Essays on Exchange Rate and Inflation Dynamics

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Abstract

This thesis explores the relationship between the exchange rate and the domestic price level in three essays. The first essay (Chapter 2) examines the causality between the exchange rate and consumer prices, and estimates the extent of the exchange rate pass-through to consumer prices for 12 OECD countries for the period 1974 to 2016. Using the adoption of the Euro and the adoption of the policy of targeting inflation in these countries, which represent changes in the monetary policy regime, I divide this time period into two groups and examine causality and pass-through behaviour separately for each country. Based on a newly developed causality measure for multiple horizons, I found that the direction of causality from consumer prices to exchange rate becomes stronger for the countries with the Euro while the direction of causality from the exchange rate to consumer prices becomes stronger for the inflation targeting countries after their respective regime change. By deriving the impulse response functions from a recursive vector autoregressive model, I found that the exchange rate pass-through to consumer prices is not statistically different from zero for the countries with the Euro while the pass-through is statistically significant in four out of the six remaining countries. Before the regime change, the evidence on both fronts was somewhat mixed among these two sets of countries.

The second essay (Chapter 3) examines whether the aggregate price level responds asymmetrically to exchange rate appreciations and depreciations in 12 Asian countries for the period 1994 to 2016. Using a recently developed response-based test, I found evidence of asymmetric responses of the consumer price index to exchange rate appreciations and depreciations in 6 out of the 12 countries. The slope-based test also provides evidence of asymmetry for 6 countries, but the results are the same as the response-based tests only for 4 countries. Further, depreciations are not necessarily passed-through to prices more than the appreciations.

The third essay (chapter 4) examines the purchasing power parity (PPP) hypothesis for our selected 12 Asian countries for the period 1974 to 2016. Since stationarity of the real exchange rate implies that PPP holds, I employ unit root tests on the real exchange rate in the presence of multiple structural breaks. Our findings support the PPP hypothesis for six countries. Further, there is no additional evidence of trend stationarity of the series in these countries, so that there is no support for the Harrod-Balassa-Samuelson hypothesis.

Abrégé

Cette thèse explore la relation entre le taux de change et le niveau des prix intérieurs dans trois essais. Le premier essai (chapitre 2) examine la causalité entre le taux de change et les prix à la consommation et évalue l'ampleur de la répercussion des taux de change sur les prix à la consommation de 12 pays de l'OCDE pour la période 1974 à 2016. Avec l'adoption de l'euro et l'adoption de la politique de ciblage de l'inflation dans ces pays, qui représente les changements de politique monétaire, je divise cette période en deux groupes et examine séparément le comportement causal et le comportement de transmission pour chaque pays. Sur la base d'une mesure de causalité nouvellement développée pour plusieurs horizons, j'ai trouvé que la direction de la causalité entre les prix à la consommation et le taux de change devient plus forte pour les pays ayant adopté l'euro. Au contraire, la direction de la causalité du taux de change vers le prix à la consommation devient plus forte pour les pays ciblant l'inflation après leurs changements de régime respectifs. En dérivant les fonctions de réponse impulsionnelle d'un modèle vectoriel autorégressif récursif, j'ai constaté que l'impact des taux de change sur les prix à la consommation n'est pas statistiquement différent de zéro pour les pays euro, alors que l'impact est statistiquement significatif dans quatre des six pays restants. Avant le changement de régime, les éléments de preuve sur les deux fronts étaient quelque peu mitigés entre ces deux ensembles de pays.

Le deuxième essai (chapitre 3) examine si le niveau agrégé des prix répond de façon asymétrique aux appréciations et aux dépréciations de taux de change dans 12 pays asiatiques pour la période 1994-2016. En utilisant un test basé sur la réponse récemment développé, j'ai trouvé des preuves de réponses asymétriques de l'indice des prix à la consommation aux appréciations et aux dépréciations de taux de change dans 6 des 12 pays. Le test basé sur la pente fournit également des preuves d'asymétrie pour 6 pays, mais les résultats sont les mêmes que les tests basés sur la réponse seulement pour 4 pays. En outre, les dévaluations ne sont pas nécessairement reflétées dans les prix plus que les appréciations.

