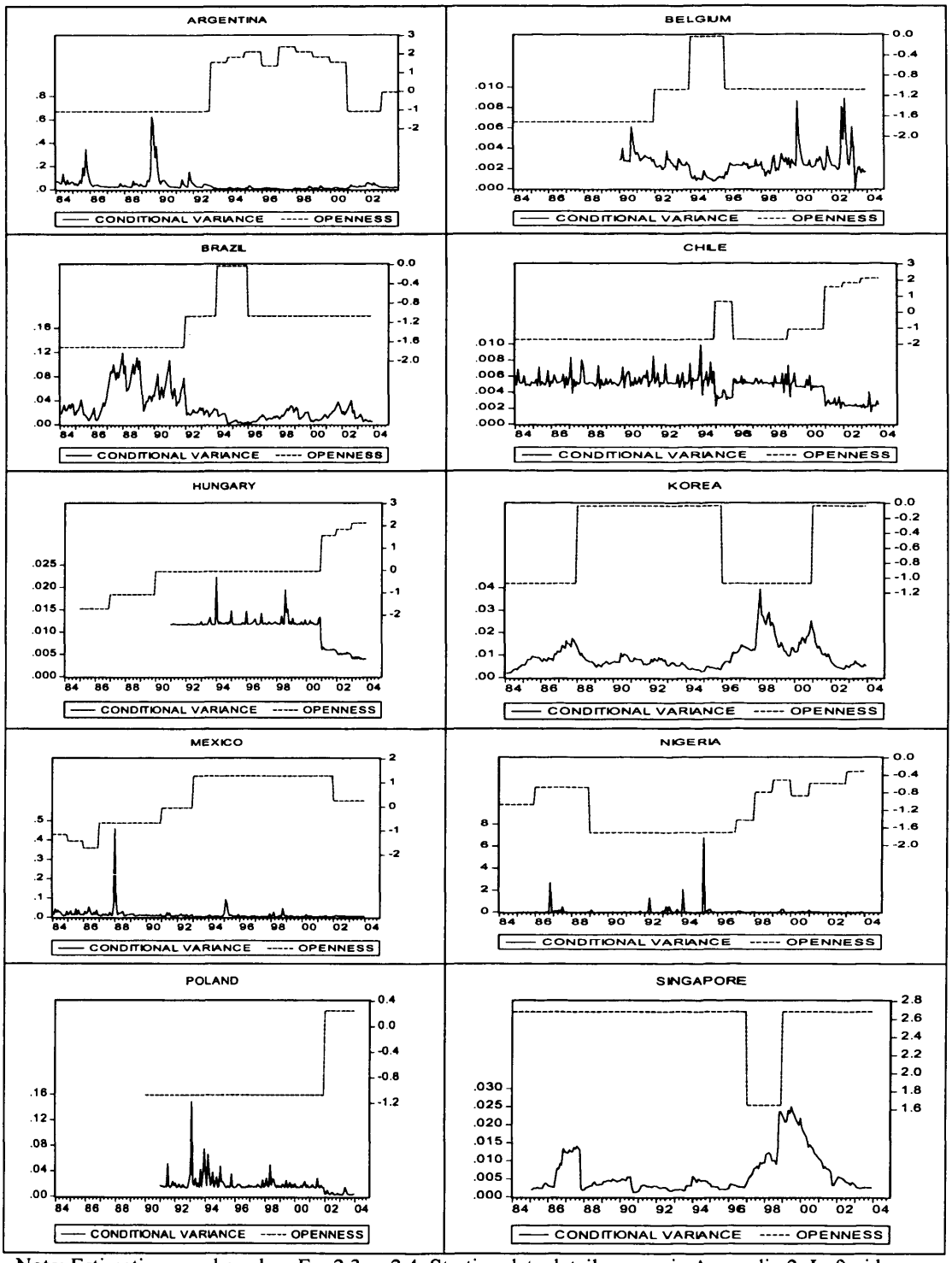
for the data. We found a significant effect of liberalization on volatility for several countries. Specifically, the findings of the liberalization effect can be grouped into three categories: decreasing, increasing and unchanging returns volatility. Based on Table 2.3, we observed a significantly negative liberalization coefficient, λ, at the 5% level for 10 countries (Argentina, Belgium, Brazil, Chile, Hungary, Korea, Mexico, Nigeria, Poland, and Singapore). This means that the intensifying liberalization is associated with decreasing volatility. The results also suggest a statistically significant positive coefficient between the openness variable and volatility for Colombia, Peru and Sri Lanka. Market returns linked with the liberalization programmes in these three countries seem to be more volatile. The other markets do not show a significant change in the level of volatility related to the liberalization variable.

Figures 2.3 and 2.4 show time series plots for the estimated conditional variance series and the openness index. It is clear that Figure 2.3 shows an opposite direction movement of conditional volatility and openness for countries that have a decreasing volatility impact in the process of liberalization. For instance in Argentina, after opening its capital market in 1993 the conditional variance series exhibited a rather substantial calm period right up to the beginning of the financial crisis in the early 2000s. The Figure also shows that before the opening of its capital market, returns were volatile. On the other hand, Figure 2.4 shows that the

Table 2.3: Conditional Variance Estimates of Model A[a] from 1984:01 to 2003:12

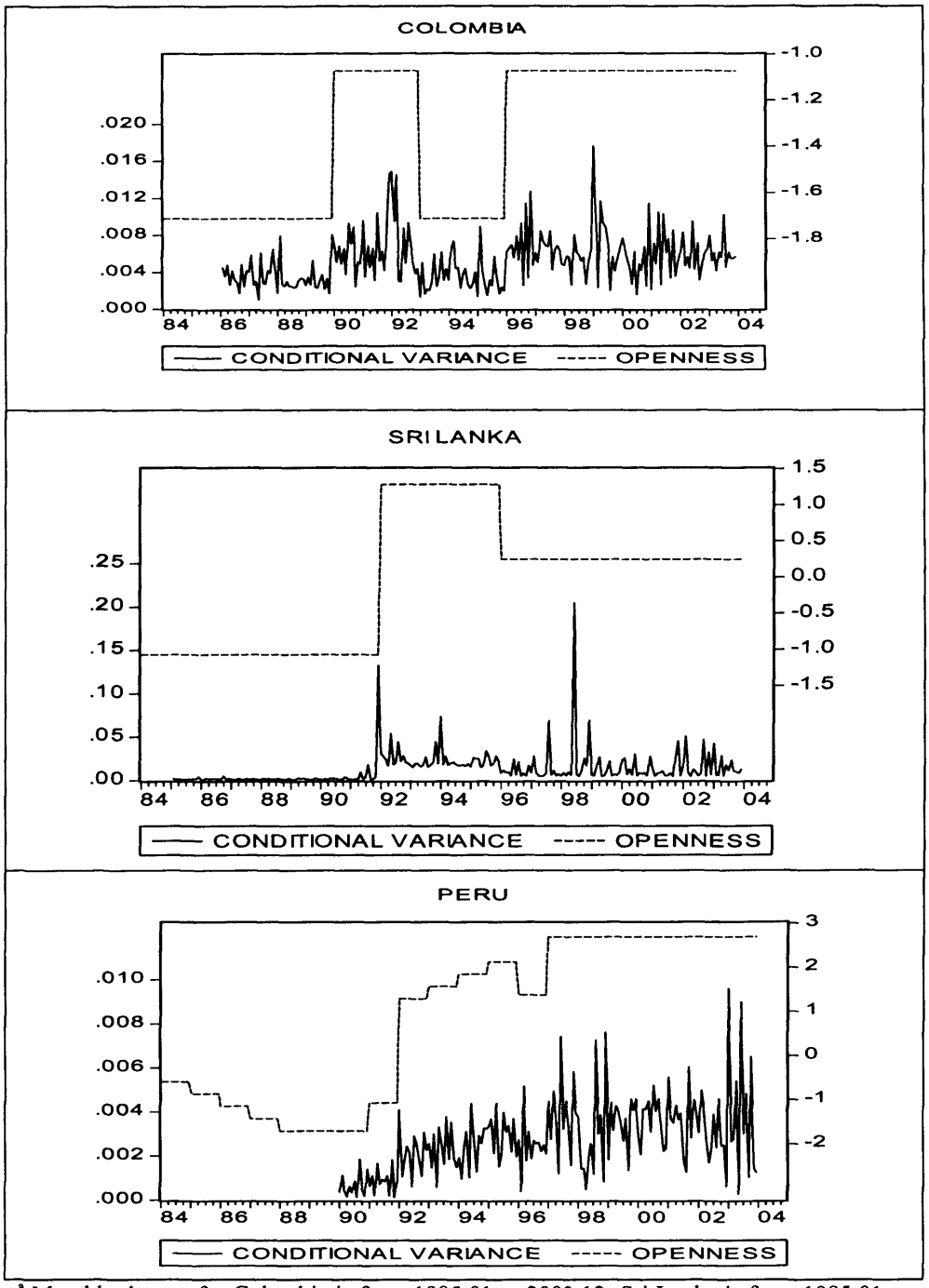
	ω	ζ_1	η_1	κ	λ	BIC	Q(12)[b]	$Q^2(12)$[c]	Skew	Kurt	J-B
ARG	0.012	0.112	0.634*	-0.449	-0.003*	-0.876	0.927	0.499	0.273	4.315	19.2*
	(0.004)	(0.145)	(0.285)	(0.299)	(0.001)						
AUS	0.002	0.741**	0.093	-0.023	-0.001	-2.279	0.721	0.598	-0.2	3.32	2.341
	(0.002)	(0.226)	(0.070)	(0.095)	(0.001)						
BAN	-1.272	0.952**	1.364**	0.396**	0.084	-2.453	0.798	0.261	-0.083	3.728	3.858
	(0.222)	(0.028)	(0.206)	(0.153)	(0.174)						
BEL	-11.997	-0.821**	-0.173	0.143	-1.06**	-2.729	0.958	0.13	-0.256	3.429	3.138
	(1.360)	(0.202)	(0.141)	(0.096)	(0.255)						
BRA	-0.344	0.920**	-.216*	-0.190**	-0.15**	-0.674	0.391	0.504	-0.479	3.769	14.53*
	(0.111)	(0.020)	(0.103)	(0.057)	(0.034)						
CHI	0.005	-0.337	0.152	-0.174	-0.001*	2.215	0.724	0.268	-0.064	2.447	3.22
	(0.001)	(0.326)	(0.103)	(0.100)	(0.000)						
COL	-8.516	-0.763**	0.096	0.198	0.886*	-2.126	0.577	0.558	0.28	3.261	3.41
	(0.928)	(0.121)	(0.097)	(0.109)	(0.505)						
HUN	0.012		0.061		-0.04**	-1.391	0.891	0.97	-0.132	4.694	19.12*
	(0.003)		(0.090)		(0.001)						
IND	0.003	0.540**	0.257	-0.318*	0.000	-1.819	0.423	0.673	0.042	2.795	0.466
	(0.001)	(0.167)	(0.135)	(0.129)	(0.001)						
INA	0.002	0.596**	0.548	-0.173	-0.000	-1.487	0.138	0.92	0.153	3.845	7.64*
	(0.002)	(0.161)	(0.647)	(0.696)	(0.001)						
ISR	-9.951	-0.754**	0.430**	-0.148**	0.220	-2.225	0.247	0.977	-0.161	3.757	6.69*
	(0.390)	(0.074)	(0.101)	(0.049)	(0.147)						
JOR	-1.819	0.716**	-0.012	0.222*	-0.013	-3.276	0.476	0.067	0.255	3.224	3.004
	(1.136)	(0.176)	(0.110)	(0.099)	(0.025)						
KEN	0.001		1.701**		0.000	-2.324	0.509	0.903	-0.101	3.754	4.277
	(0.000)		(0.282)		(0.000)						
KOR	0.005		0.237*		-0.003*	-1.694	0.68	0.21	0.168	3.317	2.134
	(0.001)		(0.117)		(0.001)						
MAL	-1.009	0.876**	0.515**	-0.113	-0.042	-2.021	0.44	0.544	0.013	3.4	1.209
	(0.420)	(0.070)	(0.155)	(0.128)	(0.032)						
MEX	0.006		0.548**		-0.02**	-1.543	0.204	0.064	-0.25	3.155	2.703
	(0.001)		(0.109)		(0.001)						
MOR	-4.098	0.410	0.189	0.216*	-0.047	-3.303	0.85	0.971	0.426	3.421	7.19*
	(2.122)	(0.318)	(0.132)	(0.093)	(0.180)						
NIG	-6.649	-0.021	0.153	0.231	-1.653*	-1.385	0.625	0.999	-0.507	5.847	73.1*
	(1.770)	(0.302)	(0.125)	(0.149)	(0.617)						
PAK	-1.171	0.952**	0.384**	0.129	-0.589	-2.018	0.429	0.551	0.256	4.034	12.82*
	(0.398)	(0.015)	(0.100)	(0.111)	(0.328)						
PER	-8.984	-0.404*	-0.736**	-0.468**	0.521**	-1.547	0.811	0.743	-0.379	3.474	5.60
	(1.165)	(0.173)	(0.198)	(0.098)	(0.081)						
PHI	-1.231	0.766**	0.110	-0.004	-0.087	-1.579	0.44	0.755	0.234	4.333	19.06*
	(0.872)	(0.175)	(0.151)	(0.102)	(0.081)						
POL	-5.304	0.062	0.327		-0.86**	-0.901	0.077	0.448	0.101	4.178	10.01*
	(0.278)	(0.146)	(0.191)		(0.295)						
SAF	-1.420	0.788**	0.364**	-0.144	0.015	-1.984	0.796	0.647	-0.388	3.351	5.82
	(0.707)	(0.129)	(0.135)	(0.093)	(0.083)						
SIN	0.002	0.953**	0.146**	-0.206**	-0.000*	-2.157	0.065	0.257	-0.139	3.554	3.638
	(0.001)	(0.025)	(0.052)	(0.073)	(0.000)						
SPA	0.000	0.648**	0.287**		0.000	-2.407	0.123	0.213	0.087	3.271	1.03
	(0.000)	(0.096)	(0.088)		(0.000)						
SRI	-5.641		0.887**	0.174	1.113**	-2.225	0.628	0.201	-0.215	4.68	20.94*
	(0.212)		(0.210)	(0.154)	(0.121)						
THA	0.023	0.607**	0.383	-0.165	0.619	-1.6	0.572	0.87	0.073	4.536	22.63*
	(0.018)	(0.155)	(0.257)	(0.274)	(0.489)						
TUR	0.040	-0.084	-0.109**		0.005	-0.313	0.341	0.542	0.292	3.295	3.791
	(0.018)	(0.492)	(0.022)		(0.006)						
VEN	0.015	-0.178**	0.760**	-0.732**	0.001	-0.816	0.587	0.978	-0.431	4.453	26.78*
	(0.003)	(0.062)	(0.223)	0.219	(0.002)						
ZIM	-3.881	0.957**	0.184	-0.066	-2.064	-1.346	0.628	0.201	-0.215	4.68	20.94*
	(2.044)	(0.041)	(0.109)	(0.067)	(1.091)						

[a]Variant of Eq. 2.4, $h_t = \omega + \sum_{i=1}^{p} \zeta_i h_{t-i} + \sum_{j=1}^{q} \eta_i \varepsilon_{t-j}^2 + \kappa \varepsilon_{t-1}^2 d_{t-1} + \lambda X$ or alternatively Eq. 2.3 for EGARCH is chosen based on the minimum BIC. The numbers in parentheses are Bollerslev-Wooldrige robust standard errors. [b] and [c] indicate the *p*-value of the *Q*-statistic for the standardized residuals and squared standardized residuals at lag 12 that tests the null hypothesis of no autocorrelation in the series up to order 12 respectively. *, ** indicate significance at the 5% or 1% level, respectively.



Note: Estimations are based on Eq. 2.3 or 2.4. Starting date details are as in Appendix 2. Left-side vertical axis is the variance and right-side vertical axis is the KAOPEN index scale.

Figure 2.3: Conditional Variance and KAOPEN Index for Decreasing Volatility



[a] Monthly dataset for Colombia is from 1985:01 to 2003:12; Sri Lanka is from 1985:01 to 2003:12; and Peru is from 1991:01 to 2003:12. Left-side vertical axis is the variance and the right-side vertical axis is the KAOPEN index scale.

Figure 2.4: Conditional Variance and KAOPEN Index for Increasing Volatility [a]

conditional volatility moves together with the openness index for countries that experience an increased volatility effect in the process of liberalization. In the case of Peru, returns seem to be more volatile starting from 1992. In that year, a decree on the Private Sector Investment Guarantee Regime was enacted, under which the rights and guarantees that were accorded to domestic investors were extended to foreign investors. Therefore, we could conclude that the effect of liberalization on volatility in emerging countries, whether increasing, decreasing or unchanged, varies depending on the country being studied. These results are consistent with previous findings such as Kassimatis (2002) and Jayasuriya (2005) and our prior expectations.

Further, we investigated the question of why volatility effects behave differently in emerging countries. We aim at answering this question by examining market characteristics and the quality of institutions for different countries in our sample. If we look closer at the market characteristics and institutional quality of these markets, we find that countries that have a significantly negative coefficient estimate for the liberalization variable are those that on average are characterized by higher institutional quality and better market characteristics. Table 2.4 presents the calculation of arithmetic mean (average) for the institutional quality indices when we group countries based on the estimated coefficient of the KAOPEN index. Countries with a negative liberalization volatility relationship are, generally, characterized by lower institutional quality compared to the countries that are experiencing decreased or at least unchanged volatility. The averages of Law and Order, Bureaucratic Quality and Corruption indices for Colombia, Peru and Sri Lanka are 1.90, 2.05 and 2.71 respectively compared to the group with decreasing volatility where the

Table 2.4: Various Indices of Institutional Quality for Model A

Country	Law & Order[a]	Bureaucratic Quality[b]	Corruption[c]
Decreasing volatility			
Argentina	3.68	2.35	3.18
Belgium	5.61	4.00	4.55
Brazil	3.06	2.65	3.41
Chile	4.51	2.43	3.30
Hungary	5.01	3.38	4.33
Korea	3.97	2.61	3.73
Mexico	4.17	2.13	2.81
Nigeria	2.46	1.96	1.99
Poland	5.12	2.67	4.52
Singapore	3.67	3.14	3.19
Average	**4.13**	**2.73**	**3.50**
Increasing volatility			
Colombia	1.32	2.62	2.67
Peru	2.31	1.52	2.21
Sri Lanka	2.06	2.00	3.26
Average	**1.90**	**2.05**	**2.71**
Unchanged Volatility			
Austria	6.00	3.89	4.81
Bangladesh	1.77	0.92	1.11
India	3.13	2.97	2.63
Indonesia	2.68	1.31	1.43
Israel	3.91	3.53	4.32
Jordan	3.40	2.15	3.34
Kenya	3.02	2.57	2.70
Malaysia	2.78	2.24	2.77
Morocco	2.05	1.28	1.67
Pakistan	2.13	1.35	2.92
Philippines	4.52	2.44	3.97
South Africa	2.36	3.23	4.42
Spain	4.61	3.22	4.15
Thailand	4.17	2.83	2.61
Turkey	3.57	2.27	2.73
Venezuela	3.65	1.65	2.76
Zimbabwe	2.48	2.44	2.42
Average	**3.31**	**2.37**	**2.99**

[a, b, c] Average of monthly ICRG index from 1984 to 2004.

averages are 4.13, 2.73 and 3.50 respectively. The average of the institutional quality index for countries that show no dependence between volatility and liberalization is also much higher than for the increasing volatility countries. *Ceteris paribus*, this could be a strong indication that high-quality institutions in emerging countries would efficiently absorb undue volatility originating from the liberalization process.

Good market characteristics are also likely to play a relevant role in determining the effect of financial liberalization on volatility. Table 2.5 reports the market transparency, investment climate and capital/income exit openness index. On average, countries with a decreasing volatility effect have higher accounting standards, investment profile, economic risk, and capital /income exit openness index compared to the countries that have an increasing volatility effect. The unchanged volatility countries also show a higher market characteristics index as compared to the increasing volatility effect countries. The details are illustrated in Table 2.5. The findings further support the hypothesis that market transparency, investment climate and the restriction of income or capital repatriation could be an explanation for different volatility effects in the process of emerging market liberalization. This is because favourable market characteristics would be a strong basis for the efficient functioning of the equity market. The findings are consistent with most of the current literature, like Bekaert and Harvey (2000), Stiglitz (1999 and 2000), La Porta (1998) and Johnson *et al.* (2000) that emphasise the importance of good financial market systems and institutions to ensure that the liberalization process in emerging markets a successful story.

Table 2.5: Various Indices of Market Characteristics for Model A

Country	Accounting Standards		Investment Profile[c]	Economic Risk[d]	Exit Openness[e]	
	La Porta[a]	Jayasuriya[b]			Income	Capital
Decreasing volatility						
Argentina	45	2.0	5.57	54.4	4.77	4.77
Belgium	61	3.0	8.64	76.6	0.00	0.00
Brazil	54	3.0	6.04	56.7	5.00	5.00
Chile	52	2.0	7.68	61.6	5.00	4.33
Hungary	-	2.0	7.73	-	5.00	5.00
Korea	62	3.0	7.24	64.4	5.00	4.67
Mexico	60	3.0	6.81	56.1	4.70	4.70
Nigeria	59	2.0	4.95	54.5	4.70	4.70
Poland	-	2.0	7.84	-	5.00	5.00
Singapore	78	3.0	7.82	82.2	5.00	5.00
Average	**58.87**	**2.5**	**7.03**	**63.31**	**4.91**	**4.81**
Increasing volatility						
Colombia	50	2.0	6.31	59.5	4.73	4.73
Peru	38	2.0	6.28	50.4	5.00	5.00
Sri Lanka	-	-	6.52	59.3	3.00	3.67
Average	**44**	**2.0**	**6.37**	**56.4**	**4.46**	**4.46**
Unchanged Volatility						
Austria	54	3.0	9.12	80.6	-	-
Bangladesh	-	-	5.19	59.2	5.00	5.00
India	57	2.7	6.33	60.9	5.00	5.00
Indonesia	-	1.0	6.12	66.4	3.00	3.00
Israel	64	3.0	6.88	64.4	-	-
Jordan	-	2.0	6.73	70.1	5.00	5.00
Kenya	-	-	6.69	55.5	5.00	5.00
Malaysia	76	3.0	7.68	78.8	4.90	4.90
Morocco	-	-	5.40	63.1	4.40	4.40
Pakistan	-	2.0	6.10	63.4	5.00	5.00
Philippines	65	3.0	7.25	59.4	-	5.00
S. Africa	70	3.0	7.51	69.1	5.00	5.00
Spain	64	3.0	8.92	73.3	-	-
Thailand	64	2.0	7.01	74.4	5.00	5.00
Turkey	51	2.0	6.41	56.4	5.00	5.00
Venezuela	40	2.0	5.31	63.2	4.43	4.43
Zimbabwe	-	2.0	4.54	51.7	5.00	5.00
Average	**60.5**	**2.41**	**6.66**	**65.27**	**4.77**	**4.74**

[a]Accounting Standards Index created by examining and rating companies' annual reports on their inclusion or omission of 90 items taken from La Porta (1998). [b, e] Index adapted from Jayasuriya (2005) Appendix A with extension years from various issue of Emerging Stock Market Factbook. [c, d] An average of the sub-component of the monthly ICRG political stability index from 1984 to 2004. [f] an economic risk rating ~ <50 is very high risk. 50-60 range is high risk. 60-70 range is moderate risk. 70-80 range is low risk and the 80-100 range is very low risk. Taken from Bekaert *et.al.* (2005). – indicates that data are not available.

2.5.2 Model B (Bekaert et al. 2005, Official Liberalization Dates Dummy)

To check the robustness of our analysis based on the KAOPEN liberalization intensity index, we also consider an alternative proxy given by the official liberalization date. We follow Bekaert and Harvey (2000) in defining the official liberalization dates to investigate its effect on stock returns volatility. They define official liberalization as the introduction of either American Depositary Receipt (ADR) or country-specific funds in the US market, official regulatory changes that allow for foreign ownership in the local stock market or capital flow break point. Austria, Belgium, Hungary, Poland, and Singapore are not included in the analysis as they are considered to have been open markets long before the beginning of the sample period. The estimations use the same emerging stock market returns and control variables in the mean equation but use the official liberalization dummy instead of the openness intensity index in the variance equation. The estimation is from T - 60 to T + 60, where T is the month of the official liberalization. The dates included in the analysis are shown in Appendix 2.

Referring to λ in Table 2.6, we noticed that countries like Bangladesh, Colombia, Morocco, Peru, Sri Lanka, Thailand and Venezuela experienced higher volatility after the liberalization date while for Brazil, Malaysia, Mexico and Nigeria the opposite holds. No statistical evidence could be detected that any changes occur in the volatility in pre- and post-liberalization periods for the rest of the countries. Model B results are consistent with the openness intensity index analysis (Model A) where countries like Brazil, Malaysia, Mexico and Nigeria have better market characteristics (62.3 for La Porta accounting standards, 2.75 for Jayasuriya accounting standards, 6.37 for investment profile, 61.5 for economic risk and 4.83 for exit openness of both income and capital) and higher institutional quality (3.12

points for law and order, 2.25 for bureaucratic quality and 2.75 for corruption) compared to Bangladesh, Colombia, Morocco, Peru, Sri Lanka, Thailand and Venezuela which have 48 for La Porta accounting standards, 2.0 for Jayasuriya accounting standards, 6.00 for investment profile, 61.3 for economic risk and 4.51 and 4.60 for exit openness of income and capital respectively and 2.48 points for law and order, 1.83 for bureaucratic quality and 2.33 for corruption respectively. Tables 2.7 and 2.8 show the details of the average of the institutional quality and market characteristics index for Model B. Furthermore, the inclusion of the official liberalization dummy variable has induced more countries to experience more volatile stock returns in the after-liberalization period compared to Model A. In Model A, only three countries experience positive and statistically significant λ coefficients compared to 7 countries in Model B. This could be due to the relatively shorter sample period used to estimate Model B, i.e. five years before and after liberalization. The results are not surprising because as Kaminsky and Schumkler (2003) also find, volatility may increase for the first 4 to 5 years after liberalization before reaching a calm and tranquil period in the longer time horizon.

Table 2.6: Conditional Variance Estimates for Model B[a, b]

	ω	ζ_1	η_1	κ	λ	BIC	Q(12)[c]	Q^2(12)[d]	Skew	Kurt	J-B
ARG	0.008	0.067	1.017**	0.624	-0.002	-0.313	0.456	0.346	0.506	4.062	10.68*
	(0.003)	(0.079)	(0.368)	(0.856)	(0.003)						
BAN	-4.251	0.232	-0.830*	0.836**	0.692*	-1.837	0.715	0.141	0.429	5.192	17.32*
	(1.332)	(0.233)	(0.324)	(0.244)	(0.292)						
BRA	0.010	0.891**	-0.039	-0.071	-0.07**	-0.106	0.363	0.566	-0.043	3.636	2.082
	(0.003)	(0.082)	(0.055)	(0.093)	(0.002)						
CHI	0.003	0.187	0.203	-0.202	-0.001	-1.959	0.617	0.210	0.233	2.287	3.592
	(0.002)	(0.621)	(0.152)	(0.162)	(0.007)						
COL	-2.194	0.542**	-0.678**	0.705**	0.218*	-2.190	0.781	0.125	0.223	3.069	1.017
	(0.530)	(0.095)	(0.154)	(0.108)	(0.101)						
IND	-1.813	0.665**	0.344*	0.170	-0.150	-1.372	0.107	0.660	0.255	2.964	1.302
	(0.909)	(0.186)	(0.172)	(0.090)	(0.157)						
INA	-0.494	0.818**	-0.556*	-0.370**	-0.085	-1.579	0.136	0.286	0.150	3.186	0.598
	(0.248)	(0.050)	(0.223)	(0.098)	(0.050)						
ISR	-1.887	0.741**	0.452*	0.064	0.216	-2.080	0.112	0.381	-0.421	3.219	3.770
	(1.098)	(0.169)	(0.214)	(0.120)	(0.202)						
JOR	0.000	0.944**	0.061	-0.246**	-0.000	-3.287	0.542	0.628	0.311	2.555	2.908
	(0.000)	(0.090)	(0.059)	(0.076)	(0.000)						
KEN	-4.680	0.367**	1.724**	-0.394*	-0.455	-2.178	0.146	0.135	0.245	3.390	1.950
	(0.564)	(0.092)	(0.270)	(0.163)	(0.296)						
KOR	0.000	1.015**	-0.229	0.216	0.000	-1.952	0.582	0.172	-0.167	2.369	2.293
	(0.000)	(0.081)	(0.143)	(0.154)	(0.001)						
MAL	-0.685	0.898**	0.312*	0.179	-0.187*	-2.024	0.457	0.899	-0.287	3.765	4.279
	(0.505)	(0.079)	(0.154)	(0.101)	(0.091)						
MEX	-6.110	-0.658**	-0.477	-0.439**	-2.89**	-1.293	0.116	0.487	-0.398	2.928	3.170
	(0.547)	(0.108)	(0.324)	(0.112)	(0.423)						
MOR	-7.162	0.161	0.522	-0.073	1.554**	-3.697	0.312	0.074	0.006	2.917	0.034
	(1.847)	(0.228)	(0.371)	(0.205)	(0.540)						
NIG	-3.371	-0.072	-0.711*	0.834**	-2.459*	-1.292	0.986	0.903	-0.695	7.730	120.5**
	(0.967)	(0.144)	(0.344)	(0.304)	(0.336)						
PAK	0.000	0.858**	0.091	-0.270*	0.001	-2.267	0.831	0.798	0.404	2.917	3.298
	(0.000)	(0.092)	(0.064)	(0.137)	(0.000)						
PER	-8.957	0.126	1.195**	-0.416**	2.637**	-3.785	0.150	0.058	-0.164	3.787	3.618
	(1.497)	(0.163)	(0.294)	(0.148)	(0.690)						
PHI	0.000	0.985**	0.018	-0.142**	-0.000	-1.393	0.845	0.719	0.124	3.823	3.695
	(0.000)	(0.077)	(0.062)	(0.052)	(0.000)						
SAF	0.001	0.306*	0.435	-0.117	-0.000	-2.221	0.453	0.820	-0.435	3.443	4.735
	(0.001)	(0.153)	(0.326)	(0.302)	(0.001)						
SPA	0.002	0.429	0.061	-0.143	-0.000	-1.768	0.099	0.998	-0.203	4.036	3.873
	(0.002)	(0.616)	(0.177)	(0.105)	(0.000)						
SRI	-1.522	0.807**	0.170	0.265**	0.407*	-2.654	0.882	0.487	0.250	2.992	1.276
	(0.527)	(0.067)	(0.142)	(0.102)	(0.165)						
THA	-9.751	-0.384**	1.271**	0.173	1.534**	-2.077	0.962	0.983	-0.492	3.166	4.236
	(0.699)	(0.098)	(0.199)	(0.145)	(0.456)						
TUR	-0.365	0.846**	-0.210*	0.237**	0.002	-0.024	0.374	0.383	0.309	2.661	2.035
	(0.309)	(0.065)	(0.106)	(0.087)	(0.053)						
VEN	-7.731	-0.577**	0.179	0.579**	1.054**	-0.969	0.659	0.769	-0.275	3.137	1.597
	(0.577)	(0.121)	(0.155)	(0.137)	(0.339)						
ZIM	-1.697	0.726**	0.300	-0.161	0.257	-1.489	0.696	0.430	-0.306	3.370	2.538
	1.219	(0.214)	(0.162)	(0.130)	(0.237)						

[a] Variant of Eq. 2.4, $h_t = \omega + \sum_{i=1}^{p} \zeta_i h_{t-i} + \sum_{j=1}^{q} \eta_j \varepsilon_{t-j}^2 + \kappa \varepsilon_{t-1}^2 d_{t-1} + \lambda X$ or alternatively Eq. 2.3 for EGARCH is chosen based on BIC. [b] Sample periods are determined using monthly observation of T-60 and T+60 where T is Bekaert *et al.* (2005) official liberalization date. Numbers in parentheses are Bollerslev-Wooldrige robust standard errors. [c] and [d] indicate the p-value of the Q-statistic for the standardized residuals and squared standardized residuals at lag 12 that tests the null hypothesis of no autocorrelation in the series up to order 12 respectively. *, ** indicate significant at the 5% or 1% level, respectively.

Table 2.7: Various Indices of Institutional Quality for Model B

Country	Law & Order[a]	Bureaucratic Quality[b]	Corruption[c]
Decreasing volatility			
Brazil	3.06	2.65	3.41
Malaysia	2.78	2.24	2.77
Mexico	4.17	2.13	2.81
Nigeria	2.46	1.96	1.99
Average	**3.12**	**2.25**	**2.75**
Increasing volatility			
Bangladesh	1.77	0.92	1.11
Colombia	1.32	2.62	2.67
Morocco	2.05	1.28	1.67
Peru	2.31	1.52	2.21
Sri Lanka	2.06	2.00	3.26
Thailand	4.17	2.83	2.61
Venezuela	3.65	1.65	2.76
Average	**2.48**	**1.83**	**2.33**
Unchanged Volatility			
Argentina	3.68	2.35	3.18
Chile	4.51	2.43	3.30
India	3.13	2.97	2.63
Indonesia	2.68	1.31	1.43
Israel	3.91	3.53	4.32
Jordan	3.40	2.15	3.34
Kenya	3.02	2.57	2.70
Korea	3.97	2.61	3.73
Pakistan	2.13	1.35	2.92
Philippines	4.52	2.44	3.97
South Africa	2.36	3.23	4.42
Spain	4.61	3.22	4.15
Turkey	3.57	2.27	2.73
Zimbabwe	2.48	2.44	2.42
Average	**3.43**	**2.49**	**3.23**

[a, b, c] Average of monthly ICRG political stability index from 1984 to 2004.

Table 2.8: Various Indices of Market Characteristics for Model B

Country	Accounting Standards		Investment Profile[c]	Economic Risk[d]	Exit Openness[e]	
	La Porta[a]	Jayasuriya[b]			Income	Capital
Decreasing volatility						
Brazil	54	3.0	6.04	56.7	5.00	5.00
Malaysia	76	3.0	7.68	78.8	4.90	4.90
Mexico	60	3.0	6.81	56.1	4.70	4.70
Nigeria	59	2.0	4.95	54.5	4.70	4.70
Average	**62.3**	**2.75**	**6.37**	**61.5**	**4.83**	**4.83**
Increasing volatility						
Bangladesh	-	-	5.19	59.2	5.00	5.00
Colombia	50	2.0	6.31	59.5	4.73	4.73
Morocco	-	-	5.40	63.1	4.40	4.40
Peru	38	2.0	6.28	50.4	5.00	5.00
Sri Lanka	-	-	6.52	59.3	3.00	3.67
Thailand	64	2.0	7.01	74.4	5.00	5.00
Venezuela	40	2.0	5.31	63.2	4.43	4.43
Average	**48**	**2.0**	**6.00**	**61.3**	**4.51**	**4.60**
Unchanged Volatility						
Argentina	45	2.0	5.57	54.4	4.77	4.77
Chile	52	2.0	7.68	61.6	5.00	4.33
India	57	2.7	6.33	60.9	5.00	5.00
Indonesia	-	1.0	6.12	66.4	3.00	3.00
Israel	64	3.0	6.88	64.4	-	-
Jordan	-	2.0	6.73	70.1	5.00	5.00
Kenya	-	-	6.69	55.5	5.00	5.00
Korea	62	3.0	7.24	64.4	5.00	4.67
Pakistan	-	2.0	6.10	63.4	5.00	5.00
Philippines	65	3.0	7.25	59.4	-	5.00
S. Africa	70	3.0	7.51	69.1	5.00	5.00
Spain	64	3.0	8.92	73.3	-	-
Turkey	51	2.0	6.41	56.4	5.00	5.00
Zimbabwe	-	2.0	4.54	51.7	5.00	5.00
Average	**58.8**	**2.36**	**6.71**	**62.2**	**4.79**	**4.73**

[a] Accounting Standards Index created by examining and rating companies' annual reports on their inclusion or omission of 90 items taken from La Porta (1998). [b, e] Index adapted from Jayasuriya (2005) Appendix A with extension years from various issue of Emerging Stock Market Factbook. [c, d] An average of the sub-component of the monthly ICRG political stability index from 1984 to 2004. [f] an economic risk rating ~ <50 is very high risk. 50-60 range is high risk. 60-70 range is moderate risk. 70-80 range is low risk and the 80-100 range is very low risk. Taken from Bekaert *et. al.* (2005). – indicates that the data are not available.

2.5.3 Specification Tests

Overall, both Models A and B are statistically valid for making inferences about the relationship between volatility and liberalization. The choice of a GARCH-type model was based on the minimum of values of the BIC criterion, (see Tables 2.3 and 2.6). The specification tests in Tables 2.3 and 2.6 reveal that most of the residuals series are normally distributed except for a few countries where the Jarque-Bera normality test is still rejected at the 5 percent significance level. However the Jarque-Bera, skewness and kurtosis statistics show considerable improvement in normality compared to pre-estimation statistics (see Table 2.2).

Further, we computed the Ljung-Box (Q) statistic on the standardized residuals and squared standardized residuals to test the null hypothesis of no autocorrelation up to the twelfth order. This test is an alternative to the Langrange Multiplier test proposed by Engle (1982) to evaluate the specification of a GARCH process. Bollerslev and Mikkelsen (1994) show that this test has considerably more power in detecting the model misspecification. Both Q(12) and $Q^2(12)$ for all markets suggest no serial correlation in the standardized and squared standardized residuals.

2.6 Conclusion

This paper investigates the effect of financial liberalization on the volatility of stock returns for 30 emerging markets. We consider models from the GARCH family, including ARCH, GARCH, EGARCH and TGARCH models that have been well developed to model the dynamics of the conditional variance. The empirical analysis suggests a significant relationship between volatility and financial

liberalization. We also explain our findings by considering the different levels of institutional quality and market characteristics in those markets.

We can make several points by comparing our results with those of the existing literature. First, the results show that the level of volatility in emerging markets could increase, decrease, or remain unchanged in the post-liberalization era. This could be due to the heterogeneous nature of the emerging market economies. The results are consistent with previous findings such as Kassimatis (2002) and Jayasuriya (2005).

Second, heterogeneity of volatility effects in the process of liberalization across emerging markets has attracted a lot of attention; see Bekaert and Harvey (2000), Kaminsky and Schmukler (2003), Jayasuriya (2005) and Caner and Onder (2005) among others. We further investigated differences in market characteristics (accounting standards, investment profile, economic risk and restrictions on repatriation of income and capital) and institutional quality (law and order, level of corruption, and bureaucratic quality) that could be the possible reasons underlying the differences of the volatility effects in the liberalization process. Based on the visual observation our findings show that countries with a higher quality of institutions and better market characteristics tend to experience decreased, or at least unchanged, volatility effects after liberalization. On the other hand, countries with a low quality of markets and institutions suffer from more volatile markets in and after the liberalization process. Employing both the openness intensity index and the official liberalization dates dummy provides consistent results. The findings therefore could be a guide to the importance of high quality institutions and market characteristics in absorbing undue volatility that could attract more international funds.

Third, comparing the results of Model A and Model B could further reveal the nature of the volatility effects after the liberalization process. Returns tend to be more volatile right after the liberalization process and to reach a calm and tranquil period approximately 4 to 5 years after the liberalization. The results are consistent with the famous phrase of Kaminsky and Schmukler (2003): 'short run pain, long run gain'.

Finally, the results could also support the importance of proper sequencing in the liberalization process. McKinnon (1973, 1991), and Arestis and Demetriades (1999) suggest that financial liberalization should only be implemented after institutional reform and trade liberalization. Incorrect liberalization sequencing could be the reasons for the recent series of financial crises in emerging markets as has been reported by many researchers (Stiglitz, 1999; Indrawati, 2002). It is therefore very important for emerging economies to first establish or strengthen their markets and institutions to benefit from liberalization.

Appendix 1: Literature Review of Stock Returns and Volatility for Emerging Markets

Study	Type of Study	Definition of Liberalization	Date of Liberalization	Model Specification	Results
Grabel (1995) The Journal of Development Studies	Construction of a volatility index for individual countries for six emerging countries.	Official financial liberalization programme by the government.	Various official financial liberalization dates ranging from 1973 for Uruguay to 1983-1984 for Indonesia.	Construction of Neo-Classical Volatility Index (NC-VI) and the Keynesian Volatility Index (Keynesian-VI).	Finds that the volatility for most of the countries in the sample is increasing
Aitken (1996) IMF Working Papers No 96/34.	A study of the autocorrelation of total return for the Overall Emerging Market Index from the IFC's Emerging Market Database to determine whether bubble-like price behaviour is evident in emerging stock markets.	Volume of capital flow to emerging markets.	1988 to 1991 as pre-liberalization and 1992 to 1995 as post-liberalization.	A simple variance ratio test on total returns data.	This paper finds that volatility increases profoundly after liberalization compared to before liberalization.
Levine and Zervos (1998) World and Development	Individual country for 16 EMs. 1976-1993 monthly data.	Liberalize capital flow and free repatriation of dividends.	Perron (1989).	Schwert (1989) event study and Perron (1989) test for structural breaks and volatility changes.	Volatility tends to be higher after liberalization.

Note: EMs refers to emerging markets

Appendix 1: (Continued) Literature Review of Stock Returns and Volatility for Emerging Markets

Study	Type of Study	Definition of Liberalization	Date of Liberalization	Model Specification	Results
Bekaert and Harvey (1997) Journal of Financial Economics	Cross-sectional and time-series data for 17 EMs, 1976-1992 monthly data.	Capital market reform.	From Bekaert (1995).	OLS and GLS on liberalization dummy.	Liberalization significantly reduced volatility.
Bekaert and Harvey (2000) Journal of Finance	Pooled cross-section and time-series study of 20 EMs, 1976-1995 monthly data.	Regulatory changes, ADR and country funds and structural breaks.	First date of liberalization ~official, ADR, country fund.	Pooled cross-sectional and time-series regression on 4 liberalization dummies. Controlling for financial and macroeconomic development. They used Semi-Parametric Autoregressive Conditional Heteroskedasticity SPARCH	Insignificant increase in volatility, increase in correlation and beta with the world market for post-liberalization.
Kim and Singal (2000) Journal of Business	14 EMs around liberalization.	Removal of restrictions on capital controls.	Official date.	ARCH/GARCH event study by comparing before and after liberalization.	Fourth and fifth year after liberalization's volatilities are significantly less than before liberalization.

Note: EMs refers to emerging markets

Appendix 1: (Continued) Literature Review of Stock Returns and Volatility for Emerging Markets

Study	Type of Study	Definition of Liberalization	Date of Liberalization	Model Specification	Results
Law (2005) unpublished thesis	Individual country of 5 Asia-Pacific emerging markets: Thailand, Malaysia, Korea, Taiwan and Philippines.	Regulatory changes, ADR and country funds and structural breaks.	Based on Bekaert and Harvey (2000).	GARCH/EGARCH model by considering possible structural break in the model and looking for volatility effects after liberalization.	Decreasing volatility after liberalization for Korea, Taiwan, Philippines and Malaysia, while increasing volatility for Thailand.
Kaminsky and Schmukler (2002) World Bank Working Paper	Construct new own liberalization index that includes 28 developed and EMs and then examined the behaviour of the booms and busts in stock prices over the financial cycle.	Evolution of regulation of the acquisition of shares in the domestic markets, repatriation of capital, interest and dividends.	Using Liberalization Intensity Index developed by Kaminsky and Schmukler (2002).	Short-run and long-run effects of financial liberalization.	Volatility is higher in the short run and more stable in the long run.
Caner and Onder (2005), Applied Economics	Volatility of short-term returns of composite index from 17 EMs is examined to find the factors that influencing the variation of returns.	Not directly defined as liberalization but uses crisis-hit countries due to their high capital flows.	Using post-1990's data for most of the countries.	Employing the VAR model, including the variance decompositions method, to compare source of volatility in EMs and developed markets.	Factors that affect volatility including dividend yield 43%, lagged returns 46%, real interest rate 3.7%, exchange rate 3.2% and world market index 1.6%.

Note: EMs refers to emerging markets

Appendix 1: (Continued) Literature Review of Stock Returns and Volatility for Emerging Markets

Study	Type of Study	Definition of Liberalization	Date of Liberalization	Model Specification	Results
Jayasuria (2005), Emerging Markets Review	Individual of 18 EMs as in Bekaert and Harvey's study of the effect of market liberalization (2000). Removal of restrictions on capital flow.	Regulatory changes, ADR and country funds available to foreign markets.	Official date of liberalization from Bekaert and Harvey (1997).	Variant of GARCH models and then classified volatility into three groups: A- decrease, B-increase and C- no change. Market characteristics of each country that fall into these categories are scrutinized further.	Group A are those with high quality institutions. Group B are those with low quality institutions.
Bekaert and Harvey (2003), Journal of Empirical Finance	Bekaert and Harvey (2000).	Regulatory changes, ADR and country funds and structural breaks.	Official liberalization, ADR, country funds.	Coefficient variation over mean for examining the long-run EMs' volatility patterns and then comparing with developed markets.	Higher volatility pattern after liberalization, but more stable in the long run.
Kassimatis (2002), Applied Financial Economics	The impact of liberalization on 6 EMs' stock exchanges: Argentina, India, Pakistan, Philippines, Korea and Taiwan.	Abolition of limit on foreign capital ownership or as foreign ownership increased in local companies.	Official date of implementation of new liberalization policy.	EGARCH model and compare the news impact curves for pre- and post-liberalization periods.	Decreasing volatility for Argentina, India, Pakistan, Korea and Taiwan. Increasing volatility for Philippines.