Le troisième essai (chapitre 4) examine l'hypothèse de la parité de pouvoir d'achat (PPA) pour nos 12 pays asiatiques sélectionnés pour la période 1974 à 2016. Étant donné que la stationnarité du taux de change réel implique que la PPA est maintenue, j'utilise des tests de racine unitaire sur le taux de change réel en présence de ruptures structurelles multiples. Nos résultats soutiennent l'hypothèse PPA pour six pays. De plus, il y avait peu de signes de stationnarité tendancielle de la série dans ces pays de sorte que l'hypothèse Harrod-Balassa-Samuelson n'est pas soutenue.

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Contribution of Authors

The essays of this thesis are solely my own work.

Professor Jagdish Handa has advised me from time to time as the supervisor of this thesis.

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Chapter 1

Introduction

An exchange rate is simply the rate at which transactions between two currencies take place. Under floating exchange rates, this rate is determined by the demand and supply of foreign currencies in every period. Any internal and/or external shock will be reflected in the exchange rate in a rather short period of time, if not instantaneously. The shocks can be economic or political. For example, during the period of the "Brexit" vote, the pound sterling depreciated against the US Dollar to a 31-year low. In that period, nothing fundamentally changed in UK but there was an expectation that the economic fundamentals will change in the future and the first hit was on the exchange rate. During the Asian financial crisis, the exchange rate of the East Asian countries depreciated by a large extent against the US Dollar. During the sub-prime crisis in the United States, the adjustment and re-adjustment of the exchange rate took place in many developed and developing countries. These changes in the exchange rate have effects on both the demand side and supply sides of the economy. For this reason, most central banks are tasked with a dual responsibility: exchange rate stability and domestic price stability.

The objective of this thesis is to examine the empirical relationship between the exchange rate and aggregate domestic price level. In particular, this thesis concentrate on three issues: (1) the direction of causality between the exchange rate and consumer prices, and the response of consumer prices to exchange rate shocks; (2) whether the response of consumer price index is symmetric to exchange rate appreciations and depreciations; and (3) whether the nominal exchange rate between two currencies varies proportionally with the relative price level in the two countries (i.e., whether purchasing power parity holds). These issues are discussed with three separate essays in Chapters 2 to 4. These essays make several empirical contributions to the literature, which are listed below.

Chapter 2 examines causality between (1) the exchange rate and consumer prices and (2) the exchange rate and short-term interest rate; and estimates the response of consumer prices and the interest rate to exchange rate shocks for 12 OECD countries (Australia, Canada, Finland, France, Germany, Italy, Netherlands, New Zealand, Spain, Sweden, UK and USA) for the period 1974 to 2016. All of these countries, except the USA, experienced a change in their monetary policy regime in the 1990s. The adoption of inflation targeting by 5 of these countries implies that, in order to achieve its target, the central bank will become very active in these countries. Another 6 countries adopted the Euro as their currency so that they could no longer pursue independent monetary policies. We divide the time period into two groups based on these structural changes and measure the causality and pass-through behaviour. This allows us to compare whether the pass-through and causality pattern changed for each country after the regime change and whether the pass-through and causality pattern is different in the countries with the Euro compared to the inflation targeting countries.

In chapter 2, we used a forecast based causality measure proposed by Dufour and Taamouti (2010) to measure the strength of causality at multiple horizons. To the best of our knowledge, this is the first paper to use this method to examine causality between (1) the exchange rate and the consumer price level, and (2) the exchange rate and the interest rate. We also estimate the response of consumer prices and interest rates to exchange rate shocks using a recursive vector autoregression (VAR) model. This adds another contribution to the literature as it adds additional evidence to the debate on whether the pass-through is decreasing across countries. A comparison of the two sets of countries provides further insights on the differences in pass-through behaviour among the inflation targeting countries and the countries with the Euro.

The aim of chapter 3 is to test for potential asymmetries in the responses of the consumer price index to exchange rate appreciations and depreciations for 12 Asian countries (Bangladesh, China, India, Indonesia, Japan, Republic of Korea, Malaysia, Pakistan, the Philippines, Singapore, Taiwan and Thailand). In particular, we test the following hypothesis of directional symmetry: does the consumer price index respond symmetrically to exchange rate appreciations and depreciations?

The empirical evidence on this issue is currently limited to a few developed countries. This chapter contributes to this issue by analyzing economies that are at various stages of economic development. Furthermore, all the previous studies use slope-based tests to examine the degree of asymmetry. This slope-based test does not provide insights on the responses of prices to exchange rate shocks. The responses, derived from a structural VAR model, of prices to exchange rate appreciations and depreciations, can be asymmetric even when the slope-based test provides evidence against asymmetry. This can be due to the fact that impulse response functions are highly nonlinear functions of the slope parameters and the size of the shocks. Therefore, the response-based test is more informative about the asymmetric adjustments of prices to exchange rate changes. We use a response-based test, recently proposed by Kilian and Vigfusson (2011), to examine asymmetric responses of consumer prices to exchange rate appreciations and depreciations. To the best of our knowledge, this is the first paper to examine this issue by a response-based test.

In chapter 4, we examine whether the purchasing power parity (PPP) hypothesis holds for 12 Asian countries, located in East, Southeast and South Asia (same as the 12 countries of chapter 3). Since stationarity of the real exchange rate implies that PPP holds, this chapter employs unit root tests, both without and with structural breaks.

The contributions of this chapter are: (1) compared to the earlier studies, we examine stationarity of the real exchange rate for a longer period of time and with more efficient unit root tests. This provides a new benchmark for future studies. (2) For our selected countries, the previous empirical studies that include structural breaks into the model use tests that only examine trend stationarity of the real exchange rate, while we examine level stationarity of the series allowing for multiple breaks. This is a more appropriate test to examine the PPP hypothesis since trend stationarity of the real exchange rate provides support for the Harrod-Balassa-Samuelson (HBS) hypothesis, which is an explanation of the failure of the PPP to hold. (3) All the previous empirical studies on our selected countries assume gradual adjustment of the trend function due to structural break(s), while this chapter makes no such assumption: it uses tests based on both the gradual adjustment of the trend function and instantaneous adjustment of the trend function. Examinations of such models are necessary since structural shocks can affect the real exchange rate differently across countries. (4) We also examine trend stationarity using models of both types of trend adjustment. This allows us to compare our results with those of earlier studies on the validity of the HBS hypothesis and provide a new point of reference for future empirical research.

The final chapter (Summary and Conclusions) of this thesis summarizes the major findings of each essay and discusses future avenues for research.

Chapter 2

Exchange Rates and Consumer Prices: Measuring the Causality and Pass-Through Behaviour in 12 OECD Countries

2.1. Introduction

The objective of this chapter is to examine the causal relationship between the exchange rate and the aggregate price level; and to estimate the degree of exchange rate pass-through (ERPT) to the aggregate price level. Understanding this relationship is important for the formulation of macroeconomic policies. These two variables also determine the real exchange rate, which in turn determines the competitiveness of an economy in international trade and, hence, the trade balance. Broadly, ERPT refers to the extent of change in the domestic price indices for a 1 percent change in the exchange rate. If the response is one-to-one, the pass-through is known as complete; if it is less than one, the pass-through is incomplete or partial. Almost all of the previous empirical studies find evidence of incomplete pass-through.

A priori, the ERPT to import prices should be higher than to producer or consumer prices as imported goods are only one component of all goods and services sold (or an intermediate good in the production process) in the economy. Empirical evidence (e.g., McCarthy (2007) for industrialized countries, and Ito and Sato (2008) for Asian countries) also supports this expectation. Firms' pricing strategies and market structure play a major role in ERPT. If prices are set in the exporter's currency, known as *Producer Currency Pricing*, and prices are sticky, then ERPT to import prices would be complete. If the foreign country's export prices vary considerably with the exchange rate and import prices in the domestic market remain fairly stable, this shows the presence of *Local Currency Pricing*[1].