Note: EMs refers to emerging markets

Appendix 2: Data Characteristics for All Emerging Markets

Country	Abbreviation	Stock Market Indices	Interest Rate	Official Liberalization Date (T)[a]	Sample Period	
					Model A[b]	Model B[c]
Argentina	ARG	General Index	Money Market	Nov/89	Jan/84-Dec/03	Dec/84-Oct/94
Austria	AUS	ATX	Deposit	#	Feb/86-Dec/03	-
Bangladesh	BAN	Bangladesh All Share Price	Deposit	Jun/91	Jan/90-Dec/03	Jul/86-May/96
Belgium	BEL	BEL 20	T Bill 3M	#	Jan/90-Dec/03	-
Brazil	BRA	IBOVESPA	Money Market	May/91	Jan/84-Dec/03	Jun/86-Apr/96
Chile	CHI	IGPA	Lending	Jan/92	Jan/84-Dec/03	Feb/87-Dec/96
Columbia	COL	IGP	Lending	Feb/91	Feb/85-Dec/03	Feb/86-Jan/96
Hungary	HUN	Budapest BUX	Deposit	#	Jan/91-Dec/03	-
India	IND	Bombay Sensitivity Index	Lending	Nov/92	Jan/84-Dec/03	Dec/87-Oct/97
Indonesia	INA	Jakarta SE Composite	Money Market	Sep/89	Feb/86-Dec/03	Oct/84-Aug/94
Israel	ISR	TA100	Lending	Nov/93	Jan/84-Dec/03	Dec/88-Oct/98
Jordan	JOR	JORDAN Index	Discount Rate	Dec/95	Jan/84-Dec/03	Jan/91-Nov/00
Kenya	KEN	Nairobi SE 20	T Bill 3M	Jan/95	Jan/90-Dec/03	Feb/90-Dec/99
Korea	KOR	Seoul Composite	Lending	Jan/92	Jan/84-Dec/03	Feb/87-Jan/96
Malaysia	MAL	Kuala Lumpur Composite	T Bill 3M	Dec/88	Jan/84-Dec/03	Jan/84-Nov/93
Mexico	MEX	IPC	T Bill 3M	May/89	Jan/84-Dec/03	Jun/84-Apr/94
Morocco	MOR	MASI	Lending	Jun/96	Feb/88-Dec/03	Jul/91-May/01
Nigeria	NIG	All-Share Index 100	Lending	Aug/95	Jan/85-Dec/03	Sep/90-Jul/00

Note: [a] Official liberalization date is based on Bekaert *et al* (2005). The detail discussion can be accessed from http://www.duke.edu/%7Echarvey/Country_risk/chronology/chronology_index.htm. [b] Liberalization intensity index is based on the Chinn and Ito (2002) KAOPEN index. [c] Dummy sample period is based on T - 60 and T + 60, where T is official liberalization date. # Countries that had been considered open before 1984.

Appendix 2: (Continued) Data Characteristics for All Emerging Markets

Country	Abbreviation	Stock Market Indices	Interest Rate	Official Liberalization Date (T)[a]	Sample Period	
Pakistan	PAK	S&P/IFCG Pakistan	Discount Rate	Feb/91	Jan/85-Dec/03	Feb/86-Jan/96
Peru	PER	IGBL	Discount Rate	Jun/92	Feb/91-Dec/03	Jul/87-May/97
Philippines	PHI	Manila Composite	Lending	Jun/91	Jan/85-Dec/03	Jul/86-May/96
Poland	POL	Warsaw General Index	Lending	#	Jan/91-Dec/03	-
S. Africa	SAF	South Afri-DS Market	Lending	Jan/96	Jan/84-Dec/03	Feb/91-Dec/00
Singapore	SIN	Singapore Straits Industrial)	T Bill 3M	#	Jan/85-Dec/03	-
Spain	SPA	Madrid General	T Bill 3M	Jun/85	Jan/84-Dec/03	Jan/84-May/90
Sri Lanka	SRI	Colombo All Share	T Bill 3M	Oct/90	Jan/85-Dec/03	Nov/85-Sep/95
Thailand	THA	Bangkok S.E.T. Index	Lending	Sep/87	Jan/84-Dec/03	Jan/84-Aug/92
Turkey	TUR	S&P/IFCG Turkey	Money Market	Aug/89	Feb/87-Dec/03	Sep/84-Jul/94
Venezuela	VEN	S&P/IFCG Venezuela	Lending	Jan/90	Feb/85-Dec/03	Feb/85-Dec/94
Zimbabwe	ZIM	Zimbabwe Industrial Index	Lending	Jun/93	Jan/84-Dec/03	Jul/88-May/98

Note: [a] Official liberalization date is based on Bekaert *et al* (2005). The detail discussion can be accessed from http://www.duke.edu/%7Echarvey/Country_risk/chronology/chronology_index.htm. [b] Liberalization intensity index is based on the Chinn and Ito (2002) KAOPEN index. [c] Dummy sample period is based on T - 60 and T + 60, where T is official liberalization date. # Countries that had been considered open before 1984.

Appendix 3: Conditional Mean Estimates for All Markets

	ARG[1]	AUS	BAN	BEL	BRA[1]	CHI	COL	HUN	IND	INA[1]
Constant	0.0544	-0.2154	-0.0133	0.0077	-0.0851	0.0536	-0.0038	-0.1844	0.0548	-0.1112
	(0.0910)	(0.1143)	(0.0099)	(0.0137)	(0.0345)	(0.0832)	(0.0727)	(0.1852)	(0.0338)	(0.0636)
Inflation	-0.0865	1.8615	-0.2736	-3.3078	0.1298	-0.0028	0.0307	0.5955	-0.4846	-0.0243
	(0.1457)	(1.3154)	(0.2264)	(1.5610)	(0.0722)	(0.0020)	(0.5801)	(0.6366)	(0.5909)	(0.0439)
Interest rate	-0.0556	0.0160	-0.0428	-0.1009	-0.0026	-0.0015	-0.4933	0.2002	-0.2166	0.9683
	(0.0186)	(0.1198)	(0.0296)	(0.0528)	(0.0202)	(0.0005)	(0.0940)	(0.0836)	(0.2419)	(0.6079)
Real exchange rate	0.0751	0.0361	-0.2120	-0.0630	-1.1874	0.0000	-0.7309	0.7973	-0.3527	-0.1132
	(0.1028)	(0.1613)	(0.1905)	(0.1404)	(0.1210)	(0.0001)	(0.1991)	(0.3056)	(0.1664)	(0.3082)
Political stability	-0.0009	0.0025	-0.0002	0.0004	0.0018	0.0005	0.0007	0.0021	-0.0010	0.0016
	(0.0012)	(0.0014)	(0.0002)	(0.0003)	(0.0006)	(0.0010)	(0.0012)	(0.0029)	(0.0007)	(0.0012)
S&P500	0.7560	0.1739	-0.0237	0.2305	-0.4255	0.5765	0.1543	-0.2597	-0.0403	0.3473
	(0.2773)	(0.1370)	(0.0286)	(0.1623)	(0.2275)	(0.1341)	(0.1128)	(0.2493)	(0.1645)	(0.1408)
FT100	0.1676	0.0335	0.0096	0.0345	0.6337	0.0963	-0.0310	-0.0021	0.2482	-0.0678
	(0.2343)	(0.1173)	(0.0230)	(0.1337)	(0.1958)	(0.1180)	(0.1057)	(0.2474)	(0.1492)	(0.1663)
Nikie225	0.1239	0.0674	-0.0033	-0.0795	0.1261	-0.0333	-0.1409	0.0072	-0.0121	0.0178
	(0.1151)	(0.0651)	(0.0108)	(0.0475)	(0.1028)	(0.0656)	(0.0654)	(0.1236)	(0.0934)	(0.0705)
Time trend	0.0001	0.0000	0.0001	-0.0001	-0.0001	0.0010	-0.0002	0.0001	0.0000	0.0001
	(0.0001)	(0.0001)	(0.0000)	(0.0000)	(0.0000)	(0.0009)	(0.0001)	(0.0003)	(0.0001)	(0.0001)
Autoregressive term	-0.0577	-0.8339	0.4527	0.8277	-0.9033		0.0910			0.0877
	(0.4414)	(0.1250)	(0.1873)	(0.0710)	(0.0182)		(0.2361)			(0.2290)
Moving average term	-0.0506	0.7962	-0.4740	-0.9768	0.9643		0.1867			0.2862
	(0.4491)	(0.1448)	(0.1918)	(0.0438)	(0.0079)		(0.2324)			(0.2114)

Note: [1] Zero-one dummy variable has been added in the estimation of mean equation. Number in parentheses is Bollerslev-Wooldrige robust standard errors. * significant at 5% or 1% level.

Appendix 3: (Continue) Conditional Mean Estimates for All Markets

	ISR	JOR	KEN	KOR[1]	MAL[1]	MEX[1]	MOR	NIG	PAK	PER
Constant	0.0030	-0.0298	-0.0468	-0.0188	0.0706	0.0220	0.3523	-0.0864	0.0129	-0.0078
	(0.0178)	(0.0139)	(0.0238)	(0.0321)	(0.0452)	(0.1482)	(0.0980)	(0.0997)	(0.0195)	(0.0198)
Inflation	0.1671	0.0892	0.2601	-2.9034	2.3542	-0.1257	-0.0026	-0.0136	-0.8598	-0.0229
	(0.0987)	(0.2265)	(0.2079)	(1.4282)	(1.7205)	(0.3401)	(0.0012)	(0.2719)	(0.3530)	(0.0136)
Interest rate	-0.0285	-0.0233	-0.0890	-0.2121	-0.0389	-0.3365	-0.0056	-0.3367	-0.3794	0.0256
	(0.0273)	(0.0784)	(0.0191)	(0.2546)	(0.1008)	(0.0334)	(0.0032)	(0.2623)	(0.2087)	(0.0055)
Real exchange rate	-0.0239	0.1747	-0.1907	-0.1984	-0.6180	0.0312	-0.0158	-0.0553	-0.8009	-0.1059
	(0.1534)	(0.1474)	(0.0807)	(0.5031)	(0.4268)	(0.0927)	(0.0042)	(0.0407)	(0.1816)	(0.0470)
Political stability	0.0002	0.0005	0.0012	0.0006	-0.0010	0.0000	-0.0003	0.0016	-0.0001	0.0002
	(0.0004)	(0.0003)	(0.0004)	(0.0004)	(0.0007)	(0.0020)	(0.0006)	(0.0020)	(0.0005)	(0.0006)
S&P500	0.0709	-0.1393	-0.1574	0.6379	0.1679	0.6574	-0.2237	-0.5067	0.0393	0.2514
	(0.1087)	(0.0885)	(0.0727)	(0.1709)	(0.1325)	(0.1697)	(0.0968)	(0.2566)	(0.0534)	(0.0958)
FT100	0.1049	0.1501	0.2892	-0.0402	0.1203	0.0739	0.2417	0.2293	0.0003	-0.0900
	(0.1020)	(0.0647)	(0.0820)	(0.1442)	(0.1166)	(0.1480)	(0.0675)	(0.2033)	(0.0481)	(0.0804)
Nikie225	-0.0579	0.0406	-0.0550	0.3395	0.0990	0.0660	-0.0179	0.1860	0.0803	-0.0520
	(0.0573)	(0.0361)	(0.0382)	(0.0812)	(0.0650)	(0.0805)	(0.0439)	(0.1008)	(0.0379)	(0.0236)
Time trend	0.0000	0.0000	-0.0002	0.0000	0.0000	-0.0001	0.0008	0.0001	0.0000	0.0000
	(0.0001)	(0.0001)	(0.0000)	(0.0001)	(0.0001)	(0.0001)	(0.0003)	(0.0001)	(0.0000)	(0.0001)
Autoregressive term	-0.0740	-0.8380		0.9536	-0.2565	0.0537	0.8402		-0.5893	-0.1154
	(0.0426)	(0.0281)		(0.0241)	(0.0801)	(0.2373)	(0.0463)		(0.0758)	(0.0562)
Moving average term	0.0974	0.9276		-0.9975	0.4130	-0.2018	-0.9190		0.6852	
	(0.0429)	(0.0178)		(0.0000)	(0.0972)	(0.2222)	(0.0188)		(0.0550)	

Note: [1] Zero-one dummy variable has been added in the estimation of mean equation. Number in parentheses is Bollerslev-Wooldrige robust standard errors.

Appendix 3: (Continue) Conditional Mean Estimates for All Markets

	PHI[1]	POL	SAF	SIN	SPA	SRI	THA[1]	TUR	VEN	ZIM
Constant	0.0499	0.3504	0.0612	0.3129	-0.1061	0.0323	1.1130	0.1128	-0.1876	0.0259
	(0.0396)	(0.3685)	(0.0424)	(0.1074)	(0.0652)	(0.0198)	(0.1212)	(0.1069)	(0.0001)	(0.0837)
Inflation	-0.5113	0.3936	-0.6523	0.3323	0.4103	0.1173	1.1130	-1.2499	-0.4344	-0.5912
	(0.7635)	(0.2644)	(1.0410)	(1.2576)	(1.0024)	(0.3546)	(0.1212)	(0.4042)	(0.2619)	(0.4942)
Interest rate	-0.0493	-0.0394	-0.0142	0.0058	-0.0556	-0.0152	1.1130	0.0131	-0.1822	-0.2735
	(0.0709)	(0.0262)	(0.0855)	(0.0128)	(0.0943)	(0.0303)	(0.1212)	(0.0406)	(0.0817)	(0.1000)
R.exchange	-1.3491	0.8646	-0.5020	-0.8808	-0.0773	-0.0167	1.1130	-0.2359	-0.0851	-0.0034
	(0.2596)	(0.2927)	(0.1798)	(0.3617)	(0.1502)	(0.2122)	(0.1212)	(0.2615)	(0.0965)	(0.0017)
P. Stability	-0.0004	-0.0026	-0.0008	-0.0044	0.0016	-0.0008	1.1130	-0.0006	0.0026	0.0013
	(0.0010)	(0.0030)	(0.0006)	(0.0016)	(0.0010)	(0.0005)	(0.1212)	(0.0019)	(0.0000)	(0.0018)
S&P500	0.0952	-0.5195	0.0381	0.3509	0.3984	-0.0388	1.1130	0.6486	-0.1706	-0.0416
	(0.1822)	(0.3411)	(0.1851)	(0.1653)	(0.1344)	(0.0834)	(0.1212)	(0.4143)	(0.2695)	(0.1743)
FT100	0.1526	0.3759	0.0910	-0.2074	-0.0174	0.0395	1.1130	0.1740	0.3335	-0.1222
	(0.1747)	(0.2527)	(0.1345)	(0.1305)	(0.1107)	(0.0905)	(0.1212)	(0.2982)	(0.2165)	(0.1482)
Nikie225	0.0978	0.1277	0.0353	0.0098	-0.0212	0.0518	1.1130	-0.0383	-0.0214	0.0496
	(0.0923)	(0.1618)	(0.0664)	(0.0618)	(0.0700)	(0.1172)	(0.1212)	(0.1657)	(0.0629)	(0.0841)
Time trend	-0.0002	-0.0006	0.0000	0.0004	-0.0001	0.0000	1.1130	-0.0002	0.0002	0.0004
	(0.0002)	(0.0006)	(0.0001)	(0.0002)	(0.0001)	(0.0001)	(0.1212)	(0.0002)	(0.0000)	(0.0002)
AR term	-0.8742	0.8698	0.5534	0.8572	0.2750	-0.6428	1.1130		-0.9977	0.8265
	(0.0345)	(0.0384)	(0.3071)	(0.0478)	(0.2142)	(0.0683)	(0.1212)		(0.0062)	(0.1187)
MA term	0.9407	-0.9724	-0.6481	-0.9321	-0.4936	0.8053	1.1130		1.0464	-0.6572
	(0.0261)	(0.0104)	(0.2819)	(0.0236)	(0.2032)	(0.0539)	(0.1212)		(0.0258)	(0.1576)

Note: [1] Zero-one dummy variable has been added in the estimation of mean equation. Number in parentheses is Bollerslev-Wooldrige robust standard errors.

CHAPTER THREE

UNCOVERED INTEREST PARITY IN EMERGING ECONOMIES: DOES THE PREDICTION TIME HORIZON MATTER?

3.1 Introduction

Uncovered interest rate parity (UIP) is one of the oldest macroeconomic propositions and is still a building block of many economic models. Contrary to its widespread theoretical use, empirical tests of UIP reject the predicted relation between interest rate differential and exchange rate changes. The exchange rates of countries with high nominal interest rates tend to appreciate rather than depreciate as hypothesized by the UIP. Empirically, the average regression coefficient of nominal exchange rate changes on interest differential regressions is -0.88 (Froot and Thaler, 1990); -0.3 (McCallum 1994); -0.4 (Engel 1996); and -0.8 (Chinn and Meredith, 2004). Excellent reviews of the long-outstanding puzzle are provided by Engel (1996) and Chinn (2006). Some of the explanations offered for the rejection include: expectational errors (Mark and Wu, 1998; Kirikos, 2002), the presence of time-varying risk premia (Francis *et al.* 2002; Sarantis, 2006), or policy behaviour (McCallum, 1994; Christensen 2000; Chinn and Meredith, 2004).

Recently, some studies have attempted to find new ground for UIP by testing its validity at longer horizons[12]. Fujii and Chinn (2001) and Chinn and Meredith (2004) have shown that the status of the UIP could crucially depend on the long-term variables. Chinn and Meredith (2004) show that using longer maturity financial instruments (five- to ten-year bonds) substantially changes the sign of the interest rate coefficient from negative to positive, with three (out of six) currencies not

[12] A contrary view to the longer horizons estimation power can be found in the works by Chaboud and Wright (2005), and Yang and Shintani (2006). They believe that UIP only holds for a very short time span. However, these studies do not consider emerging markets.

rejecting the hypothesis that the slope coefficient is equal to 1. Augmenting McCallum's (1994) model, they argue that at short horizons, shocks in the exchange markets lead to monetary policy responses that result in negative correlation between exchange rate changes and interest rate differentials. Contrary to short horizons, at longer time horizons interest rates and exchange rates are both driven by macroeconomic "fundamentals" factors that results in a more consistent relationship with the UIP.

However, Valkanov (2003) argues that using long-horizon regression could provide misleading statistical inferences compared to the short-horizon regression. Extra caution is required in long-horizon regression because of the overlapping sums of the original series (close to a unit root process) that might lead to *t*-statistics that do not converge to a well-defined asymptotic distribution. This may result in inconsistent OLS estimators and inadequate measures for coefficient of determination, R^2. Similar arguments can be found in Kilian (1999). He employs bootstrap methods on monetary models to show that there is no significant increase in predictive power by using longer-horizon estimation methods.

The arguments used by Valkanov (2003) are no different to those in Granger and Newbold (1974) and Phillips (1986). The analogy among them lies in finding a spurious correlation between persistent variables when they are statistically independent. All these facts are related to the non-stationary behaviour that is usually exhibited by long-horizon variables.

All of the above-mentioned studies concentrate on developed and industrialized economies. Given the current status of liberalization in emerging markets (Levy-Yeyati and Sturzenegger 2005, and Chinn and Ito, 2005) and their growing importance in global financial markets (Bekaert and Harvey, 2003; Stiglitz,

2004), in this paper we re-examine UIP for emerging economies focussing on different time horizons to evaluate whether UIP holds or not. Further we use different based-currency for relative-country choice sensitivity as a means of robustness checking.

Our main contributions to the literature are as follows. First, until recently only very few studies dealt with UIP in emerging countries, among others are Bansal and Dalhquist (2000), Flood and Rose (2001), Francis *et al.* (2002), and Frankel and Poonawala (2004). This is because emerging markets were relatively closed until the mid 1980s. Previously, excessive constraints were imposed by local authorities either on capital movements or exchange rate changes, which makes the testing of UIP uninteresting. In this sense, we complement the existing literature on UIP, since empirical work on emerging markets is still lacking.

Second, the majority of studies considering emerging countries use short-term forecast horizons (k) in the regression of UIP models: Bansal and Dahlquist (2000) use one- and three-month intervals, while Flood and Rose (2001), Francis *et al.* (2002) and Frankel and Poonawala (2004) use one-month horizons. Contrary to these papers, we extend the test of UIP by focussing on the different exchange and interest rate maturities from short- to medium-term, i.e. one-, three- and twelve-month horizons ($k=1$, $k=3$ and $k=12$) using both dynamic time series and panel regression. Our findings confirm the earlier results for emerging economies, although, at longer horizons the slope coefficients are getting closer to unity for most of the markets. As a robustness check, we further test the UIP hypothesis using different combinations of base countries using the same prediction horizon. The results are similar. In practice, there is no sound basis for choosing other than the US

as a base country since 89% of exchange rate trading in the world uses the US dollar[13].

The remainder of the paper is organized as follows: In section 3.2, we briefly discuss the theory and recent evidence of the UIP in emerging markets. Section 3.3 describes the dataset used in the empirical analysis and the layout of the econometric procedures. Section 3.4 discusses the estimation results. Finally, section 3.5 offers some concluding remarks.

3.2 Theory and Evidence in Emerging Markets

The UIP states that the interest differential between two countries should equal the expected exchange rate changes. If the nominal interest rate in the foreign market is higher compared to the local market, it allows investors to borrow at the relatively low local rate and invest the proceeds at the foreign higher rate. Then, at the end of the k-th period, the foreign currency proceeds are converted back to local currency. The local currency is expected to appreciate just to reach an equilibrium point and cancel out the excess profit between these two markets. Ideally, this proposition holds true if the market satisfies the condition of no economic and/or political barriers (i.e. risk premium and political risk), between countries. In addition, the agents are assumed to be risk-neutral and behave rationally. Then, active arbitrage trading ensures that the UIP hypothesis holds.

The above explanation is one specification of UIP, which can be expressed, in the following equation:[14]

$$\Delta s_{t,k} = \alpha + \beta(i_{t,k} - i^{*}_{t,k}) + \varepsilon_{t,k} \qquad \text{Equation 3.1}$$

[13] Based on the report publish by the Bank for International Settlement, Basel, 2003.
[14] Model has been constructed base on Equation 2 of Bansal and Dahlquist (2000).

Where, $\Delta s_{t,k}$ is the change of the domestic exchange rate over time period k, $(i_{t,k} - i^{*}_{t,k})$ is the interest rate differential between domestic and foreign markets for maturity in k periods, subscript t represents time, and $\varepsilon_{t,k}$ is an error term. Given that markets are efficient with regard to arbitrage activities and neither political nor economic barriers exist between markets, the estimated parameters of α and β should not be statistically different from 0 and 1, respectively, and the error term should be white noise. The failure of any hypothesis from which the model is derived indicates the presence of a time-invariant risk premium, a time-varying risk premium or political risk.

Testing of the UIP in emerging markets is still relatively lacking. This can be for at least two reasons. The first is the relatively fixed exchange rate regimes and extensive controls on the economy in some of these markets until the mid '80's and early '90's. These restrictions violate the theoretical framework of UIP and may cause the "peso problem" in its empirical testing (Krasker, 1980). In this study, we try to avoid this problem by dropping countries with excessive capital control and adopting hard peg exchange rate regimes. We thus consider only countries that have a free capital account and a relatively floating exchange rate regime which allows the exchange rate to fluctuate i.e. from a band to a free-floating regime. Recent literature has found the difficulty in establishing whether a declared flexible or fixed exchange rate regime is in fact just *de jure* or also *de facto* (Reinhart and Rogoff, 2004).

Second, the empirical failure of the UIP suggest either a time varying risk premium that is negatively correlated with the interest differential or a departure from rationality. Considering this problem, many studies use to assume rational expectation hypothesis to hold (Francis *et al.* 2002; and Cheung *et al.* 2005; and

2006) [15] or to add shock to the UIP equation (see e.g., McCallum 1994). This shock is often referred to as a "risk premium" shock for a variety of reasons, including transaction costs (see *inter alia*, Baldwin, 1990; Dumas, 1992; Hollifield and Uppal, 1995), central bank intervention (e.g. Moh, 2006), and the existence of limits to speculation (e.g. Lyons, 2001). Even though introducing "risk premium" shocks improve the fit of the UIP equation, these shocks could induce counterfactual correlation between interest rate and aggregates quantities. Therefore, allowing for "risk premium" shocks amounts to introducing an important source of model misspecification that is likely affect policy analyses. In addition, due to lack of expectation (*ex-ante*) exchange rate datasets, many researchers like Francis *et al.* (2002) and Cheung *et al.* (2005 and 2006), carried out investigation of UIP in emerging markets by assuming rational expectations and using an *ex-post* instead of an *ex ante* series even though, strictly speaking, the UIP is an *ex ante* concept defined by expectations rather than *ex post* realized depreciation rates.[16]

Recent work on UIP in emerging economies can be found in Bansal and Dahlquist (2000). They assume a latent factor model for both cross-sectional and time series data from 12 emerging economies. They show that UIP performs better in emerging economies compared to developed economies. Their findings indicate that the deviation from UIP occurs only in two specific scenarios: the first is when the US interest rate exceeds the foreign interest rate, second, if the foreign interest is higher than the local rate[17]. They also find that country-specific attributes such as *per capita*

[15] Assuming rational expectation hypothesis to hold, Cheung et al. (2005) assume that the *ex post* realizations are unbiased predictor of the *ex ante* exchange rate. They term this un-standard UIP concept as Rational Uncovered Interest Parity (RUIP).