In perfectly competitive markets for homogeneous goods, price equals marginal cost so that there is no mark-up for firms in the pricing equation. In this case, the ERPT will be complete. In imperfectly competitive markets, firms can charge a mark-up over their marginal cost and can

[1] Also known as pricing-to-market (PTM), a term coined by Krugman (1987).

adjust this mark-up in response to changes in the exchange rate or marginal cost; thus paving the way for local currency pricing, which results in incomplete pass-through. This mark-up depends on the functional form of the demand curve, the substitutability between domestic and foreign goods and market segmentation. Along with these microeconomic factors, macroeconomics factors such as the exchange rate and price stability also play major roles in pass-through behavior. We discuss these factors in greater detail in section 2.2.

For the empirical analysis of ERPT to domestic prices, one strand in the literature estimates a single equation where the domestic price is a function of a set of explanatory variables and the exchange rate is one of the explanatory variables (e.g., Knetter (1993), and Campa and Goldberg (2005)), while another strand employs a structural vector autoregression (VAR) framework where all variables are treated as endogenous (e.g., Choudhri, Faruquee and Hakura (2005), and McCarthy (2007) among others). VAR estimation is preferable to single equation estimation since it enables simultaneous identification of the effect of exchange rate shocks on the selected price indices and uses the covariance structure among the endogenous variables. Further, single equation estimation tends to suffer from the endogeneity problem. Presenting dynamic stochastic general equilibrium (DSGE) models in the VAR framework is also popular in the contemporary literature. However, this approach uses prior information from the specified DSGE model; and thus, can impose a very strict structure on the data.

Structural VAR models can use recursive or non-recursive identification scheme to identify the structural shocks. The recursive model uses the Cholesky decomposition to derive the structural shocks. On the other hand, a non-recursive model relies on the structural shocks derived from prior information, such as from previous studies, or by estimating the parameters that explain causal effects. Uhlig (2005) proposes an alternative procedure where the focus is to identify the shock(s) of interest rather than all the shocks in the VAR model. This is achieved by imposing sign restrictions on the impulse responses of some variables in response to shocks in some other variables in the model. An and Wang (2012) use this sign restriction approach to examine ERPT to domestic prices for nine OECD countries.

In this chapter, we examine (1) causality between the exchange rate and consumer price index (CPI) and (2) the extent of ERPT to CPI for 12 OECD countries for the period 1974 to 2016. The countries are: Australia (AUS), Canada (CAN), Finland (FIN), France (FRA), Germany

(GER), Italy (ITA), Netherlands (NET), New Zealand (NZL), Spain (SPA), Sweden (SWE), United Kingdom (UK) and the United States of America (USA). These countries are at similar stage of economic development and went through a noticeable structural change in their monetary policy regimes in the 1990s. Finland, France, Germany, Italy, Netherlands and Spain adopted the Euro as their currency in January 1, 1999. On the other hand, Australia, Canada, New Zealand, Sweden and UK officially started targeting inflation in the early 1990s.[2] Due to the change in monetary policy regime, there is a large change in the role of the central banks of the respective countries. In the inflation targeting countries, the central bank is expected to become very active to achieve the inflation target. On the other hand, countries that adopted Euro as the currency can no longer pursue independent monetary policies. This selection of countries allows us to compare whether the causality and pass-through behaviour changed for each country after the regime change and whether the countries that adopted the Euro behave differently to exchange rate shocks when compared with the inflation targeting countries.