[16] Explanation of the limited dataset problem in developing economies can be found in Chinn and Frankel, (1993).

[17] Bansal and Dahlquist (2000) define the local rate as the UK interest rate and the foreign interest rate is for the other developed or developing countries under study. All currencies are in US dollars.

income, inflation, volatility, country risk rating and nominal interest rate are important in explaining the deviation from the UIP hypothesis.

Motivated by liberalization policies in most of the emerging countries, Francis *et al.* (2002) further investigated the empirical puzzle of the UIP for 9 developing countries (Chile, Columbia, Mexico, India, Korea, Malaysia, Pakistan, Thailand and Turkey) in pre- and post-liberalization eras using a multi-factor conditional asset-pricing model estimated in a multivariate GARCH framework. This research confirms that the deviation from the UIP prevails in most of the emerging countries and the phenomenon is country-specific in nature.

Using the one-month forward exchange rate, Frankel and Poonawala (2004) tested the unbiasedness hypothesis for fourteen emerging countries from 1996 to 2004. The results from the individual market time-series regressions are mixed. Eight markets experienced a positive estimated forward discount coefficient, β (although smaller than unity), and the remaining were negative and statistically insignificant. They also found a positive slope for β by pooling together the emerging countries.

Summarizing, the evidence against the UIP puzzle in the post-liberalization era in emerging economies is not as severe as was commonly thought in the pre-liberalization period. However, the evidence is still far from conclusive and it is country-specific in nature.

3.3 Data and Econometric Specification

3.3.1. Data Description

In this study, the UIP hypothesis is tested using monthly data of exchange rate changes and interest rate differentials spanning from January 1995 to December 2005 period for 15 emerging markets with the US as a base country(hereafter we call

this a 'US-base model'). The countries included are four Latin American emerging markets (Brazil, Chile, Mexico and Venezuela), four Asian emerging markets (Indonesia, Korea, Philippines and Thailand), five European emerging markets (Hungary, Poland, Portugal, Romania and Russia), one Middle-East (Israel) and an African emerging market (Morocco). The basic criterion for selecting these countries is based on the capital account openness, an exchange rate regime that at least allows for large movement bands for managed regimes (base on the visual inspection of exchange rate series in Figure 3.1 to 3.3) and data availability. The determination of the exchange rate regime in each market is based on work by Reinhart and Rogoff (2004) Appendix III.[18] The interest rates used are the 1-month, 3-month and 12-month deposit rate, inter-bank rate or Treasury Bill rate (with priority to the deposit rate, if available) of monthly frequency. All interest rate series are downloaded from Datastream. The monthly exchange rate series are extracted from the International Financial Statistic (IFS) and expressed in terms of US dollars per unit of emerging market currency. Details of the data set used in the analysis are presented in Table 3.1.

To check the robustness of our results, we add Japan and Germany as relative countries (hereafter the estimation that uses Japan as the relative country is called the 'Japan-base model' and likewise for Germany, the 'Germany-base model'). The selection of these two countries is based on the total volume of imports and exports in the direction of trade (DoTs) which shows that these two markets are placed either first or second after the US in the sample countries. All Asian countries have Japan as their major trading partner after the US. On the other hand, all the Middle-East,

[18] Countries with hard peg exchange rate regimes to the US dollar, like Malaysia (1998 to 2005) and Argentina (1991 to 2001), or capital control regimes, like India, are omitted from the dataset.

African and European emerging markets have Germany as their major trading partner. Mixed combinations between Japan and Germany are found for Latin American emerging markets. Altogether the US, Japan and Germany are accounted for approximately 40% of emerging countries' total imports and exports.

Table 3.1: Data Specification for Emerging Countries from 1995:01 to 2005:12

	Interest Rate	Time Horizon	Period	Base Country
Latin America				
Brazil	Deposit	1-Month	95:01-05:12	US, Germany
Chile	Deposit	1-, 3-12-Month	95:01-05:12	US, Japan
Mexico	Deposit	1-, 3- , 12-Month	96:01-05:12	US, Japan
Venezuela	Deposit	1-, 3-Month	96:12-05:12	US, Germany
Asia				
Indonesia	Deposit	1-, 3- , 12-Month	95:01-05:12	US, Japan
Korea	Deposit	3- , 12-Month	95:01-05:12	US, Japan
Philippines	Deposit	1-, 3- , 12-Month	95:01-05:12	US, Japan
Thailand	Deposit	1-, 3- , 12-Month	95:01-05:12	US, Japan
Middle-East				
Israel	T-Bill	3- , 12-Month	95:01-05:12	US, Germany
Morocco	Deposit	1-, 3- , 12-Month	95:01-05:12	US, Germany
Europe				
Hungary	Interbank	1-, 3- , 12-Month	95:01-05:12	US, Germany
Poland	Interbank	1-, 3- , 12-Month	95:01-05:12	US, Germany
Portugal	Interbank	1-, 3-Month	95:01-05:12	US, Germany
Romania	Interbank	1-, 3- , 12-Month	95:01-05:12	US, Germany
Russia	Interbank	1-, 3-Month	95:01-05:12	US, Germany

Data for nominal interest rates are collected from Datastream. The selection of relative country is base on the first two largest trading partners with respective emerging economies in direction of trade (DoT) statistics.

3.3.2 Econometric Methodology

We follow the research procedure use in Bansal and Dalhquist (2000) and Frankel and Poonawala (2004) by considering both time series and panel data estimation to investigate the status of UIP in emerging markets. The empirical analysis of Equation 3.1 is carried out by developing the following basic steps for the

three different models, i.e. US-base model, Japan-base model and Germany-base model. The name of the model is chosen depending on the relative country used in the exchange rate arrangements.

3.3.2.1 Time Series Analysis

For preliminary analysis, we implement unit root tests using the Augmented Dickey Fuller (*ADF*) test in level and first difference of the series covering various time-lag terms. The optimal lag is chosen using the *AIC* specification. The results of the test applied to the series in level indicate that we do not reject the null hypothesis of a unit root for interest rate differential at all horizons and for all markets, except for the 1- and 3-month maturity of Romania and Russia. In the case of exchange rate changes, we only fail to reject the null of unit roots at the 12-month horizon for all countries. The first difference series are stationary. In general, the results show that all interest rate series are *I(1)*, while exchange rates are *I(0)* for 1- and 3-month horizons and *I(1)* for the 12-month horizon. Table 3.2 provides a summary of the ADF unit root tests.[19]

Due to the stationarity property of 1- and 3-month horizons for dependent variables (exchange rate changes), which at $k=12$ becomes non-stationary, we split the estimations of UIP into two procedures. First we use the standard ordinary least squares (OLS) method for $k = 1$ and $k = 3$ with additional dummy variables to capture the crises that affected some of the countries during the sample period: Asian financial crisis 1997 and the Russian crisis 1998. The Newey-West robust standard

[19] However results of unit root testing from the ADF statistics should be interpreted with extra caution. It may suffer from difficulties in distinguishing between I(1) and near unit root process due to overlapping date in higher k. The test also can have low power especially in small sample. Elliot et al (1996), Kwiatkowski et al (1992) and Ng and Peron (2001), among others, provide better option of unit root tests.

errors are used to give consistent covariance matrices in the presence of both serial correlation and heteroskedasticity.

Table 3.2: Summary of Unit Root Properties of Exchange Rate Movement and Interest Rate Differential Using Augmented Dickey Fuller (ADF) Unit Root Tests for Data from 1995:01 To 2005:12

Country	Exchange rate			Interest rate		
	1-m	*3-m*	*12-m*	*1-m*	*3-m*	*12-m*
BRAZIL	*I(0)*	-	-	*I(1)*	-	-
CHILE	*I(0)*	*I(0)*	-	*I(1)*	*I(1)*	-
MEXICO	*I(0)*	*I(0)*	*I(1)*	*I(1)*	*I(1)*	*I(1)*
VENEZUELA	*I(0)*	*I(0)*	-	*I(1)*	*I(1)*	-
INDONESIA	*I(0)*	*I(0)*	*I(1)*	*I(1)*	*I(1)*	*I(1)*
KOREA	-	*I(0)*	*I(I)*	-	*I(1)*	*I(1)*
PHILIPPINES	*I(0)*	*I(0)*	*I(1)*	*I(1)*	*I(1)*	*I(1)*
THAILAND	*I(0)*	*I(0)*	*I(1)*	*I(1)*	*I(1)*	*I(1)*
ISRAEL	-	*I(0)*	*I(1)*	-	*I(1)*	*I(1)*
MOROCCO	*I(0)*	*I(0)*	*I(1)*	*I(1)*	*I(1)*	*I(1)*
HUNGARY	*I(0)*	*I(0)*	*I(1)*	*I(1)*	*I(1)*	*I(1)*
POLAND	*I(0)*	*I(0)*	*I(1)*	*I(1)*	*I(1)*	*I(1)*
PORTUGAL	*I(0)*	*I(0)*	-	*I(1)*	*I(1)*	-
ROMANIA	*I(0)*	*I(0)*	*I(1)*	*I(0)*	*I(0)*	*I(1)*
RUSSIA	*I(0)*	*I(0)*	-	*I(0)*	*I(0)*	-

Note: *I(0)* refers to stationary at level form and *I(1)* refers to stationary at first difference. We used 1% and 5% critical value that was provided by MacKinnon (1996) to test the significance level. The lag length has been selected based on AIC to ensure white noise residual. – indicates non availability of series.

Second, due to the persistency problem in dependent and independent variables for $k=12$, estimation and testing of longer-horizon variables cannot be carried out using standard OLS regression methods (Valkanov, 2003). The same argument has been used by Granger and Newbold (1974), Phillips (1986) and Ferson *et al.* (2003) for the possibility of finding spurious regression correlations between persistence variables when they are statistically independent. The OLS estimated parameters will be super-consistent for the true value, but we need to correct for

biases and the distributional divergence of *t*-statistics. In this case, alternatives to the OLS estimation method are required.

We employ Stock and Watson's (1993) Dynamic OLS (*DOLS*) to estimate the long-run parameters of UIP for $k = 12$. The DOLS procedure basically involves regressing any cointegrated *I(1)* variables on other *1(1)* variables, any *I(0)* variables and leads and lags of the first differences of any *I(1)* variables. It can be represented in the following econometric specification;

$$\Delta s_{t,k} = \alpha + \beta_D (i_{t,k} - i^*_{t,k}) + \sum_{q=-q_1}^{q=q_2} \delta_q \Delta (i_{t-q,k} - i^*_{t-q,k}) + \varepsilon_{t,k} \quad \text{Equation 3.2}$$

where Δs, $(i_{t,k} - i^*_{t,k})$ and $\varepsilon_{t,k}$ are the same as in Equation 3.1. β_D, the Stock-Watson DOLS parameter, estimates the long-run parameters with the interest rate differential appearing in level. q is the optimum number of lead and lag terms included in the estimation to provide an efficient estimator of the cointegrating coefficient. We also use heteroskedasticity consistent covariance proposed by Newey and West (1987) to avoid the problem of whether or not the regression errors are heteroskedastic and autocorelated.

3.3.2.2. Panel Data Analysis

Bansal and Dalhquist (2000) pointed out that it is difficult to produce reliable and precise point estimates of UIP, especially from emerging markets. Therefore, to complement individual time series estimates, we further investigated the relationship between exchange rate depreciation and interest rate differential using a panel technique. Comparing these two techniques, panel data estimates are more powerful than those obtained by applying individual time series estimations especially in short-span data sets. Levin *et al.* (2002) have shown that using panel analysis will

eventually increase the power of the test, and Baillie and Bollerslev (1997) further explain that pooling the data could minimize the problem of statistical inferences.

In the panel technique, we pool all countries by stacking the series according to cross-section. The empirical investigation test procedure is carried out using the following steps. First, we investigate the unit root properties for each panel using the methodology proposed by Levin *et al.* (2002, LLC hereafter) and Im *et al* (2003, IPS hereafter). We are testing the null of unit root by comparing the IPS *w*-statistics and LLC t^*-statistics to 95% critical values. These two techniques are robust over the problems of homogeneity and heterogeneity across units on the lagged variable. Second, for $k = 1$ and $k = 3$, where exchange rate depreciation and interest rate differential are stationary; we employ the standard panel OLS techniques to the Equation 3.1 with and without fixed effects.

On the other hand, for $k = 12$ where both series are persistent and non-stationary, we utilise two types of the heterogeneous panel cointegration test developed by Pedroni (1999 and 2004) and Kao (1999). Basically, both (Pedroni and Kao) extend the Engel-Granger two-step residual-based cointegration framework to tests involving panel data. In this study, specifically we consider two type of the heterogeneous panel cointegration test developed by Pedroni (1995, 1999 and 2004) which allows different individual effects across N or cross-sectional interdependency. The first type of tests includes the panel rho (ρ), panel non-parametric (*PP*) and panel parametric (*ADF*) statistics. The panel parametric statistics is similar the single-equation *ADF*-test and panel non-parametric statistics is analogous to the Phillips and Perron (1988) test. The second type of tests proposed by Pedroni (1999 and 2004) is comparable to the group mean panel tests of Im *et al.* (2003). Pedroni argues that both types of test are appropriate for testing the null of

cointegration in bivariate panel models with heterogeneous dynamic, fixed effects and heterogeneous cointegrating slope coefficients. Further, Pedroni claims that this method also will take into account the off-diagonal terms in the residual long-run covariance and the effect of spurious regression in the heterogeneous panel. The calculation of the heterogeneous panel and heterogeneous group mean panel tests of rho (ρ), non-parametric (*PP*) and parametric (*ADF*) statistics are as follow;

Panel *ρ-statistics*

$$Z_t = \left(\hat{S}^{*2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{2} \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{*} \Delta \hat{e}_{it}^{*} \qquad \text{Equation 3.3}$$

Panel parametric *ADF-statistics*

$$Z_\rho = \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{2} \right)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right) \qquad \text{Equation 3.4}$$

Panel non-parametric *PP-statistic*

$$Z_{pp} = \left(\hat{\sigma}^2 \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{2} \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right) \qquad \text{Equation 3.5}$$

Group *ρ-statistics*

$$\tilde{Z}_\rho = \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{e}_{it-1}^{2} \right)^{-1} \sum_{t=1}^{T} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right) \qquad \text{Equation 3.6}$$

Group parametric *ADF-statistics*

$$\tilde{Z}_t = \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{S}_i^{-2} \hat{e}^{*2}_{it-1} \right)^{-1/2} \sum_{t=1}^{T} \hat{e}_{it-1}^{*} \Delta \hat{e}_{it}^{*} \qquad \text{Equation 3.7}$$

Group non-parametric *PP-statistic*

$$\tilde{Z}_{pp} = \sum_{i=1}^{N} \left(\hat{\sigma}^2 \sum_{t=1}^{T} \hat{e}_{it-1}^{2} \right)^{-1/2} \sum_{t=1}^{T} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right) \qquad \text{Equation 3.8}$$

where $\hat{\sigma}^2$ is the pooled long-run variance for the non-parametric model given as $1/N\sum_{i=1}^{N}\hat{L}_{11i}^{-2}\hat{\sigma}_i^2$; $\hat{\lambda}_i = 1/2(\hat{\sigma}_i^2 - \hat{S}_i^2)$ where $\hat{L}_i$ is used to adjust for autocorrelation in panel parametric model, $\hat{\sigma}_i^2$ and $\hat{S}_i^2$ are the log-run and contemporaneous variances for individual i and $\hat{S}_i^2$ are obtained from individual *ADF*-test of $e_{it} = \rho_i e_{it-1} + \upsilon_{it}$. S^{*2} is the individual contemporaneous variance from the parametric model, $\hat{e}_{it}$ the estimated residual from the parametric cointegration in Equation 3.1 while $\hat{e}_{it}^*$ the estimated residual from the parametric model and $\hat{L}_{11i}$ the estimated log-run covariance matrix for $\Delta\hat{e}_{it}$ and L_i is the ith component of the lower-triangular Cholesky decomposition of matrix Ω_i for $\Delta\hat{e}_{it}$ with the appropriate lag length determined by the Newy-West method.

Further, we consider the panel cointegration tests of Kao (1999). The Kao test follows the same basic approach as the Pedroni tests, but specifies cross-section specific intercepts and homogeneous coefficient on the first stage regressors. The limiting distribution of the residual-based cointegration tests using DF test and ADF are as follows;

$$DF_\rho = \frac{T\sqrt{N}(\hat{\rho}-1)+3\sqrt{N}}{\sqrt{10.2}} \qquad \text{Equation 3.9}$$

$$DF_t = \sqrt{1.25}t_\rho + \sqrt{1.875N} \qquad \text{Equation 3.10}$$

$$DF_\rho^* = \frac{\sqrt{N}T(\hat{\rho}-1)+3\sqrt{N}\hat{\sigma}_\upsilon^2/\hat{\sigma}_{0\upsilon}^2}{\sqrt{3+36\hat{\sigma}_\upsilon^4/(5\hat{\sigma}_{0\upsilon}^4)}} \qquad \text{Equation 3.11}$$

$$DF_t^* = \frac{t_\rho + \sqrt{6N}\hat{\sigma}_\upsilon/(2\hat{\sigma}_{0\upsilon})}{\sqrt{\hat{\sigma}_{0\upsilon}^2/(2\hat{\sigma}_\upsilon^2)+3\hat{\sigma}_\upsilon^2/(10\hat{\sigma}_{0\upsilon}^2)}} \qquad \text{Equation 3.12}$$

$$ADF = \frac{t_{ADF} + \sqrt{6N}\hat{\sigma}_{\upsilon} / (2\hat{\sigma}_{0\upsilon})}{\sqrt{\hat{\sigma}_{0\upsilon}^2 / (2\hat{\sigma}_{\upsilon}^2) + 3\hat{\sigma}_{\upsilon}^2 / (10\hat{\sigma}_{0\upsilon}^2)}} \qquad \text{Equation 3.13}$$

where the $\hat{\sigma}_v^2$ and $\hat{\sigma}_{0v}^2$ are the estimated variance and long run variances. The covariance of $w_{it} = [u_{it}\, \varepsilon_{it}]'$ is estimated as $\hat{\Sigma} = \begin{bmatrix} \hat{\sigma}_u^2 & \hat{\sigma}_{u\varepsilon} \\ \hat{\sigma}_{u\varepsilon} & \hat{\sigma}_{\varepsilon}^2 \end{bmatrix} = \frac{1}{NT} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{w}_{it} \hat{w}_{it}'$. On the other hand, the long run covariance is estimated as $\hat{\Omega} = \begin{bmatrix} \hat{\sigma}_{0u}^2 & \hat{\sigma}_{0u\varepsilon} \\ \hat{\sigma}_{0u\varepsilon} & \hat{\sigma}_{0\varepsilon}^2 \end{bmatrix} = \frac{1}{N} \sum_{i=1}^{N} \left[\frac{1}{T} \sum_{t=1}^{T} \hat{w}_{it} \hat{w}_{it}' + \kappa(\hat{w}_i) \right]$ where κ is any kernel function. Under the null of no cointegration, Kao shows that all the DF_{ρ}, DF_t, DF_{ρ}^*, DF_t^*, and ADF tests statistics are converge to a standard normal asymptotic distribution.

However the result from both panel unit root and panel cointegration tests should be treated with extra careful. It is because these two tests depend on the assumption of no cross-country correlation among the errors which not always the case in many studies.[20] Maddala and Wu (1999) have shown that panel unit root test based on the assumption of cross-sectional independence perform poorly for cross-sectionally correlated panels. O'Connell (1998) added that with cross-sectional dependence, the limiting distributions under the *i.i.d.* assumption may not correct and even if the correct distribution is available, the power to reject the unit root can be greatly diminished. Considering this deficiency new panel unit root tests have been proposed in the literature by Chang (2002), Phillips and Sul (2003), Bai and Ng (2004), Smith et al. (2004) and Pesaran (2007), among others. Similarly, panel cointegration tests that consider cross-sectional dependence in the specification always finds good performance in the statistical inference. Empirical works that

[20] Cross correlations dependency can be caused by common components induced by the numeraire countries in based currency in the UIP testing.

consider this issue can be found in Coakley, et al. (2005) and Bannerjee and Carrion-I-Silvester, (2006) among others.[21]

If there is evidence of cointegration, further we estimate the cointegration coefficients for panel using bias-corrected ordinary least squares (bias-corrected-OLS), fully modified ordinary least squares (FM-OLS) and dynamic ordinary least squares (DOLS) under the homogenous covariance structure proposed by Kao and Chiang (2000). We used these three different methods to avoid and compare any estimation bias at longer horizons. The Kao's DOLS specification can be represented as follows;

$$\Delta s_{jt,k} = \alpha_j + \beta_D(i_{jt,k} - i^*_{jt,k}) + \sum_{q=-q_1}^{q=q_2} \delta_{jq}\Delta(i_{jt-q,k} - i^*_{jt-q,k}) + \mu_{jt,k} \quad \text{Equation 3.14}$$

where subscript j is an individual emerging economy. The parameter α_j is the member-specific intercept or a fixed effect parameter to cater for omitted variables that differs between markets but is constant over time. β_D is DOLS long-run parameter estimate and q is the number of lead and lag terms to correct the nuisance parameter in order to obtain coefficient estimates with nice limiting distribution properties as described in Kao and Chiang (2000).

3.4 Empirical Results

3.4.1. Time Series Analysis

In this section we report results for individual series estimations of Equation 3.1. First we graph the exchange rate changes and interest rate differentials relative to the US market for all emerging economies for $k = 1$, $k = 3$ and $k = 12$ in Figure 3.1,

[21] Please refer to Baltagi (2005) Chapter 12.3 for further discussion on the panel estimation that allow for cross-sectional dependence.

3.2 and 3.3 respectively.[22] Visual inspection clearly shows that in most markets these two variables move in opposite directions over the sample period and series tend to be more persistent for longer horizons.