To examine causality, we use the Granger causality test and a forecast based causality measure proposed by Dufour and Taamouti (2010). To best of our knowledge, this is the first paper to examine the causality between exchange rate and consumer prices using this causality measure. Using the recursive VAR model, we estimate the response of consumer prices to exchange rate shocks over a horizon of three years. We also examine whether the causality and pass-through relationship between the exchange rate and the interest rate altered after the regime change. This also allows us to verify whether the central banks of the inflation targeting countries behaved differently when compared to the European Central Bank (ECB) to exchange rate shocks. For the countries with the Euro, we expect the causality and pass-through behaviour to show a similar pattern, though with some variation since the nominal effective exchange rate (NEER) is determined by the weight of the exchange rate with the country's trading partners. Further, movements in the CPI among these countries may not be synchronized because of differences in transaction costs, non-tradable services and the composition of their consumption basket. We also expect that the causal relationship between the exchange rate and domestic price level to be weaker for the countries with the Euro compared to the inflation targeting countries. Although both the currency union and inflation targeting will increase monetary stability, the currency union will

[2] In addition, we included USA in our analysis to examine whether any differences exist in its pass-through behavior since it does not target inflation officially but achieved low and stable inflation over the period.

increase trade among the participating countries because of Geographical proximity and free-trade. Also, as the market size increase in the participating countries, foreign exporters will adjust their mark-up to keep price constant in the currency union.

The overview of our main findings is as follows. We found that the direction of causality from CPI to the exchange rate is stronger in the countries that adopted the Euro while for inflation targeting countries the direction of causality from the exchange rate to the CPI has more strength. The strength of the direction of causality has changed for some of the countries after both types of regime changes. We also found that the ERPT to CPI is not statistically different from zero for the countries that adopted the Euro as their currency. Before the adoption of Euro, the response was statistically different from zero for some of these countries. On the other hand, the responses of CPI to exchange rate changes do not exhibit any distinctive pattern for the inflation targeting countries after the regime change. In all countries, after the regime change, the interest rate increases in response to exchange rate appreciation; however, the response are not statistically different from zero for some countries.

The remaining chapter is organized as follows. Section 2.2 presents a brief review of the empirical literature. Section 2.3 discusses the data. Section 2.4 provides the results from the causality analysis. We present our econometric methodology and results of the VAR model in Section 2.5. Section 2.6 presents the concluding remarks.

2.2. Literature Review

The Granger causality tests, introduced by Granger (1969), examine the predictability at horizon 1 of a variable X from its own past, the past of variable Y and the past of a set of auxiliary variables. For example, the exchange rate is said to Granger cause the inflation rate if the inflation rate can be predicted better using the past values of exchange rate and inflation rate, rather than only with the past values of inflation rate. A point to be noted here is that Granger-causality does not necessarily measure true causality. It only examines whether the lagged values of one variable help to predict another variable and if they do so, it establishes causality in the 'Granger-causality sense'.

An important extension of the Granger causality test was proposed by Dufour and Renault (1998) who generalized the notion of Granger causality to a horizon h $(1 \leq h \leq \infty)$ in the presence of indirect causality through the auxiliary variables. This indirect causality can distinguish between short-run and long-run (non)causality. For example, even if Y does not Granger cause X at horizon one, it may help to predict X at horizon $h > 1$ through transmission via a set of auxiliary variables. These auxiliary variables can cause indirect causality between the X and Y variables at a horizon higher than 1, even if there is no direct causality at horizon 1. This is known as conditional causality. If the auxiliary variables are dropped from the information set, the representation becomes unconditional causality. For the case of unconditional causality, non-causality at horizon 1 implies non-causality at any other horizon.

Instead of tests for non-causality at horizon h, Dufour and Taamouti (2010) propose a procedure for measuring the causality between two vector processes. This causality measure allows for both conditional and unconditional causality. By construction, this causality measure is nonnegative and becomes zero when there is no causality at horizon h. The higher the value of the measure, the stronger is the causal relationship, so that this measure can reveal the strength of the direction of causality. Additionally, the confidence interval of the causality measure can help to determine how long the causal effects will last. The causality measure is statistically significant if the 95% confidence interval does not include zero. The consistency and asymptotic normality of this estimator is provided in Dufour and Taamouti (2010). Estimating the standard error of this estimate is typically difficult, and Dufour and Taamouti (2010) suggested a bootstrap procedure to compute the confidence interval for this causality measure. In their empirical work, they examined the causality between monetary policy and real GDP and found that the direction of causality from monetary policy to real GDP has more strength.

As already mentioned, we did not find any study that uses this causality measure to examine the causality between the exchange rate and CPI or between the exchange rate and the interest rate. The most relevant study for our work in this chapter is Zhang, Dufour and Galbraith (2016) who examine the causal relationships between three commodities (crude oil, gold and copper) and the exchange rate for four countries (Canada, Australia, Norway and Chile) over the period 1986 to 2015. Using the Granger causality test, they found evidence of Granger causality in both directions. However, using the causality measure of Dufour and Taamouti (2010), they found that the measured intensity of the causality is much stronger in the direction of commodity prices to exchange rates, especially at horizon one.

Causality measures can predict the strength of the direction of causality but do not shed lights on the response of one variable to changes in any other variable. For that purpose, we will look into the impulse response functions (IRFs) derived from a VAR model. For rest of this section, we will review the factors that determines the degree of pass-through and review some empirical studies on some developed countries to get an idea on the degree of ERPT.

The import prices for any country P_t^m equals the export price of trading partner P_t^* converted to domestic currency using the exchange rate E_t (domestic currency per unit of foreign currency): $P_t^m = E_t P_t^*$. Exports prices for foreign firms are a markup over their marginal cost (MC): $P_t^* = (markup_t^*)(MC_t^*)$. Substituting this relationship in the previous equation, taking logs and using the lower case to denote variables in logs, we get: $p_t^m = e_t + markup_t^* + mc_t^*$. Allowing markup to have industry-specific fixed effects and sensitivity to exchange rate, this can be written as: $markup_t^* = \Phi + \Theta e_t$. Hence, the import price equation becomes: $p_t^m = \Phi + (1+\Theta)e_t + mc_t^*$. If $\Theta = 0$, producer currency pricing takes place and pass-through is found to be complete. If $\Theta = -1$, local currency pricing take place and exporters fully absorb the fluctuations in the exchange rate in their markup. Consumer prices of the final good producer can be written as the weighted average of the price of domestically produced intermediate goods ($P_{h,t}$) and imported intermediate goods ($P_{m,t}$): $P_t = \left[\alpha P_{h,t}^{1-\theta} + (1-\alpha) P_{m,t}^{1-\theta}\right]^{1/(1-\theta)}$, where θ is the elasticity of substitution between these two types of intermediate goods and α is the share of intermediate domestic goods in final goods production (see Obstfeld and Rogoff (1995), Chari et al (2000, 2002) etc. for details).

From the above analysis, it is clear that the major determinant of ERPT to domestic prices is foreign firms' adjustments of markups, the degree of substitutability between domestic and foreign goods and the relative share of foreign goods in the domestic market. Dornbusch (1987) and Feenstra, Gagnon, and Knetter (1996) also argue along these lines. Adjustments of markups can also depend on the size and duration (temporary vs. permanent) of the exchange rate change. Firms absorb unexpected small and temporary exchange rate fluctuations in their markup and adjust prices in the local market only if the fluctuation is large (see Krugman (1987), Dixit (1989)). Froot and Klemperer (1989) also show that the size and direction of the pass-through depend on whether the exchange rate change is perceived to be temporary or permanent. Mann (1986) argues that high volatility in the exchange rate may make importers adjust profit margins instead of adjusting prices frequently, thus reducing pass-through. In a similar vein, Taylor (2000) argues that pass-through will be greater if the exchange rate and import price fluctuations show greater persistence since firms will adjust prices rather than markups in this case. Krugman (1987) found the presence of local currency pricing by analyzing the data on US import prices and US-German trade from 1980 to 1984. He found that 35 to 40 percent of the real exchange rate appreciation of the US Dollar had been absorbed by foreign exporters when raising their prices in the US market relative to other markets. Knetter (1993) also found evidence of local currency pricing by analyzing the industry level data for the USA (1973-87), UK (1974-87), Germany (1975-87) and Japan (1973-87).