Given that exchange rate movements for *k=1* and *k=3* are stationary in levels and stationary only in first difference for *k=12* (from previous section), we move on estimating the UIP hypothesis using standard OLS (Equation 3.1) and cointegration techniques, respectively. Table 3.3 depicts the results of country-by-country standard OLS coefficient (β_O) for *1*- and *3*-month maturity, while Table 3.4 presents the dynamic OLS (β_D) for 12-month maturity.[23] Since both exchange rate and interest rate differentials for *k=12* are of first differenced stationary series *I(1)*, it is necessary to check whether these two series are cointegrated or not to ensure the β_D estimates are efficient. The last column of Table 3.4 under the *ADF* heading shows the bivariate residual-based two-step cointegration test for *k=12* using the Augmented Dickey-Fuller (ADF) technique. All *ADF* statistics are much smaller than the critical values, leading to the conclusion that we reject the null hypothesis of a unit root for all estimated residuals for all emerging market models irrespective of their relative countries (the US, Japan or Germany). This finding confirms that the exchange rate changes and interest rate differentials in these markets are cointegrated. Therefore, the Stock-Watson parameter estimates of the long run parameter (β_D) are valid and not spurious. This time-series model (Equation 3.2) was estimated including up to $q = \pm 3$ leads and lags without altering the results to any significant degree.[24]

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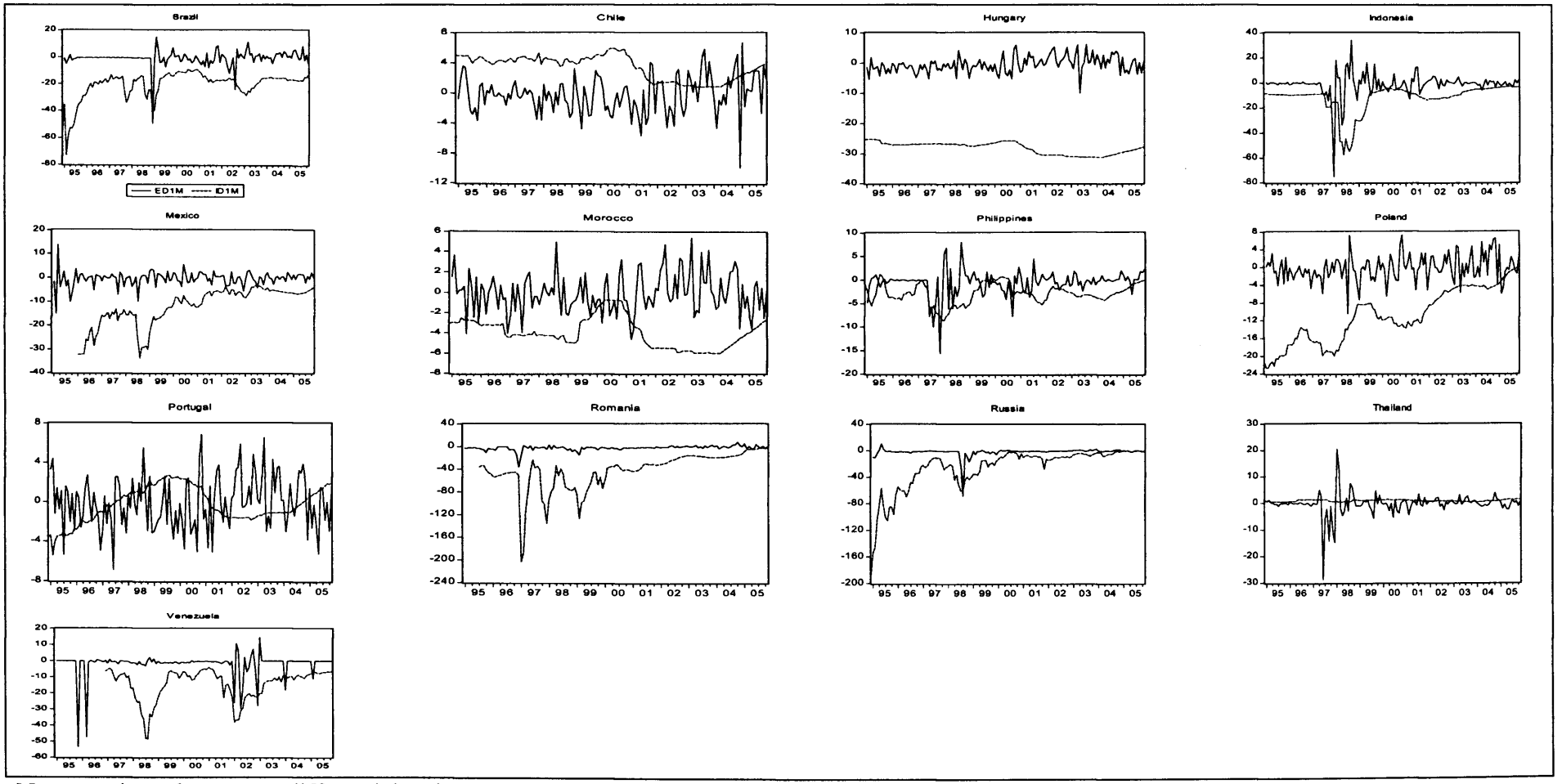
[22] Graph of exchange rate changes and interest rate differential relative to Japan and German markets have very similar characteristics.

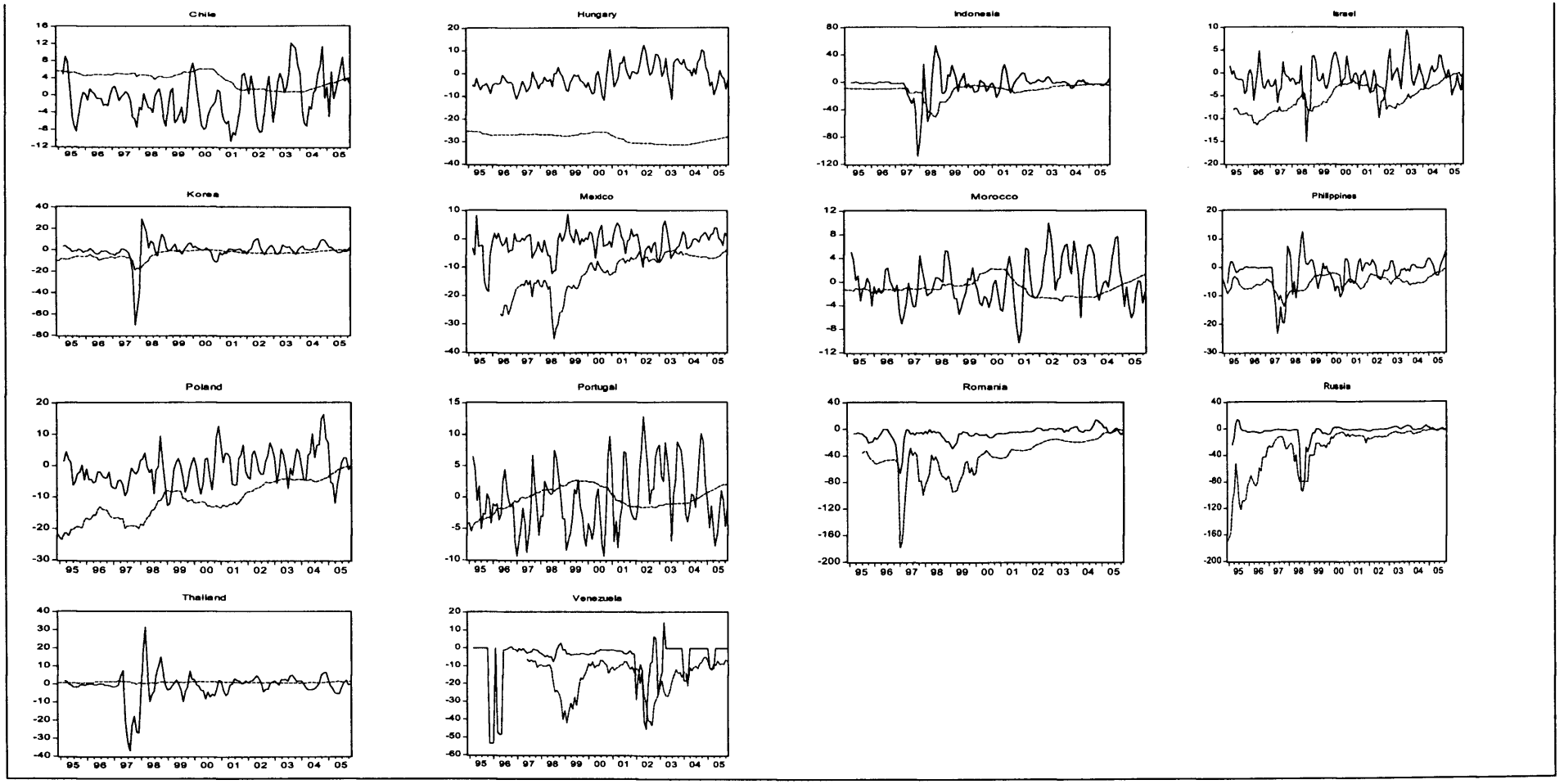
[23] The estimation

[24] However, δ_q is not reported in Table 3.4 for brevity purposes.



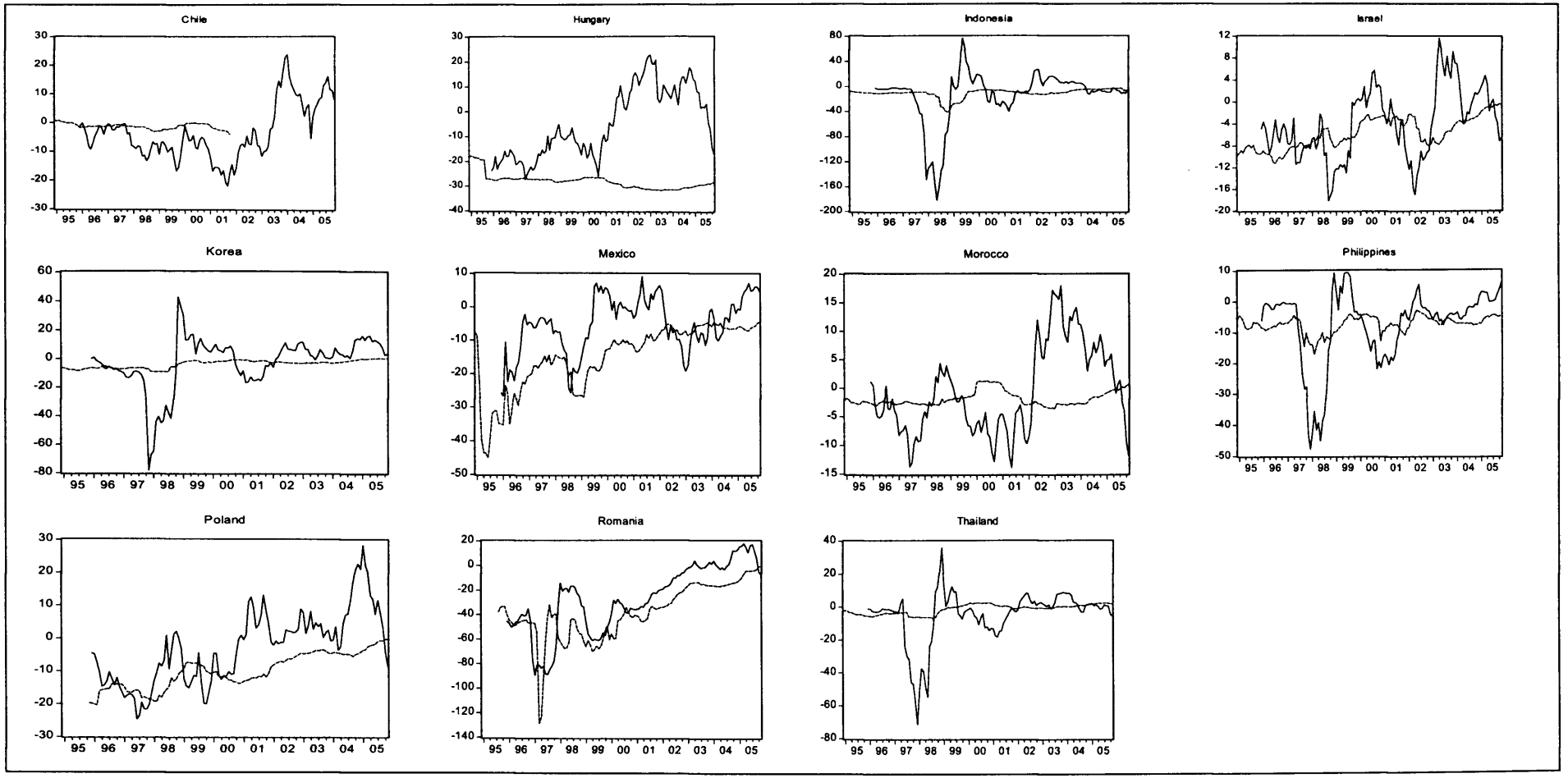
Note: — is exchange rate differential and ----- is interest rate differential.

Figure 3.1: Exchange Rate Changes and Interest Rate Differential of Emerging Economies and the US from 1995:01 to 2005:12 for *k = 1*



Note: — is exchange rate differential and ----- is interest rate differential.

Figure 3.2: Exchange Rate Changes and Interest Rate Differential of Emerging Economies and the US from 1995:01 to 2005:12 for $k = 3$



Note: — is exchange rate differential and ----- is interest rate differential.

Figure 3.3: Exchange Rate Changes and Interest Rate Differential of Emerging Economies and the US from 1995:01 to 2005:12 for *k = 12*

The striking result of the estimated coefficient for US-based regression, β (inclusive of both β_O and β_D), is that, at longer horizons (higher k), the UIP regression tends to produce estimates that are positive and not significantly different from unity. In Table 3.4 Panel A, when $k=12$, nine β_D are positive and statistically significant compared to only five and two for $k = 3$ and $k = 1$, respectively. Furthermore, five β_D estimates, out of nine are statistically not different from unity. The results discussed above are robust since the same pattern of results is also reported for the UIP regression under the Japan and Germany models (Panel B and C of Table 3.3a, Table 3.3b, and 3.4 respectively).

These results are consistent with previous empirical UIP testing in emerging markets where emerging markets' regression generally produces more favourable results compared to developed markets as documented in Bansal and Dahlquist (2000), Madarassy and Chinn, (2002) and Frankel and Poonawala, (2004).

However the results of OLS regression for $k = 1$ and $k = 3$ should be interpreted cautiously since inference based on OLS might not be appropriate in the case of regression between I(0) (the dependent variable; exchange rate changes) and I(1) (the independent variable; interest rate differential) series. Although the OLS estimate is consistent, the asymptotic distribution is non-standard. Using the OLS makes the standard errors of the estimate smaller and the t-statistics larger.[25] However, in $k = 1$ and 3, our estimation does not find much evidence of relationship between exchange rate movements and interest rate differential compared to the longer horizon $k = 12$. On this basis, that why we are not claiming that the UIP holds in short horizon.

[25] The argument is similar to the estimation of unit root testing using ADF tests. In ADF test, we regress changes of a variable I(0) on the level of the variable I(1). Dickey and Fuller (1979) show that under the null hypothesis of a unit root, this statistic does not follow the conventional Student's t-distribution, and they derive asymptotic results and simulate critical values for various test and sample sizes.

Table 3.3a: Ordinary Least Square (OLS) Regression of $\Delta s_{t,k} = \alpha + \beta_O(i_{t,k} - i^*_{t,k}) + \varepsilon_{t,k}$ for Individual Emerging Market for $k = 1$ from 1995 to 2005

Country	α	$SE(\alpha)$	β_O	$SE(\beta_O)$	$\beta_O = 1$	$\overline{R}^2$
A: US						
Brazil♣	0.919	(1.368)	0.070	(0.061)	0.000	0.089
Chile	0.663	(0.699)	-0.259	(0.170)	0.000	0.015
Mexico	0.236	(0.388)	0.046	(0.034)	0.000	0.008
Venezuela	0.654	(1.056)	0.142	(0.093)	0.000	0.012
Indonesia♣	-4.255*	(1.891)	-0.464	(0.236)	0.000	0.191
Korea♣	-	-	-	-	-	-
Philippines♣	-0.300	(0.457)	0.109	(0.194)	0.000	0.054
Thailand♣	1.340	(1.209)	-1.981	(1.322)	0.026	0.002
Israel	-	-	-	-	-	-
Morocco	-1.59***	(0.375)	-0.367***	(0.097)	0.000	0.083
Hungary	-15.90**	(2.788)	-0.551**	(0.101)	0.000	0.121
Poland	0.767	(0.598)	0.086	(0.043)	0.000	0.015
Portugal	-0.155	(0.248)	0.233*	(0.138)	0.000	0.030
Romania	1.681**	(0.706)	0.090***	(0.019)	0.000	0.552
Russia♣	0.284	(0.439)	0.046	(0.038)	0.000	0.246
B: Japan						
Chile	-0.246	(0.647)	-0.393	(1.369)	0.000	0.027
Mexico	0.034	(0.757)	0.022	(0.042)	0.000	0.014
Indonesia♣	-4.746	(2.649)	-0.366	(0.232)	0.000	0.110
Korea♣	-	-	-	-	-	-
Philippines♣	-0.146	(1.652)	0.049	(0.284)	0.000	0.072
Thailand♣	-0.105	(0.582)	0.115	(0.277)	0.001	0.019
C: Germany						
Brazil	0.010	(1.362)	0.017	(0.053)	0.000	0.141
Venezuela	0.144	(1.288)	0.101	(0.100)	0.000	0.023
Israel	-	-	-	-	-	-
Morocco	0.950	(1.704)	0.214	(0.330)	0.019	0.020
Hungary	-11.653	(11.004)	-0.385	(0.484)	0.000	0.028
Poland	1.331	(0.845)	0.150***	(0.060)	0.000	0.043
Portugal	0.136	(0.560)	0.386*	(0.231)	0.009	0.060
Romania	2.175**	(0.920)	0.101***	(0.022)	0.000	0.409
Russia♣	0.293	(0.595)	0.048	(0.033)	0.000	0.221

Note: $SE(\bullet)$ is Newey-West Standard Errors. $\beta_O = 1$ refers to *p-value* of the *F*-statistic. ***, ** and * indicate significant at 1 percent, 5 percent and 10 percent respectively.♣ Financial crisis dummy has been considered in the regression. – indicates non availability of dataset.

Table 3.3b: Ordinary Least Square (OLS) Regression of $\Delta s_{t,k} = \alpha + \beta_O(i_{t,k} - i^*_{t,k}) + \varepsilon_{t,k}$ for Individual Emerging Market for $k = 3$ from 1995 to 2005

Country	α	$SE(\alpha)$	β_O	$SE(\beta_O)$	$\beta_O = 1$	$\overline{R}^2$
A: US						
Brazil ♣						
Chile	1.327	(1.899)	-0.519	(0.442)	0.000	0.107
Mexico	0.203	(0.881)	0.123*	(0.059)	0.000	0.277
Venezuela	-0.413	(2.416)	0.251	(0.191)	0.000	0.112
Indonesia♣	-7.252*	(3.010)	0.716**	(0.343)	0.000	0.452
Korea♣	1.073	(1.461)	0.445	(0.389)	0.156	0.451
Philippines♣	1.228	(2.431)	0.588	(0.499)	0.411	0.221
Thailand♣	1.355	(2.822)	-3.086	(2.571)	0.000	0.150
Israel	-0.429	(0.671)	0.145	(0.100)	0.000	0.118
Morocco	-1.35***	(0.475)	-1.260***	(0.263)	0.000	0.189
Hungary	-50.56**	(6.494)	-1.740**	(0.236)	0.000	0.398
Poland	1.648	(1.644)	0.239*	(0.121)	0.000	0.044
Portugal	-0.165	(0.635)	0.833**	(0.364)	0.000	0.047
Romania	5.382**	(2.075)	0.288***	(0.057)	0.000	0.674
Russia♣	1.404	(1.382)	0.132	(0.088)	0.000	0.454
B: Japan						
Chile	-0.906	(1.162)	-2.499	(3.823)	0.000	0.019
Mexico	-0.374	(2.171)	0.034	(0.117)	0.000	0.033
Indonesia♣	-7.454	(4.149)	-0.555	(0.337)	0.000	0.410
Korea♣	1.160	(2.169)	0.2333	(0.325)	0.020	0.353
Philippines♣	0.192	(2.865)	0.203	(0.331)	0.018	0.119
Thailand♣	-0.228	(1.541)	0.070	(0.588)	0.000	0.057
C: Germany						
Brazil	-	-	-	-	-	-
Venezuela	-2.728	(2.760)	0.088	(0.190)	0.000	0.042
Israel	1.580	(1.808)	0.490*	(0.199)	0.012	0.120
Morocco	1.505	(2.191)	0.911	(0.879)	0.919	0.023
Hungary	-57.838	(37.030)	-1.956	(1.282)	0.000	0.059
Poland	3.729	(2.357)	0.446**	(0.174)	0.002	0.136
Portugal	0.349	(1.586)	0.677	(0.782)	0.681	0.023
Romania	7.23***	(2.688)	0.333***	(0.063)	0.000	0.569
Russia♣	1.204	(1.761)	0.122	(0.080)	0.000	0.438

Note: $SE(\bullet)$ is Newey-West Standard Errors. $\beta_O = 1$ refers to *p-value* of the *F*-statistic. ***, ** and * indicate significant at 1 percent, 5 percent and 10 percent respectively.♣ Financial crisis dummy has been considered in the regression. – indicates non availability of dataset.

Table 3.4: Stock-Watson Dynamic Ordinary Least Square (DOLS) Regression of $[\Delta s_{t,k} = \alpha + \beta_D(i_{t,k} - i^*_{t,k}) + \sum_{q=-q_1}^{q=q_2} \delta_i \Delta(i_{t-n,k} - i^*_{t-q}) + \varepsilon_{t,k}\ ]$ **for** $k = 12$ **for Individual Emerging Market from 1995 to 2005**

Country	α	$SE(\alpha)$	β_D	$SE(\beta_D)$	$\beta_D = 1$	$\overline{R}^2$	ADF
A: US							
Chile	-3.219	(1.611)	2.980***	(0.883)	0.028	0.189	-3.191**
Mexico	2.697	(2.493)	0.647***	(0.140)	0.013	0.339	-2.965**
Indonesia♣	2.390	(4.454)	0.859**	(0.412)	0.733	0.771	-3.080**
Korea♣	9.506**	(3.557)	2.587**	(0.691)	0.023	0.614	-3.91***
Philippines♣	7.700*	(4.603)	1.764***	(0.705)	0.280	0.702	-2.963**
Thailand♣	-3.973**	(1.855)	2.584**	(1.109)	0.155	0.431	-3.368**
Israel	0.317	(1.700)	0.897**	(0.276)	0.711	0.337	-2.747**
Morocco	-5.42***	(1.113)	-2.74***	(0.816)	0.000	0.193	-2.714**
Hungary	-219.1**	(16.969)	-7.495**	(0.604)	0.000	0.857	-4.51***
Poland	11.55***	(2.497)	1.575***	(0.191)	0.003	0.603	-3.038**
Romania	17.774***	(2.569)	1.167***	(0.106)	0.118	0.724	-5.67***
B: Japan							
Chile	87.141	(47.377)	14.974	(7.542)	0.039	0.123	-2.017**
Mexico	-5.019	(5.075)	-0.054	(0.257)	0.000	0.030	-2.894**
Indonesia♣	5.210	(6.378)	0.784	(0.440)	0.625	0.680	-3.185**
Korea♣	2.608	(4.400)	0.146	(0.639)	0.184	0.456	-3.82***
Philippines♣	8.234	(7.418)	1.373*	(0.723)	0.607	0.130	-2.70***
Thailand♣	-0.994	(3.323)	1.010	(1.041)	0.991	0.181	-3.07**
C: Germany							
Israel	6.758	(4.685)	1.849**	(0.550)	0.126	0.183	-2.651**
Morocco	6.057	(7.628)	2.590	(2.092)	0.448	0.055	-2.629**
Hungary	-550**	(77.530)	-18.59**	(2.643)	0.000	0.531	-2.008**
Poland	24.8***	(3.133)	2.88***	(0.220)	0.000	0.794	-3.76**
Romania	29.5***	(3.749)	1.462***	(0.124)	0.003	0.741	-5.19**

Note: $SE(\bullet)$ is Newey-West Standard Errors. $\beta = 1$ refers to *p-value* of the *F*-statistic. ***, ** and * indicate significant at 1 percent, 5 percent and 10 percent respectively.♣ Financial crisis dummy has been considered in the regression. ADF is unit root test for $\varepsilon_{t,k}$ of Equation 3.1 and test using the critical value from MacKinnon 1991.

3.4.2 Panel Analysis

The analysis is further complemented using the panel data framework under the same strategy employed for the time series. The panel technique is about increasing testing power and then boosting-up the confidence level on the estimated parameters.

Prior to testing for panel regression and cointegration, two panel unit root tests, as described in Section 3.3, were carried out. The results of the LLC and IPS tests are presented in Table 3.5 for both exchange rate and interest rate for $k=1$, $k=3$ and $k=12$. The results clearly show that the IPS w-statistics and LLC t^*-statistics reject the null hypothesis of a unit root at 5% significance level for both exchange rate and interest rate differentials for $k=12$ only at first difference. For $k=1$ and $k=3$; both IPS w-statistics and LLC t^*-statistics reject the null of unit root for exchange rate series. However, a mixed combination of 'reject' and 'fail to reject' IPS w-statistics and LLC t^*-statistics for interest rate differentials in levels. This is true whether we allow for a deterministic trend to appear in the unit root test specification or not. Generally, the results are consistent with individual series where both variables are differenced stationary $I(1)$ at $k=12$, while for $k=1$ and $k=3$, exchange rates are stationary at level but interest rates are only stationary at first difference. For $k=12$ we need to further confirm whether these two $I(1)$ variables are cointegrated or not to establish an efficient long-run relationship. Table 3.7 shows the bivariate panel cointegration test proposed by Kao (1999) and Pedroni (1999 and

Table 3.5: Panel Unit Root Test

Variables			US		Japan		Germany	
A: Level			LLC	IPS	LLC	IPS	LLC	IPS
Exchange rate	*k=1*	No trend	-35..8**	-33.11**	-26.9**	-27.2**	-16.4**	-15.3**
		Trend	-41.39**	-34.78**	-29.9**	-28.3**	-24.2**	-19.5**
	k=3	No trend	-5.88**	-8.46**	-3.55**	-8.46**	-3.22**	-5.60**
		Trend	-6.95**	-8.34**	-3.04**	-7.39**	-5.51**	-6.88**
	k=12	No trend	0.30	-0.86(13)	-0.33	-3.51**	-0.73	-0.86
		Trend	2.23	-1.32(2)	0.34	-1.28(1)	-1.12	-0.45
Interest rate	*k=1*	No trend	-2.71**	-4.26**	-0.91	-2.41**	-3.72**	-4.08**
		Trend	-1.24	-3.50**	-1.07	-2.47**	-3.28**	-4.15**
	k=3	No trend	-1.30	-2.29*	-1.06	-2.14*	-4.29**	-3.37**
		Trend	-0.23	-2.42**	-0.24	-2.70**	-3.08**	-3.46**
	k=12	No trend	-0.15	-1.00	-1.12	-1.05	-1.03	-1.23
		Trend	-0.49	-0.88	-1.26	-1.50	-0.80	-1.02
B: First Difference								
Exchange rate	*k=1*	No trend	-	-	-	-	-	-
		Trend	-	-	-	-	-	-
	k=3	No trend	-	-	-	-	-	-
		Trend	-	-	-	-	-	-
	k=12	No trend	-33.28**	30.82**	-31.4**	-	-16.8**	-15.2**
		Trend	-36.97**	31.31**	-34.9**	-30.9**	-18.8**	-14.9**
Interest rate	*k=1*	No trend	-	-	-18.3**	-	-	-
		Trend	-37.48**	-	-20.6**	-	-	-
	k=3	No trend	-32.31**	-	-26.0**	-	-	-
		Trend	-36.27**	-	-29.3**	-	-	-
	k=12	No trend	-28.78**	-26.79**	-24.9**	-21.63**	-19.97**	-17.50**
		Trend	-32.13**	-27.03**	-27.6**	-21.90**	-22.35**	-17.66**

Note: IPS is *w*-statistic from Im *et al* (2003) and LLC is t^*-statistic from Levin *et al* (2002). * and ** indicate significant at 5% and 1% level.

Table 3.7: Panel Cointegration

	US N=11	Japan N=6	Germany N=5
A: Kao (1999)			
DF_ρ	-7.539**	-11.35**	-6.03**
DF_t	-3.997**	-5.49**	-3.37**
DF^*_ρ	-17.42**	-24.03**	-14.95**
DF^*_t	-3.739**	-4.62**	-3.31**
ADF	-5.23**	-4.85**	-4.82**
B: Pedroni (1999;2004)			
	Intercept and no trend		
Z_ρ	-1.504	-0.843	-2.099*
Z_{pp}	-3.524**	-2.257*	-1.521
Z_t	-5.286**	-2.955**	-4.171**
$\tilde{Z}_\rho$	-0.663	-0.432	-0.897
$\tilde{Z}_{pp}$	-3.778**	-2.548	-0.688
$\tilde{Z}_t$	-5.318**	-2.830**	-2.905**
	Intercept and trend		
Z_ρ	-0.849	-0.194	-0.982
Z_{pp}	-2.061*	-2.109*	-0.448
Z_t	-2.996**	-2.444*	-2.048*
$\tilde{Z}_\rho$	-0.572	-0.477	-0.110
$\tilde{Z}_{pp}$	-2.817**	-1.811	-0.423
$\tilde{Z}_t$	-4.147**	-2.181*	-0.829

Note: Cointegration test statistics are calculated through the residuals from the panel OLS estimation. ** indicate significant at 1% level.

2004). All test statistics for Kao (1999) i.e. DF_γ, DF_t, DF^*_γ, DF^*_t and *ADF* reject the null of no cointegration at the 1% significance level for all models.[26] For Pedroni

[26] The testing of panel cointegration and the estimation of panel coefficient are carried out using the GAUSS programme (NPT 1.3 procedure written by Chiang and Kao (2002)and EViews 6.

tests statistics, as indicated by the panel non-parametric (Z_{pp}-statistics) and parametric ($\tilde{Z}_t$-statistics) as well as their group statistics, the null is rejected at the 1% level of significance for the US model. The Japan and Germany models are also supports for cointegration between these two variables. This finding confirms that at longer maturity periods, taken as a group, exchange rate and interest rate differentials are cointegrated and this could be an indication of the existence of the UIP.

The analysis is pursued therefore by estimating the cointegrating coefficient using panel bias corrected OLS, FMOLS and DOLS under the heterogeneous covariance structure proposed by Kao and Chiang (2000) for *k=12* and standard panel OLS for *k=1* and *k=3*. The results for estimated coefficients with their *t*-statistics in parentheses are presented in Tables 3.6 and 3.8, respectively. One main feature of the results is that the estimated interest rate differential coefficient has the correct sign as predicted by the hypothesis (positive) and is getting closer to unity at longer time horizons for all models. For instance, for the US-base model, as maturity (*k*) increases from *1* to *3* and then to *12*, β increases from 0.05 to 0.20 and 0.641 (for DOLS or 1.064 for bias-corrected and 0.681 for FM) respectively[27]. The other two models (Japan-base and Germany-base) produce the same pattern of β as *k* increases from *1* to *3* and *12*. This finding is more favourable than the existing literature where Bansal and Dalhquist found the pool coefficient on interest rate differential for developing markets for 3-month maturity to be 0.19. However Bansal and Dalhquist do not proceed further with longer maturity periods to show the pattern of β as *k* increases. Our finding, which is new for emerging markets, is quite similar to Chinn and Meredith (2005) who found the panel coefficient on interest rate differential to be around 0.674 at 5-year maturity for developed markets. This indicates that,

[27] We prefer DOLS estimated β as it is superior to the other two estimates as pointed out by Kao and Chiang (2000).

consistent with the individual series regression, the estimated coefficient of interest rate differential in emerging markets is positive and it is converging to unity at longer horizons of k.

Table 3.6: Panel OLS Regression of Uncovered Interest Parity for Emerging Markets from 1995 to 2005

	US		Japan		Germany	
	k=1	*K=3*	*k=1*	*K=3*	*k=1*	*K=3*
A: Fixed Effect						
α	-0.054	0.315	-0.313	-0.414	0.283	1.548**
	(0.160)	(0.274)	(0.387)	(0.706)	(0.244)	(0.420)
β	0.050**	0.200**	0.010	0.088	0.060**	0.232**
$se(\beta)$	(0.008)	(0.016)	(0.036)	(0.069)	(0.009	0.017)
$\beta = 1$	0.000	0.000	0.000	0.000	0.000	0.000
$\overline{R}^2$	0.047	0.145	0.007	0.012	0.071	0.225
Obs	1662	1759	643	760	1019	999
NoID	13	14	5	6	8	8
B: No Fixed Effect						
α	-0.125	0.046	-0.229	-0.324	0.206	1.198*
$se(\alpha)$	(0.145)	(0.358)	(0.327)	(0.575)	(0.235)	0.575
β	0.045**	0.176	0.020	0.098*	0.056**	0.214**
$se(\beta)$	(0.006)	(0.013)	(0.026)	(0.049)	(0.007)	0.016
$\beta = 1$	0.000	0.000	0.000	0.000	0.000	0.000
$\overline{R}^2$	0.045	0.098	0.003	0.014	0.064	0.165
Obs	1662	1759	1662	1759	1662	1759
NoID	13	14	13	14	13	14

Note: Panel regression of $[\Delta s_{it,k} = \alpha + \beta(i_{it,k} - i^{*}_{it,k}) + \varepsilon_{it,k}]$. $\beta = 1$ is the p-value of the *F-stat*. NoID refers to number of cross-sections. Number in parenthesis is White cross-section standard errors. * and ** indicate significant at 5% and 1% level respectively.

Table 3.8: Dynamic Panel Regression

A:US	β	T-Ratio	$\overline{R}^2$
OLS	0.988**	16.741	0.174
Bias-corrected-OLS	1.064**	12.754	0.173
FM-OLS	0.681**	8.091	0.157
Dynamic-OLS	0.641**	7.430	0.090
B: Japan			
OLS	0.748**	12.971	0.112
Bias-corrected-OLS	0.808**	9.264	0.111
FM-OLS	0.499**	5.673	0.100
Dynamic-OLS	0.566**	6.274	0.042
C: Germany			
OLS	1.371**	19.055	0.215
Bias-corrected-OLS	1.490**	13.426	0.213
FM-OLS	0.797**	7.128	0.177
Dynamic-OLS	0.813**	7.086	0.126

Note: All regressions include unreported country-specific constants. The reported *t*-statistics are biased corrected *t*-statistics. ** denotes that the coefficient is significant at 1% level.

3.5 Conclusion

In this paper, we re-examine the well-known empirical puzzle of UIP using a sample of emerging economies. In particular, we focus on testing whether rejection of UIP is driven by the typically short horizons used in empirical studies.

The major finding of the paper is that the majority of emerging economies with more flexible exchange rate regimes clearly indicate that at longer maturity periods, the β coefficients of interest rate differentials for both time series and panel regressions are positive and getting closer to unity, as stated by UIP. This finding confirms and expands earlier results by Bansal and Dalhquist (2000), Frankel and Poonawala (2004), and Chinn and Meredith (2004 and 2005).

In summary, complementing work on developed economies, this study has found a supportive ground to reconcile the theoretical-empirical puzzle of the UIP testing by adopting longer horizons for the exchange rate in emerging economies. This reveals that at the longer time horizon, the model has better econometric specification, and thus more predictive power for exchange rate movements compared to the shorter time period, as has been explained by Chinn and Meredith (2005). Success or failure in testing UIP is sensitive to the selection of the prediction time horizon, k. The findings can also be a signalling of well-integrated currency markets between emerging economies and developed markets and a reliable guide to international investors as well as for the orderly conduct of monetary authorities.

CHAPTER FOUR

LONG-HORIZON EXCHANGE RATES PREDICTABILITY IN EMERGING MARKETS

4.1 Introduction

The aim of this paper is to investigate the exchange rate forecastability puzzle that suggests that macroeconomic fundamentals contain negligible predictive content about the movements of nominal exchange rates. Since the seminal papers by Meese and Rogoff (1983a, 1983b), a lot of resources has been channelled into the refinement of theoretical models and advancement of econometric techniques to explain better the puzzle. However, the empirical evidence from mature economies has consistently failed to overturn this paradox. Consequently, clarifying the puzzle remains a challenging area for the researchers.

In this paper we give monetary models another chance and investigate whether by using a dataset from emerging economies can improve their forecasting performance. Our expectation is to find exchange rate predictability for countries with unstable macroeconomic fundamentals (see for example McNown and Wallace, 1994; Rogoff, 1996 and 1999a; and Moosa, 2000). The reason underlying this hypothesis is that countries with greater monetary instability are expected to show a stronger correlation between exchange rates and monetary fundamentals. Rogoff (1999a) argues that economically stable countries like United States, Germany and Japan generally experience very modest inflation rates. In such circumstances, it is difficult to identify the effect of monetary shocks on exchange rates. On the other hand, emerging economies experience high inflation rates, trade balance deficit,

budget deficit and excess money supply.[28] These relatively weak economic fundamentals, in addition to the poor management of the economy, are postulated to be crucial in predicting exchange rates under the monetary approach. Furthermore, most of the literature in the area of exchange rate predictability deals with developed and industrialised economies. Until now not much work has been done to investigate the forecastability of exchange rates in emerging economies despite their increasingly liberalised financial markets and their growing importance in the global financial system.[29]

This study differs from most previous studies in few ways. First, our sample is limited to emerging countries that satisfy two important assumptions of the exchange rate determination model: relatively floating exchange rate and considerably open economy for a long period to allow meaningful time series analysis. It does not mean that the emerging countries that we choose are fully liberalised, rather that the markets are satisfactorily open with little market frictions and government interventions. The countries we consider are Chile, Uruguay, Philippines, Thailand, Israel, Morocco, South Africa and Tunisia. According to Levy-Yeyati and Sturzeneggar (2003) these are countries that are adopting relatively floating exchange rate regime and on the process of liberalizing their capital account.

Second and motivated by Chinn and Meese (1995), we calculate the deviation from monetary fundamentals that are suitable for the emerging economies. In particular, we consider sticky price and relative price Balassa-Samuelson monetary

[28] Refer to Table 4.1 for comparison between income volatility and inflation rate between emerging markets and the US. Countries Chile, Israel and Uruguay are categorised as high inflation countries.

[29] Bulks of related works in the emerging markets are more concern on the subject other than forecasting exchange rate movements using monetary model. Among other issues of interest are optimal exchange rate regime, (Hochreiter and Tavlas, 2004; Alfaro, 2005), exchange markets integration, (Francis et al. 2002; Cheung et al. 2006; Rogers, 2006; Tai, 2007), exchange markets and financial crisis, (Phengpis, 2006; Kan and Andreosso-O'Callaghan, 2007), and exchange rate determinants, (Civcir, 2004; Candelon et al. 2006).

models to account for emerging country characteristics, as suggested by MacDonald and Ricci (2001). These models are expected to be superior to the standard flexible monetary model especially for countries which are still in the process of liberalization period (see Crespo-Cuaresma *et al.* (2005); Candelon et al. 2007).

Third, we use an error-correction framework to investigate both in-sample predictive content and out-of-sample point forecast accuracy of the fundamental-based models by employing the bootstrap technique proposed by Kilian (1999). The technique is able to account for small sample biases and size distortion that arise in the inferences procedure. Furthermore, the methodology is designed to differentiate whether forecastability power (if any) is due to the contribution of the explanatory variables or simply due to the drift term in the model.

The plan of the paper is as follows. Section 4.2 delves with literature reviews. In Section 4.3, we describe the process of constructing the fundamental variables, the dataset and the econometric procedure for testing predictability of exchange rate using the monetary models. Section 4.4 discusses the findings and the link between predictability and economic fundamentals of emerging economies. Section 4.5 concludes.

4.2 Literature Review

The study of exchange rates predictability was pioneered by Meese and Rogoff (1983a, 1983b). Their results suggest that none of the structural exchange rate models were able to forecast out-of-sample better than a naïve random walk model. Subsequently, extensive work has been carried out using various econometric techniques and different information sets to challenge the superiority of the random walk over monetary models of exchange rate determination. However, after more

than two decades of efforts, none of the out-of-sample empirical work finds consistent evidence of superior forecastability of structural models compared to the random walk.

Mark (1995) has given a new hope for exchange rate predictability by exploiting the assumed long-run linkages between exchange rates and monetary fundamentals. He finds significant evidence of forecastability at longer horizons (12 and 16 quarter). The same conclusion can also be found in Chinn and Meese (1995) who investigate the same issue using a larger set of explanatory variables. Chinn and Meese (1995) find that fundamental-based error-correction models outperform the random walk model for long term prediction horizons. However, both the econometric techniques and the results of Mark (1995) and Chinn and Meese (1995) have not been free from criticism. Kilian (1999) finds that Mark's results suffer from inconsistencies in the testing procedure and small-sample bias. Correcting for these drawbacks, Kilian (1999) finds no support for long run predictability of exchange rate. Later, Berkowitz and Giorgianni (2001) argue that the results of Mark are not robust and heavily depend on the assumption of cointegration in the long run series. Berkowitz and Giorgianni (2001) show that using the same dataset as Mark (1995) but under the unrestricted VAR model has produced very little evidence of predictability. Therefore, unpredictability of exchange rates remains if no prior assumption is imposed.[30]

Recent studies that use different information set and econometrics approaches (mostly departing from the traditional linear time series) to analyse the association of exchange rates and economic fundamentals do find encouraging support. For example, Kilian and Taylor (2003) use an Exponential Smoothing Threshold

[30] Comprehensive debate on the reliability of long-term exchange rate forecast can be found in Berben and van Dijk (1998), Groen (1999) and Rossi (2005), among others.

Autoregressive (ESTAR) model for seven OECD countries. They show the (in-sample) relevance of nonlinearities in exchange rate dynamics at the one- and two-year horizons. However, they still could not find support for out-of-sample predictability. Manzan and Westerhoff (2007) propose a chartist-fundamentalist model which allows for nonlinear time variation in chartists' extrapolation rate that provide support for the long-term predictability for five major currencies (German mark, Japanese yen, British pound, French franc and Canadian dollar) against the US dollar. Their study shows that the fundamentalist, together with the chartist, are correcting the deviation of exchange rate from its long run equilibrium path.

Faust et al. (2003) criticise the use of revised data for the fundamental variables and propose the use of real-time (unrevised) data. They argue that revised data can be used only if economic agents have the ability to predict future data (including the revision) correctly. However, this is not the case as Faust et al. (2005) among others, has shown that revisions to preliminary fundamental values are large and are unpredictable for some countries. Faust et al. (2003) empirically show that the exchange rate determination models that use real-time data are capable of explaining about 75% of the monthly directional changes of the US dollar-Euro exchange rate.

A comprehensive review of the empirical literature on the exchange rate unpredictability for industrialised nations over the last few decades can be found in Neely and Sarno (2002) and Cheung et al. (2005). The plausible explanations for the empirical failure of the exchange rate determination models include the instability of the parameters over the period, simultaneity problems, improper modelling of expectations formation and the failure of law of one price, among others. Following these dismal findings, exchange rate economists have drawn the conclusion that

exchange rate movements cannot possibly be attributed to macroeconomic fundamentals, at least in the short term. However, they have a firm belief that exchange rates cannot move independently from macroeconomic fundamentals over long horizons.

4.3 Exchange Rate Predictability and Emerging Markets

4.3.1 Evidence

Investigating the predictability of exchange rate movements using exchange rate determination model in emerging markets has not been an easy task. Empirical attempts are hampered by the difficulty to find an appropriate market that satisfies the assumption of free floating regime, free capital mobility and stable monetary regime.[31] Consequently, there is only relatively little empirical evidence of exchange rate forecastability in emerging markets during post-liberalization eras. These very handful empirical works also produce inconsistent results and therefore no concrete conclusion can be drawn from these limited findings.

For instance, Ferreira (2006) extensively investigates the significance effect of monetary fundamentals on the exchange rates for Chile, South Korea, Malaysia, Mexico, South Africa, Thailand and Turkey from 1992 to 2002 using panel cointegration techniques. He considers the sticky price model to account for the price rigidities effect between developed and emerging markets. The empirical evidence does not show any significant support for rejecting the hypothesis of no long run co-movement between exchange rates and monetary fundamentals across time and

[31] Chinn (1998) stresses the importance of capital imperfect mobility and substitutability, and instability of money demand that are widespread in developing countries in monetary modelling in emerging countries.

models.[32] Therefore the finding casts doubt on the validity of the hypothesis introduced by McNown and Wallace (1994) who find significant co-movement between exchange rate and monetary fundamentals in some emerging markets (Argentina, Chile and Israel). On the other hand, Wang and Wong (1997) use Kalman filtering techniques and ARCH models to address the issues of parameter instability and conditional variances to predict Japanese yen, Singapore dollar and Malaysian ringgit from 1973 to 1995. They find that the predictive power improves over 6 to 12 months forecasting horizons. The out-of-sample forecast errors are significantly lower compared to the naïve random walk model. Baharumshah and Masih (2005) further confirm this finding using cointegration techniques. They find substantial evidence of strong predictive power of the monetary model, both for in-sample and out-of-sample forecast accuracy. Based on the standard root mean square error (RMSE) and the Theil's *U* statistics, their findings suggest that the structural model performs better than the random walk only when the current account is included into the VAR system. They also find the error-correction term in the exchange rate equation enters with a significantly negative coefficient. This could suggest that exchange rates converge to the equilibrium path over longer period.

4.3.2 Monetary Models and Estimation Procedure

Theoretically, economists strongly believe that the exchange rate cannot deviate significantly from its "fundamental value". In other words, the exchange rate and its fundamental value are supposed to be cointegrated and one of the two variables will pull the other toward the equilibrium path. Therefore current

[32] Panel cointegration techniques have been employed in order to mitigate the problem of small sample bias and to increase the power of the statistical test. However, Neely and Sarno (2002) cast doubt on the validity of across countries estimation since currency values in different countries may be driven by very different forces such as monetary policy and exchange rate regime.

deviations of the exchange rate from its fundamental value should help predict future exchange rate movements. As such, they may be represented in a typical dynamic error-correction framework:

$$\Delta s_{t+k} = s_{t+k} - s_t = \alpha_k + \lambda_k (s_t - f_t) + \upsilon_{t+k} \quad k = 1, 8, 12 \text{ and } 16 \quad \text{Equation 4.1}$$

where s_t is logarithm of the nominal domestic-currency price of one unit of foreign exchange at time t. f_t represents the fundamental value of the exchange rate. α_k is a constant and λ_k is the predictability parameter to be estimated. k is the forecast horizon (3 months or quarter of a year) and υ_t is an *iid* disturbance term. If λ_k is smaller than 0, Equation 4.1 predict that the exchange rate should depreciate when $s_t > f_t$ in order to revert toward the equilibrium path. A statistical test of predictability of exchange rate at horizon k is thus carried out based on the null hypothesis of no predictability, $H_0 : \lambda_k = 0$, against the alternative hypothesis of predictability, $H_1 : \lambda_k < 0$. There are at least two econometric procedures often used to estimate exchange rate predictability namely, traditional linear and non-linear time series techniques. In this study we only consider the conventional linear time series methodology.

The estimation of Equation 4.1 is implemented in 2 steps. First step consists of obtaining the fundamental value f_t and the second is to estimate the forecasting regression. Specifically, first, we use Mark (1995) methodology to construct the fundamental value but with a few alterations to suit emerging market characteristics. Instead of imposing the theoretical value on the elasticity of money stock and income elasticity of exchange rate to [1, -1] respectively, the fundamental value f_t will be constructed using the estimated elasticity of money stock and elasticity of income from the estimated cointegrating coefficient of the Dynamic Ordinary Least Squares

(DOLS) method. After constructing the fundamental values then the forecasting estimation will be carried out employing the bootstrap procedure proposed by Mark (1995) and improved by Kilian (1999) under a constrained error-correction specification.

4.3.2.1 Construction of the Fundamental Values

The fundamental values f_t is constructed using cointegrating coefficients estimated by DOLS regression using the following specification:[33]

$$s_t = \alpha + \beta f_t + \sum_{j=-q}^{q} \delta_j \Delta f_{t-j} + \varepsilon_t \qquad \text{Equation 4.2}$$

where f_t is a vector of fundamental variables obtained from either one of the following three monetary models;

flexible price model

$$f_t = [(m_t - m_t^*), (y_t - y_t^*)] \qquad \text{Equation 4.3}$$

sticky price model

$$f_t = [(m_t - m_t^*), (y_t - y_t^*), (i_t - i_t^*), (\pi_t - \pi_t^*)] \qquad \text{Equation 4.4}$$

and **relative price model**

$$f_t = [(m_t - m_t^*), (y_t - y_t^*), (i_t - i_t^*), (\pi_t - \pi_t^*), (p_t^T - p_t^N) - (p_t^{T*} - p_t^{N*})] \qquad \text{Equation 4.5}$$

where m, y, i, π and p in Equation 4.3 , 4.4 and 4.5 represent the logarithm of money stock, the logarithm of real income, nominal interest rate, the CPI inflation rate and overall prices which include T, tradable, and N, non-tradable goods, respectively. An asterisk indicates foreign markets. β in Equation 4.2 is a vector of parameters of the

[33] For *I(1)* series with one cointegration relation, the DOLS estimation procedure produces efficient estimates of the cointegrating vector.

corresponding monetary models (flexible price, $[\beta_m, \beta_y]$; sticky price, $[\beta_m, \beta_y, \beta_i, \beta_\pi]$; and relative price, $[\beta_m, \beta_y, \beta_i, \beta_\pi, \beta_p]$). The β_m represents the elasticity of money stock, β_y is the income elasticity of output, β_p is the relative price elasticity, β_i and β_π are the interest and inflation semi-elasticity, respectively. The anticipated sign for the estimated coefficients are β_m, β_p and $\beta_\pi > 0$, while β_y and $\beta_i < 0$. Δ is difference operator and following Stock and Watson (1993) we set the number of leads and lags of the regressor (q) in the DOLS estimator of Equation 4.2 equal to three (q = 3). We use Newey-West procedure to compute robust standard errors.

The estimated cointegrating coefficients, $\hat{\beta}$s in Equation 4.2 are then used to construct the fundamental values based on the following models;

flexible price,

$$\hat{f}_t = \hat{\beta}_m(m_t - m_t^*) - \hat{\beta}_y(y_t - y_t^*) \qquad \text{Equation 4.6}$$

sticky price,

$$\hat{f}_t = \hat{\beta}_m(m_t - m_t^*) - \hat{\beta}_y(y_t - y_t^*) - \hat{\beta}_i(i_t - i_t^*) + \hat{\beta}_\pi(\pi_t - \pi_t^*) \qquad \text{Equation 4.7}$$

and **relative price**,

$$\hat{f}_t = \hat{\beta}_m(m_t - m_t^*) - \hat{\beta}_y(y_t - y_t^*) - \hat{\beta}_i(i_t - i_t^*) + \hat{\beta}_\pi(\pi_t - \pi_t^*) + \hat{\beta}_p[(p_t^T - p_t^N) - (p_t^{T*} - p_t^{N*})] \qquad \text{Equation 4.8}$$

Deriving fundamental values using the standard flexible price monetary model (Equation 4.6) is the most common procedure that has been extensively used by most of the researchers in the area, Mark (1995) and Kilian (1999) among

others.[34] However, it is less appropriate in the case of emerging markets since it requires domestic and foreign asset to be perfect substitutes and uncovered interest parity (UIP) condition to hold in the markets.

In this paper, we consider also two extension of the basic monetary model as suggested by Chinn (1998). First, following the work of Dornbusch (1976) and Frankel (1979), we consider a monetary model that incorporates short-term price rigidities (Equation 4.7). This model incorporates variables that allow for short run price stickiness that violates the condition of the Purchasing Power Parity (PPP) hypothesis. In addition, the relationship includes interest rates in order to capture the short term liquidity effect of the monetary policy. Second, we consider relative price movements by including the tradable and non-tradable goods within and across countries. Following Balassa (1964) and Samuelson (1964), the relative prices model is driven by relative differentials in productivity in the tradable and non-tradable sectors as presented in Equation 4.8. These two approaches, i.e. the sticky price monetary model and the relative price Balassa-Samuelson model, are expected to represent better the fundamental values of emerging economies. Furthermore, the inclusion of sticky prices and the Balassa-Samuelson effect in f_t could be crucial to find cointegration evidence in emerging markets.

Equation 4.1, combined with the structural models discussed above, result in the following predictability equations for $k = 1, 8, 12$, and 16:

Model 1:

$$\Delta s_{t+k} = \alpha + \lambda_k [s_t - (\hat{\beta}_m (m_t - m_t^*) - \hat{\beta}_y (y_t - y_t^*))] + \varepsilon_{t+k} \qquad \text{Equation 4.9}$$

[34] Mark (1995) and Kilián (1999) impose value of [1, -1] to β_m and β_y respectively.

Model 2:

$$\Delta s_{t+k} = \alpha + \lambda_k [s_t - (\hat{\beta}_m (m_t - m_t^*) - \hat{\beta}_y (y_t - y_t^*) - \hat{\beta}_i (i_t - i_t^*) + \hat{\beta}_\pi (\pi_t - \pi_t^*))] + \varepsilon_{t+k} \quad \text{Equation 4.10}$$

Model 3:

$$\Delta s_{t+k} = \alpha + \lambda_k [s_t - (\hat{\beta}_m (m_t - m_t^*) - \hat{\beta}_y (y_t - y_t^*) - \hat{\beta}_i (i_t - i_t^*) + \hat{\beta}_\pi (\pi_t - \pi_t^*) + \hat{\beta}_p ((p_t^T - p_t^N) - (p_t^{T*} - p_t^{N*})))] + \varepsilon_{t+k} \quad \text{Equation 4.11}$$

4.3.2.2 Forecasting Regression

We consider in-sample and out-of-sample forecast to evaluate the accuracy of monetary model in predicting exchange rate movements. Analysis of in-sample forecast (base on full sample from 1984Q1 to 2005Q4) of the monetary models (Model 1, 2 and 3) has been compared to random walk model of Equation 4.12 (as a benchmark)

$$s_{t+k} - s_t = d_k + \varepsilon_{t+k} \qquad k = 1, 8, 12 \text{ and } 16 \qquad \text{Equation 4.12}$$

of the corresponding k and tested for $H_0 : \lambda_k = 0$ against $H_1 : \lambda_k < 0$ or based on joint test of all forecast horizon as $H_0 : \lambda_k = 0 \ \forall k$ against $H_1 : \lambda_k < 0$ for some k. On the other hand, for out-of-sample forecast, we use prediction mean-squared error of Equation 4.9, 4.10, 4.11 and 4.12 from the sequence of recursive forecasts to evaluate the Theil's U-statistic and DM statistic of Diebold and Mariano (1995) with and without drift. Specifically, the estimation starts from 1984Q1 to 1995Q4. To generate the next forecast k, the estimation sample is updated by one period 1996Q1 for $k = 1$, 1997Q4 for $k = 8$, 1998Q4 for $k = 12$ and 1999Q4 for $k = 16$. The procedure is repeated until we reach the end of the sample in 2005Q4.

However, forecasting exercise based on Model 1, 2 and 3 involves some econometric difficulties. First, the error-correction representation is only appropriate under the assumption of stationarity of the error correction term $(s_t - f_t)$. This is because the asymptotic null distribution of test statistics for λ_k depends on whether the error-correction term is stationary or not, as discussed in Cavanagh et al (1995) and Valkanov (2003).

Another econometric problem is that forecasting involves future horizons k; when $k > 1$, the dependent variable $(s_{t+k} - s_t)$ represents overlapping sums of the original series that may result in high persistency of the error correction term. In this case, statistical inference should be handled with care since the in-sample R^2 and the t-statistics do not converge to a well-defined asymptotic distribution and the estimated coefficient, $\hat{\lambda}_k$, is biased away from zero due to size distortions. This bias is in favour of finding predictability as the forecast horizon (k) increases (see Mark and Sul, (2001), and Berkowitz and Giorgianni, (2001) among others, for detail discussions on the subject matter).

To mitigate the above discussed problems we consider bootstrap technique proposed by Kilian (1999) to approximate the finite sample distribution of the test statistic under the null hypothesis of no exchange rate predictability. This approach consist of first, estimating the Data-Generating Process (DGP) under the null of no predictability for the Constrained Vector Error Correction Model (VECM)

$$\Delta s_t = \alpha_s + u_{1,t} \qquad \text{Equation 4.13}$$

and

$$\Delta f_t = \alpha_f - h_2(f_{t-1} - s_{t-1}) + \sum_{j=1}^{q-1} \xi_j^{21} \Delta s_{t-j} + \sum_{j=1}^{q-1} \zeta_j^{22} \Delta f_{t-j} + u_{2,t}$$

Equation 4.14

using constrained Estimated Generalised Least Squares (EGLS) technique with all coefficient but α_s set equal to zero. The system also requires the restriction of $h_2 < 0$ to be satisfied to ensure estimation stability. The lag order q has been determined under H_o using AIC criterion.[35]

Second, after estimating Equations 4.13 and 4.14, a sequence of $\{s_t^*, f_t^*\}$, pseudo observations can be generated under the assumption of i.i.d. innovations using cumulative sums of the realizations of the bootstrap data-generating process. The process has been initialized by specifying $\left(f_{t-1}^* - s_{t-1}^*\right) = 0$ and $\Delta s_{t-j}^* = 0$ and $\Delta f_{t-j}^* = 0$ for j = q-1, ..., 1 and discard the first 500 observations. The pseudo innovation term $u_t^* = (u_{1t}^*, u_{2t}^*)'$ is random and drawn with replacement from the set of observed residuals $\hat{u}_t = (\hat{u}_{1t}, \hat{u}_{2t})'$. The process has been repeated for 2000 times. Third, use these $\{s_t^*, f_t^*\}$ of 2000 bootstrap replication to estimate the following long-horizon regression;

$$s_{t+k}^* - s_t^* = \alpha_k^* + \lambda_k^*(s_t^* - f_t^*) + \upsilon_{t+k}^* \quad k = 1, 4, 8, 12, 16 \qquad \text{Equation 4.15}$$

Finally, use the empirical distribution of these 2000 replication of the bootstrap test statistics to determine the p-value of the $t(20)$, $t(A)$, U, $DM(20)$, and $DM(A)$ of Equation 4.9, 4.10 and 4.11.

Regarding the potential problem of the serial correlation of the error term due to $k > 1$, we adopt two approaches. First we use Newey-West corrected t-statistics by setting the truncation lags to 20 since the longest forecast horizon is 16. Second, we use a data-dependent formula provided by Andrews (1991) under a univariate AR(1) as an approximating model. As a result, the statistical inference is robust to highly

[35] Further details explanation on the estimation procedures please refer to Appendix of bootstrap algorithm for long-horizon regression test in Kilian (1999).

persistent or near-spurious regression problems because it has the ability to automatically adjust the critical values to the increase in dispersion of the finite sample distribution of the test statistic for different lag structures and estimation procedures.

4.3.3 Data

In the present case, which is limited by the availability of fully liberalized emerging economies, we constrain ourselves to markets that satisfy two conditions i.e. floating exchange rate regime and relatively open capital markets for long period. Based on Levy-Yeyati and Sturzeneggar (2003, 2005), and supplemented with ratios of total external trade to GDP (see Table 4.1), we choose the following 8 emerging economies: Chile, Israel, Morocco, Philippines, South Africa, Thailand, Tunisia and Uruguay (along with the US economy as a base market).[36] Levy-Yeyati and Sturzeneggar (2003) classify 3 *de-facto* exchange rate regimes: float, intermediate and fixed. We choose only markets that are under float or intermediate regimes for the whole sample periods.[37] Float and intermediate regimes also indirectly indicate that the markets are not only open but characterised by little market frictions and government intervention. As defined by Levy-Yeyati and Sturzeneggar (2003), float and intermediate regimes are characterized by indices of low reserve volatility together with high exchange rate volatility. Low volatility of reserves is considered an indicator of less government intervention in the monetary policy. Therefore

[36] The definition of emerging market is based on the International Financial Cooperation (IFC). For more details explanation refer to Global Economic Prospects and the Developing Countries, *World Bank*, (2002), among others.

[37] We do include Philippines and South Africa in our sample though these two economies had fixed exchange rate regime on the following years, 1987, 1993, 1996 and 1990, 1993, 1995 respectively. Countries that experience more than three years of fixed exchange rates regime were excluded from the analysis. Full version of exchange rate regime classification can be access from http://200.32.4.58/~ely/index.html.

countries that have adopted a hard peg exchange regime, like China and Malaysia, or excessive capital control, like Korea, are excluded from the analysis.

Table 4.1: Economic Fundamentals for Selected Emerging Markets from 1984Q1 to 2005Q4

Country	Exchange Rate Regime		Income Volatility	Inflation	Total Trade
	Lowest	Highest			(% GDP)
Chile	1	2	2.75	11.63	60.32
Israel	1	2	2.27	41.12	83.41
Morocco	1	1	4.84	4.11	60.84
Philippines	1	3♠	3.89	9.91	79.11
South Africa	1	3♦	2.59	9.99	48.50
Thailand	1	2	4.78	3.61	90.73
Tunisia	1	2	2.60	5.00	87.36
Uruguay	1	2	5.08	43.41	42.22
United States	1	1	1.53	3.11	21.64

Classification of exchange rate regime is base on Levy-Yeyati and Sturzenegger (2003 and 2005). The index ranges from 1 = float; 2 = intermediate; and 3 = fixed. We do include Philippines and S. Africa since the fixed regime only for these three years ♠ 1987, 1993 and 1996; ♦ 1990, 1993 and 1995, respectively. The indices for the remaining countries and years are either 1 or 2. Income volatility is the standard deviation of the growth rate of GDP per capita. Inflation is a measure of mean inflation over the sample period. Total trade is an average of total import and export per GDP.

The variables considered in our monetary model are end of period quarterly nominal exchange rates expressed as the US dollar per emerging markets currency to proxy the nominal exchange rate (s_t), the money stock M2 to measure money supply (m_t), the Gross Domestic Product (GDP) is used to proxy output (y_t), the Consumer Price Index (CPI) is used as broad deflator (π_t), short term interest rate is proxied by inter-bank deposit interest rates (i_t), and the relative price of tradable and non tradable price deflator (p_t) is proxied by the ratio of CPI and Producers Prices Index (PPI) or Wholesale Price Index (WPI). The sample period considered in the analysis is from 1984Q1 to 2005Q4 and retrieved from either Datastream® or the IMF's

International Financial Statistics. All variables except interest rates are converted to natural logarithms.

4.4 Results

Unlike the earlier studies (for instance Mark 1995 and Kilian 1999), this paper does not impose theoretical value for the cointegrating coefficients in constructing the fundamental values (f_t). Instead, we use the estimated parameters obtained from DOLS regressions of Equation 4.2. Table 4.2 shows the estimated cointegrating coefficients that are used in constructing the fundamental values for all models and markets.[38]

We compute the Theil's *U*-statistics (the ratio of RMSE from two competing models, monetary versus random walk), the *t*-statistics and the Diebold-Mariano, (*DM*) statistics to assess the performance of exchange rate forecast using Model 1, 2, and 3.[39] The estimation results are presented in Table 4.3a and 4.3b for the drift-less random walk benchmark model while Table 4.3c and 4.3d for the random walk with a drift term. All the test results are presented in the form of bootstrap *p*-values based on 2000 replications. We are particularly interested in testing (in-sample) the hypothesis that $\lambda_k < 0$, and the out-of-sample performance based on the one-step ahead Diebold-Mariano *DM* test statistics and Theil's *U*-statistics. Long horizon

[38] This two-stage approach will be consistent provided that the first-stage estimates are consistent. Problematic estimation however could arise in the small sample estimation and therefore the results from small sample estimations should be interpreted with due caution.

[39] The estimation procedures were conducted using the MATLAB code provided by Lutz Kilian which is available in the Journal of Applied Econometrics data and code archive (1999) Volume 5.

Table 4.2: Cointegrating Coefficient Estimates Based on Dynamic OLS (DOLS), $s_t = \alpha + \beta f_t + \sum_{j=-q}^{q} \delta_j \Delta f_{t-j} + \varepsilon_t$ for 1984Q1 to 1995Q4.

Country	Flexible Price		Sticky Price				Balassa-Samuelson Effect				
	β_m	β_y	β_m	β_y	β_i	β_π	β_m	β_y	β_i	β_p	β_π
Chile	0.744	-0.232	0.896	-0.441	0.017	0.104	-0.333	1.210	0.005	-4.089	0.009
	(0.179)	(0.218)	(0.217)	(0.265)	(0.007)	(0.055)	(0.476)	(0.634)	(0.006)	(1.674)	(0.049)
Uruguay	-2.971	4.097	-2.796	3.879	-0.015	-0.024	-3.491	4.829	-0.045	-2.255	-0.005
	(0.053)	(0.066)	(0.239)	(0.289)	(0.036)	(0.013)	(0.521)	(0.658)	(0.059)	(1.051)	(0.015)
Philippines	0.679	-0.314	0.638	-0.277	0.012	-0.008	0.504	-0.027	-0.011	-1.148	-0.074
	(0.029)	(0.038)	(0.066)	(0.077)	(0.005)	(0.029)	(0.078)	(0.114)	(0.007)	(0.303)	(0.031)
Thailand	1.430	-1.494	1.588	-1.695	-0.016	-0.005	1.546	-1.599	-0.026	-1.607	-0.062
	(0.019)	(0.028)	(0.053)	(0.076)	(0.004)	(0.027)	(0.071)	(0.113)	(0.006)	(0.714)	(0.035)
Israel	-0.813	1.149	-0.877	1.229	0.008	0.001	-0.007	0.132	0.004	2.590	0.003
	(0.067)	(0.084)	(0.087)	(0.105)	(0.004)	(0.017)	(0.277)	(0.348)	(0.008)	(0.812)	(0.019)
Morocco	-1.025	0.672	-1.010	0.649	0.021	-0.030	-0.895	0.624	-0.023	-1.208	-0.033
	(0.156)	(0.075)	(0.359)	(0.190)	(0.033)	(0.018)	(0.447)	(0.263)	(0.097)	(2.162)	(0.029)
S. Africa	-1.110	1.390	0.254	-0.072	-0.011	-0.016	0.851	-0.815	0.003	2.706	-0.016
	(0.125)	(0.139)	(0.377)	(0.407)	(0.005)	(0.018)	(0.504)	(0.568)	(0.013)	(1.320)	(0.022)
Tunisia	0.108	-0.040	-0.348	0.372	-0.082	0.023	0.667	-0.525	-0.020	2.208	0.029
	(0.248)	(0.216)	(0.151)	(0.132)	(0.009)	(0.027)	(0.260)	(0.230)	(0.015)	(0.575(	(0.025)

Numbers in the parentheses are robust standard errors. Sample from 1984Q1 to 1995Q4. $q = 3$. β_m is the elasticity of money stock, β_y is the income elasticity, β_p is the relative price elasticity, β_i and β_π are the interest and inflation semi-elasticity, respectively.

predictability arises if the p values indicate increasing significance as the horizon k becomes larger. We are also interested in testing the joint significant of $\lambda_k = 0$ for all k at the 10% level.

Based on these criteria, the results show that only two countries (Israel and Uruguay) provide strong support for long horizon out-of-sample predictability. For Israel, the forecast accuracy is improving for longer horizons. This is evident from the *U*-statistics that are significant at k = 12 and 16 under the no drift sticky price model. In addition, the p value of the joint test of *Theil's U*-statistics is also significant. However none of the test statistics for Israel are significant when a drift term is considered in the models. In the case of Uruguay, the monetary models with a drift predict better the exchange rate movements. The joint test of *DM(20)* statistics for sticky price model and *DM(A)* for all three models with a drift are significant compared to none for the driftless case.

The result shows that there is evidence of the short horizon (k = 1 and 8) predictability of Chile, Uruguay and Morocco under the monetary models with a drift term. The out-of-sample test statistics (for k = 1) of all models are significant for Chile and Uruguay but only sticky price model fits the Moroccan market. Another obvious finding from the analysis is that the Chilean, Israeli and Uruguayan markets also provide significant support for in-sample predictability. The p values of *t(A)* and *t(20)* for some of the λ_k are significant (in the case of Uruguayan market, the in-sample predictability test statistics are significant for all models with drift term). For the remaining countries (Philippines, Thailand, South Africa and Tunisia), no predictability has been detected in the analysis.[40]

[40] The results from our analysis might be suffering from small sample biased towards non-predictability. Therefore it should be interpreted with due careful. The power of predictability is expected to be higher at longer span of dataset. However it is not always the

A number of interesting observations can be drawn from the results discussed above. First, the two countries (Israel and Uruguay) for which we find support of long-horizon predictability are characterized by high inflation (see for instance Braumann (2000) for high inflation countries classification and Table 4.1 for comparison between markets under study). The results confirm the earlier proposition made by McNown and Wallace (1994) and Rogoff (1999a) who argued that forecast accuracy using monetary models should be higher in countries with unstable macroeconomic fundamentals such as high inflation countries.

Second, inclusion of a drift term in the estimation has eliminated predictability from the Israelis market. The opposite result holds for Uruguay with the drift term. This shows the importance of considering drift or no drift in the estimation, as argued by Kilian (1999). Third, considering alternatives monetary models (sticky price and relative price models) has provided very useful information in the process of predicting exchange rates movements in emerging markets. At least the sticky price model seems to be superior to the standard flexible price and the Balassa-Samuelson model. This finding is similar to Chinn (1998) where he suggested the superiority of the sticky price model over relative price Balassa-Samuelson for Philippines peso and Thailand bath.

Finally, the finding of short-term predictability (k = 1 and 8) for Chile, Uruguay and Morocco is relatively new and surprising. This could be presumably a result of the instantaneous exchange market reaction to the instability of economic fundamental in the emerging markets. The linkages between these markets are further speed up by the rapid development in information technology and other economic factors such as trade linkages [Glick and Rose, (1999)], "common lender"

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case for emerging markets. One of the possible reasons for the biased is that the analysis does not cover the whole cycle of economic cycle.

Table 4.3a: Results of the VEC Bootstrap Model with No-Drift

		Flexible Price Model					*Sticky Price Model*					*Balassa-Samuelson Effect Model*				
Country	Horizon	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*
Chile	1	**0.059**	**0.017**	**0.071**	**0.081**	**0.074**	**0.029**	**0.010**	0.149	0.133	0.140	**0.043**	**0.013**	**0.064**	**0.076**	**0.067**
	8	0.157	**0.073**	0.295	0.266	0.308	**0.071**	**0.027**	0.230	0.204	0.225	0.123	**0.055**	0.222	0.207	0.237
	12	0.198	0.137	0.251	0.283	0.282	**0.080**	**0.046**	0.178	0.178	0.178	0.149	0.111	0.196	0.197	0.196
	16	0.241	0.216	0.287	0.625	0.389	**0.081**	**0.071**	0.174	0.135	0.138	0.146	0.128	0.207	0.229	0.218
	Joint	0.285	0.190	0.319	0.296	0.293	0.120	**0.077**	0.248	0.210	0.211	0.200	0.153	0.294	0.266	0.261
Uruguay	1	0.353	0.165	0.990	0.542	0.426	0.353	0.163	0.990	0.539	0.419	0.373	0.183	0.970	0.538	0.416
	8	0.340	0.330	0.980	0.668	0.720	0.336	0.327	0.990	0.663	0.722	0.359	0.349	0.990	0.660	0.724
	12	0.285	0.301	0.980	0.639	0.674	0.290	0.303	0.980	0.648	0.688	0.319	0.340	0.980	0.658	0.689
	16	0.210	0.221	0.980	0.638	0.675	0.216	0.232	0.980	0.639	0.673	0.239	0.267	0.980	0.631	0.660
	Joint	0.283	0.300	0.980	0.732	0.612	0.288	0.300	0.990	0.738	0.616	0.310	0.330	0.980	0.731	0.618
Philippines	1	0.577	0.613	0.487	0.567	0.423	0.750	0.755	0.384	0.396	0.316	0.661	0.670	0.391	0.449	0.333
	8	0.789	0.787	0.145	0.124	0.134	0.805	0.807	0.212	0.226	0.226	0.795	0.793	0.205	0.216	0.212
	12	0.811	0.812	0.208	0.233	0.220	0.838	0.839	0.325	0.544	0.448	0.816	0.813	0.281	0.474	0.386
	16	0.883	0.887	0.211	0.183	0.193	0.874	0.879	0.360	0.545	0.593	0.867	0.868	0.320	0.482	0.472
	Joint	0.773	0.809	0.289	0.266	0.278	0.851	0.852	0.475	0.424	0.417	0.814	0.819	0.513	0.410	0.406
Thailand	1	0.641	0.500	0.988	0.602	0.506	0.636	0.581	0.551	0.735	0.628	0.595	0.500	0.632	0.689	0.649
	8	0.748	0.751	0.265	0.676	0.550	0.722	0.720	0.327	0.765	0.606	0.752	0.747	0.456	0.968	0.896
	12	0.795	0.797	0.325	0.888	0.776	0.812	0.812	0.616	0.618	0.658	0.803	0.802	0.656	0.638	0.664
	16	0.829	0.827	0.460	0.636	0.673	0.842	0.844	0.654	0.529	0.573	0.852	0.852	0.704	0.517	0.554
	Joint	0.843	0.837	0.684	0.795	0.688	0.813	0.801	0.623	0.863	0.821	0.801	0.794	0.698	0.819	0.828

Note: The figure under *t(20)*, *t(A)*, *U*, *DM(20)* and *DM(A)* headings are bootstrap *p*-values for the VEC model with or without drift (Kilian 1999). Flexible price model, sticky price model and Balassa-Samuelson effect model have been considered to construct the fundamental variables. *t(20)* refers to *t*-statistic for the slope coefficient in the long-horizon regression with robust standard errors calculated based on a fixed truncation lag of 20. *t*(A) refers to the case of standard errors using Andrew (1991) rule. *DM* and *U* refer to the corresponding Diebold-Mariano and Theil's *U*-statistics (ratio of out-of-sample and random walk model) respectively. Results are shown for alternative forecast horizons k = 1-, 8-, 12- and 16-quarter. *Joint* refers to the *p*-value for the joint test statistics for all horizons. Boldface *p* values denote significance at the 10 percent level.

Table 4.3b: Results of the VEC Bootstrap Model with No-Drift

		Standard Monetary Model					*Sticky Price Model*					*Balassa-Samuelson Model*				
Country	Horizon	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*
Israel	1	0.547	0.493	0.134	0.149	0.150	0.146	0.122	0.177	0.129	0.126	0.424	0.359	0.102	0.121	0.124
	8	0.480	0.464	0.184	0.199	0.198	0.135	0.126	0.204	0.174	0.173	0.391	0.369	0.230	0.204	0.204
	12	0.321	0.319	**0.051**	0.122	0.124	**0.066**	**0.066**	**0.047**	**0.086**	**0.090**	0.256	0.246	**0.064**	0.123	0.124
	16	0.242	0.284	**0.081**	0.133	0.141	**0.060**	**0.073**	**0.085**	0.104	0.114	0.186	0.225	**0.080**	0.109	0.116
	Joint	0.313	0.358	**0.097**	0.205	0.211	**0.086**	**0.092**	**0.086**	0.151	0.152	0.245	0.279	**0.094**	0.167	0.177
Morocco	1	0.142	0.168	0.336	0.357	0.360	0.142	0.187	0.311	0.402	0.353	**0.058**	**0.058**	**0.096**	**0.094**	**0.086**
	8	0.072	0.073	0.146	0.143	0.145	0.155	0.151	0.261	0.819	0.806	0.104	0.120	0.214	0.240	0.230
	12	0.183	0.186	0.195	0.166	0.165	0.262	0.248	0.210	0.558	0.684	0.181	0.185	0.205	0.180	0.180
	16	0.197	0.203	0.313	0.376	0.407	0.284	0.268	0.332	0.411	0.508	0.176	0.185	0.335	0.380	0.422
	Joint	0.190	0.196	0.275	0.291	0.287	0.294	0.307	0.453	0.663	0.558	0.257	0.248	0.293	0.291	0.282
S. Africa	1	0.607	0.594	0.460	0.188	0.202	0.724	0.722	0.382	0.190	0.197	0.611	0.612	0.414	0.183	0.192
	8	0.496	0.467	0.536	0.271	0.290	0.583	0.556	0.414	0.255	0.273	0.489	0.464	0.506	0.255	0.278
	12	0.569	0.544	0.355	0.238	0.250	0.651	0.637	0.266	0.218	0.224	0.566	0.538	0.340	0.240	0.245
	16	0.689	0.683	0.212	0.215	0.215	0.751	0.749	0.134	0.175	0.170	0.679	0.675	0.188	0.204	0.202
	Joint	0.676	0.644	0.329	0.344	0.343	0.750	0.725	0.169	0.279	0.263	0.662	0.634	0.259	0.320	0.313
Tunisia	1	0.525	0.581	0.780	0.939	0.767	0.187	0.239	0.121	0.121	0.121	0.415	0.451	0.608	0.326	0.355
	8	0.508	0.519	0.659	0.521	0.566	0.268	0.260	0.316	0.231	0.236	0.305	0.325	0.696	0.564	0.626
	12	0.567	0.538	0.586	0.592	0.633	0.338	0.330	0.192	0.189	0.190	0.398	0.360	0.703	0.887	0.900
	16	0.705	0.703	0.295	0.295	0.306	0.497	0.497	**0.095**	0.135	0.121	0.587	0.591	0.331	0.313	0.328
	Joint	0.538	0.626	0.815	0.474	0.510	0.395	0.401	0.119	0.196	0.171	0.382	0.438	0.671	0.538	0.565

Note: Refer to note in Table 4.3a

Table 4.3c: Results of the VEC Bootstrap Model with Drift

		Flexible Price Model					*Sticky Price Model*					*Balassa-Samuelson Effect Model*				
Country	Horizon	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*
Chile	1	**0.030**	**0.010**	**0.024**	**0.043**	**0.026**	**0.044**	**0.013**	**0.013**	**0.040**	**0.024**	**0.054**	**0.016**	**0.012**	**0.059**	**0.037**
	8	**0.070**	**0.026**	0.237	0.221	0.222	0.122	**0.051**	0.252	0.225	0.228	0.149	0.067	0.349	0.266	0.271
	12	**0.080**	**0.047**	0.582	0.373	0.433	0.147	0.111	0.596	0.383	0.408	0.200	0.136	0.694	0.422	0.454
	16	**0.080**	**0.069**	0.771	0.484	0.495	0.144	0.126	0.798	0.515	0.507	0.237	0.207	0.859	0.572	0.555
	Joint	0.120	**0.077**	0.259	0.281	0.259	0.197	0.151	0.246	0.262	0.233	0.280	0.193	0.281	0.320	0.289
Uruguay	1	**0.098**	**0.029**	**0.003**	**0.025**	**0.018**	**0.099**	**0.027**	**0.003**	**0.025**	**0.017**	0.102	**0.028**	**0.003**	**0.027**	**0.017**
	8	**0.069**	**0.056**	0.126	0.120	0.121	**0.068**	**0.057**	0.126	0.119	0.122	**0.071**	**0.058**	0.125	0.120	0.122
	12	**0.056**	**0.044**	0.124	0.122	0.123	**0.055**	**0.045**	0.120	0.118	0.121	**0.057**	**0.050**	0.121	0.120	0.121
	16	**0.043**	**0.039**	0.153	0.150	0.150	**0.043**	**0.039**	0.152	0.151	0.153	**0.047**	**0.040**	0.151	0.148	0.151
	Joint	**0.045**	**0.040**	0.121	0.102	**0.098**	**0.045**	**0.040**	0.120	**0.099**	**0.095**	**0.049**	**0.041**	0.122	0.100	**0.095**
Philippines	1	0.746	0.752	0.778	0.843	0.789	0.628	0.650	0.700	0.588	0.744	0.310	0.344	0.230	0.412	0.583
	8	0.803	0.803	0.721	0.772	0.759	0.778	0.774	0.464	0.654	0.599	0.620	0.617	0.171	0.912	0.743
	12	0.833	0.836	0.813	0.910	0.920	0.833	0.833	0.592	0.879	0.918	0.602	0.600	0.197	0.922	0.930
	16	0.870	0.877	0.874	0.853	0.896	0.877	0.879	0.631	0.828	0.864	0.539	0.542	0.244	0.905	0.854
	Joint	0.851	0.849	0.834	0.942	0.903	0.739	0.764	0.726	0.772	0.735	0.466	0.531	0.320	0.633	0.757
Thailand	1	0.636	0.581	0.593	0.711	0.720	0.720	0.584	0.942	0.712	0.624	0.585	0.445	0.993	0.675	0.586
	8	0.722	0.719	0.390	0.898	0.867	0.798	0.804	0.429	0.948	0.877	0.677	0.680	0.291	0.259	0.272
	12	0.811	0.811	0.719	0.685	0.728	0.858	0.858	0.813	0.781	0.769	0.725	0.727	0.354	0.822	0.551
	16	0.843	0.843	0.792	0.626	0.670	0.891	0.891	0.919	0.644	0.671	0.754	0.759	0.511	0.645	0.667
	Joint	0.813	0.801	0.721	0.911	0.903	0.865	0.850	0.917	0.848	0.757	0.779	0.768	0.485	0.457	0.471

Note: Refer to note in Table 4.3a

Table 4.3d: Results of the VEC Bootstrap Model with Drift

		Flexible Price Model					*Sticky Price Model*					*Balassa-Samuelson Model*				
Country	Horizon	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*	*t(20)*	*t(A)*	*U*	*DM(20)*	*DM(A)*
Israel	1	0.148	0.114	0.435	0.267	0.455	0.398	0.313	0.217	0.217	0.224	0.532	0.476	0.292	0.289	0.313
	8	0.140	0.128	0.502	0.284	0.302	0.363	0.335	0.441	0.320	0.329	0.464	0.450	0.378	0.337	0.344
	12	0.068	0.069	0.211	0.223	0.223	0.228	0.224	0.282	0.277	0.277	0.302	0.306	0.275	0.314	0.312
	16	0.063	0.070	0.389	0.284	0.288	0.170	0.203	0.293	0.301	0.301	0.234	0.272	0.357	0.370	0.370
	Joint	0.089	0.090	0.383	0.427	0.426	0.220	0.250	0.413	0.438	0.438	0.300	0.331	0.404	0.480	0.477
Morocco	1	0.142	0.189	0.317	0.317	0.282	**0.058**	**0.058**	**0.060**	**0.075**	**0.061**	0.142	0.169	0.318	0.270	0.287
	8	0.153	0.151	0.236	0.232	0.234	0.104	0.120	0.135	0.177	0.185	**0.072**	**0.073**	**0.078**	0.146	0.147
	12	0.261	0.247	0.207	0.198	0.194	0.181	0.185	0.143	0.201	0.185	0.183	0.186	0.139	0.196	0.178
	16	0.282	0.266	0.497	0.391	0.424	0.175	0.184	0.543	0.401	0.435	0.197	0.203	0.492	0.370	0.385
	Joint	0.295	0.307	0.340	0.358	0.347	0.256	0.247	0.251	0.338	0.331	0.190	0.196	0.197	0.373	0.354
S. Africa	1	0.723	0.721	0.676	0.481	0.416	0.586	0.542	0.490	0.403	0.325	0.586	0.542	0.490	0.403	0.325
	8	0.584	0.555	0.655	0.501	0.532	0.543	0.522	0.529	0.376	0.395	0.543	0.522	0.529	0.376	0.395
	12	0.651	0.635	0.664	0.455	0.495	0.615	0.591	0.429	0.358	0.372	0.615	0.591	0.429	0.358	0.372
	16	0.750	0.750	0.684	0.545	0.589	0.747	0.747	0.414	0.356	0.367	0.747	0.747	0.414	0.356	0.367
	Joint	0.748	0.722	0.782	0.748	0.674	0.665	0.630	0.528	0.546	0.441	0.665	0.630	0.528	0.546	0.441
Tunisia	1	0.186	0.240	0.134	0.130	0.129	0.415	0.451	0.796	0.771	0.671	0.524	0.579	0.868	0.753	0.787
	8	0.265	0.261	0.562	0.618	0.693	0.305	0.327	0.852	0.968	0.989	0.506	0.520	0.799	0.943	0.977
	12	0.339	0.330	0.484	0.423	0.461	0.397	0.362	0.908	0.994	0.999	0.565	0.540	0.826	0.983	0.991
	16	0.500	0.500	0.382	0.359	0.372	0.588	0.592	0.775	0.973	0.987	0.706	0.704	0.712	0.982	0.992
	Joint	0.397	0.401	0.469	0.457	0.455	0.381	0.439	0.863	0.960	0.901	0.539	0.627	0.908	0.887	0.927

Note: Refer to note in Table 4.3a

or stock market [Kaminsky and Reinhart, (2001) and Caramazza *et al.*, (2000)], and "common macroeconomic weaknesses" [Eichengreen *et al.* (1996)]. The evidence is also in favour to the growing literature on the integration of currency market with other financial markets (Francis *et al.* (2002); Frankel *et al.* (2004); and Goldstein *et al.* (2000)) in emerging economies.

4.5 Conclusion

We consider emerging markets that are open and adopt floating exchange rate regimes to investigate the exchange rate forecastability puzzle using three different monetary models. The motivation for this study is based on the hypothesis proposed by McNown and Wallace (1994) and Rogoff (1999a). The hypothesis states that exchange rate predictability should be better off in countries with unstable monetary fundamentals. In addition to the standard flexible price model, we consider two alternatives approaches that account for sticky and relative prices. The method of Kilian (1999) has been employed to reduce problems in the long horizon finite sample forecasting estimations.

Based on Levy-Yeyati and Sturzeneggar (2003 and 2005), eight emerging markets have been chosen in the analysis to gain insight on exchange rate forecastability. The results suggest that the inclusion of fundamental values derived from the sticky price monetary model appears to improve the out-of-sample forecast accuracy of the exchange rate determination models for four emerging economies, Chile, Israel, Morocco and Uruguay. Empirical evidences are in favour of the hypothesis that markets with unstable monetary fundamentals such as high inflation have higher forecast accuracy compared to the random walk model.

Overall, predictability of exchange rates in emerging markets is very sensitive to the selection of appropriate models and the results are country specific in nature. For future research in emerging markets under the same issue, it may be fruitful to explore on the potential of short- or long-term forecast accuracy using non-linear specification.

CHAPTER FIVE

CONCLUSION

5.1 Summary of the Study

This dissertation is composed of three chapters that empirically investigate relevant issues related to the financial liberalization programmes in emerging economies. The study specifically examines the effects of financial liberalization on the volatility of equity market, validity of UIP and the predictability of foreign exchange market.

In the first chapter, employing different GARCH models (including ARCH, GARCH, EGARCH and TGARCH) on monthly returns of thirty emerging stock markets from 1984 to 2003, we find that the volatility effects after liberalization were of three categories i.e. some markets have experienced increasing, decreasing or unchanged volatility. The findings are in line with the current literature that finds volatility effect are market specific in nature. We further investigate why the driving forces of this phenomenon by grouping all the markets that experience increasing, decreasing or unchanged volatility separately. Analysing their level of institutional quality and market characteristics revealed that those markets that experienced increasing volatility after the liberalization are characterized by lower quality of market characteristics and institutions. The argument is reversed for countries with better institutions and market characteristics. The findings can reconcile contradicting views on the evolution of volatility effect after financial liberalization (decreasing or increasing) as reported in the literature if it is implemented with caution. It is suggested that it is important for the local authority to upgrade the domestic institutional quality including the respect for the rule of law, reduce level of

corruption, promote better contract enforcement, effective prudent regulation and supervision, and increase transparency before liberalizing the economy.

We then divert our analysis to the foreign exchange market in the second and third chapter. Second empirical chapter explores the validity of UIP for the selected 15 opened emerging markets. The study deviates from the previous works in emerging economy by using longer horizon data. Most of the existing studies in emerging economies use either 1- or 3-month maturity data and generally the results are inconclusive to the UIP hypothesis. In the empirical analysis, we employ dynamic ordinary least squares (DOLS) and bivariate dynamic heterogeneous panel regression based on Pedroni (1999 and 2004) and Kao (1999). Both dynamic time series and panel regression techniques are able to minimise the problem of persistent and near nonstationary series in long horizon regression. In contrast to the previous results, which have used short horizon data, the results of this chapter provide more support in favour of the UIP. The coefficients of interest differential for both time series and panel regression have of the expected sign and almost all are closer to the predicted value of unity at longer horizon. The results are also robust to changes in base country between the US, Japan or Germany.

The third empirical chapter investigates the exchange rate predictability puzzle using three versions of the monetary models: flexible price, sticky price and relative price. Our interest is to test the hypothesis whether economic fundamentals forecast better exchange rate movements in countries with monetary instability, such as emerging markets. Markets that experience unstable and volatile economic fundamentals are assume to have stronger correlation between exchange rates and monetary fundamentals. This study departs from most of the previous works in two important aspects. First, it limits the sample to emerging countries that satisfy two

important criteria: relatively floating exchange rate and considerably open economy for a long period (that allows for a meaningful time series analysis). Second, apart from flexible price, the study calculates the monetary fundamental values based on sticky price and relative price monetary models. The later two models are expected to forecast exchange rate movement better as they account the characteristic of the emerging markets. The empirical analysis of eight emerging economies (Chile, Uruguay, Israel, Morocco, South Africa, Tunisia, Philippines and Thailand) from 1984Q1 to 2005Q4 employs VEC bootstrap method of Kilian (1999). The bootstrap method is able to reduce problems of long horizon finite sample forecasting estimations. The overall finding indicates that a significant evidence for the hypothesis of predictability for countries that experience unstable economy.

5.2 Findings of the Study

Several observations have been derived from this dissertation.

First, the empirical results that show returns volatility could increase, decrease or be unchanged after the financial liberalization process depending on the level of institutional quality and market characteristics of the emerging markets. The results indicate that even though financial liberalization could bring benefits to the world economic system such as efficient risk sharing and better resource allocation, it also could bring economic fragility to certain markets that are not ready for liberalization. Therefore, policy makers need to concentrate to attain satisfactory level of quality in the institutional framework in order to reduce economic fragility and minimize market uncertainty. This is because countries with high quality institutions and market characteristics are more able to sterilise against the destabilizing effect of external shock than those with weaker ones. High quality of

institutions is also expected to improve the intermediation of sudden capital inflows or to cope with sudden capital outflows. Furthermore, this study could reconcile seemingly contradicting views on the effect of financial liberalization. Apart from the above mentioned findings, the study could be an important guideline for sequential process in the liberalization programmes. Financial liberalization should only be implemented after domestic institutional reforms in order to avoid excessive or unwanted volatility effects and not to jeopardise the wisdom of globalization. Incorrect liberalization sequencing was partly blamed as one of the vital reasons for the recurrence of financial crisis and the failure of liberalization programmes in most of the emerging markets. Therefore, the policy maker should prioritise the development of local institution and market infrastructure before other moves of liberalization.

Second, the study shows that by using longer maturity period of interest rate and exchange rate differential, the hypothesis of UIP may not be rejected in emerging countries. Most of the coefficients of interest rate differentials β, for longer k, for both time series and panel regression for all based countries (the US, Japan and Germany) are conform to the theory i.e. positive and closer to unity. This finding has helped us understand better the reasons lie in the theoretical-empirical puzzle of the UIP. It is postulated that at longer horizon exchange rates could have sufficient time to adjust to the changes in the fundamental variables and thus increase the predictive power of the model compared to the short horizon analysis. This chapter could be a complement to the works of Bansal and Dalhquist (2000) and Frankel and Poonawala (2004) who find forward parity puzzle might be confined to the developed economies and less present at the emerging currencies. The evidence also could be a reliable indication of the policy preference and inability of small but open

emerging markets to defend their monetary policies from the influence of international factors. In addition, at this point, we conjecture that the finding could be a significant signal for the presence a well-integrated currency market between emerging and developed economies.

Third, the study finds that in a more unstable economy like in the emerging economies, it is easier to forecast the movements of exchange rates using the macroeconomic variables. This is evident from the study when the hypothesis of no exchange rate predictability of four, (Chile, Israel, Morocco and Uruguay) out of eight emerging countries under study, has been significantly rejected at least at 10% level. The finding is considerably new and in favour to the proposition made by several economists such as McNown and Wallace (1994), and Rogoff (1999a). Another interesting observation from the finding is that three out of four countries that the exchange market is 'predictable' have been classified as high inflation nation. Those countries are Chile, Israel and Uruguay. The study also reveals the importance of a time trend in the regression model as suggested by Kilian (1999). Dropping the variable of time trend from the analysis has significantly altered the overall significance results of the study. The finding could be a complement to the literatures that supporting trend-following strategy in currency trading. Even though the existence of trends is a questioned in academic literature, there exist many empirical studies which show the presence of trends and the benefit following the trends strategies in currency trading (for example, Levich and Thomas, 1993; Lequeux and Acar, 2001; Pojarliev, 2005).[41]

[41] There are two traditional techniques in currency trading that remain popular. They are trend-following strategy and carry strategy. The trend-following strategy relies on the belief that currencies exhibit trends. On the other hand, carry strategy is based on the belief that the UIP does not hold and the forward rate is a bias predictor of spot rate. Usually the direction of change between future and spot rate are in opposite direction.

In conclusion, the study has shown the benefits and advantages of financial liberalization programmes in emerging economy especially in equity and foreign exchange markets. Liberalization has made the emerging market to be more integrated with the developed economies and has promoted the market to be one of the crucial players in the international financial system in the post-liberalization era. However, the implementation of financial liberalization should be carried out with prudent to avoid unnecessary drawback that could jeopardise the reputation of liberalization.

5.3 Future Research

Developments in emerging market such as financial liberalization in the mid of 80s and early 90s and then the financial reforms towards the end of 90s has left the study on the emerging market wide and open. Future research in the field of financial liberalization, volatility and institutions is worth to utilize panel data technique. The methodology is expected to have a very clear picture on the effect of institution and market characteristic across markets that are stable over period but diverse across markets. In the area of uncovered interest parity hypothesis, it is worth to fine tuning the emerging markets dataset. Since the UIP is an *ex ante* concept, it is better if we could drop the assumption of rationality and use a survey data. However these data are only available for highly funded research. Finally, the future research for predictability in the exchange rate movements is much more appropriate if considering non linear model. This is because much recent research has found support for non linearity behaviour in the exchange rate movement.

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