Foreign exporters do not sell products directly to the consumer. There are costs of advertising, distributing and retail selling which must be paid in the local currency. Burstein, Neves and Rebelo (2003) estimate that such distribution costs represent more than 40 percent of the retail price in USA and create a natural wedge between retail prices in different countries. Goldberg and Campa (2010) quantify the responsiveness of consumer prices to exchange rate and import price changes via different channels for 21 industrialized countries. The major channels include the adjustment of the distribution margin to exchange rate changes and the extent of imported inputs used in different types of consumer goods. Across this 21 countries, the distribution margin on consumption goods ranges from 30 percent to 50 percent of the purchaser's price. Imported inputs account for 10 percent to 48 percent of the production cost of the tradable goods and 3 percent to 35 percent of the production cost of the non-tradable goods, attesting to the importance of these two channels. Using panel regression, they found that a nominal exchange rate depreciation of 1 percent results in a 0.35 percent decrease in the distribution margin. Also, in response to a one percent nominal exchange rate depreciation, the unweighted average of ERPT to consumer prices is 0.15 percent in the long run for these countries.

Almost all empirical studies show that the pass-through to import prices is incomplete, i.e., import prices increase proportionately less than the depreciation of the domestic currency. For example, Campa and Goldberg (2005), using single equation estimation and quarterly data for 23 OECD countries from 1975 to 2003, find that the unweighted average pass-through to import prices is 46 percent over shorter terms (1-quarter) and 64 percent over longer terms (1-year). They found that countries with lower exchange rate and inflation volatility have lower ERPT to import prices. Since ERPT to import prices is incomplete, it is likely that ERPT to consumer prices will also be incomplete and much smaller compared to import prices. The ERPT to CPI, 1-year after the shock (depreciation), in Choudhri et al. (2005) for non-US G-7 countries during 1979-2001 is in the range of 0.04 percent to 0.20 percent. An and Wang (2012) estimate ERPT to import, producer and consumer prices for nine OECD countries (7 of those countries are examined in this chapter) using the VAR model and by imposing sign restrictions on the impulse responses. Using monthly data for the period 1980 to 2007, they found that the average pass-through estimates for consumer prices are on average 0.09 and 0.08 at the 1-quarter and 1-year horizons, respectively.

The ERPT to CPI in some developing and/or emerging countries tends not to be very high. Ito and Sato (2008) found that the extent of ERPT to consumer prices is modest, ranging from as low as 0.03 percent in Malaysia to as high as 0.40 percent in Indonesia after one year of the shock for 5 east-Asian countries (Indonesia, Korea, Malaysia, the Philippines and Thailand). Kohlscheen (2010) analyzes the degree of ERPT to CPI for 8 emerging countries using a bivariate VAR with NEER and CPI. They found that the unweighted average response of CPI to 1 percent depreciation of NEER is 0.18 percent and 0.28 percent after 2-quarters and 1-year of the shock, respectively.

There seems to be enough evidence that ERPT tends to be lower under common currency and inflation targeting regime. There are several reasons to expect that, under a currency board or monetary union (with a common currency), ERPT to domestic prices should decline. First, as Frankel and Rose (2002) argue, under a common currency, trade flows among participating countries increase. Since more trade occurs in the same currency, they are less affected by relative price swings caused by exchange rate fluctuations. However, this hypothesis is debatable as Campa and Mínguez (2006) find no evidence of a shift in the trade patterns of Euro Area members after 1999. Second, Devereux et al. (2003) have hypothesised in the context of the Euro that a currency union will encourage local currency pricing and foreign exporters will adjust their profit margin when the exchange rate fluctuates. Campa and Mínguez (2006) do find evidence that the share of the euro as a currency of denomination for imports coming from third countries has increased since 1999. Faruqee (2006), analyzing the impulse response pattern from a VAR framework, also found a high degree of local currency pricing in import prices and producer currency pricing in export prices for the Euro area as a whole. Third, a monetary union will encourage monetary stability and will provide a low inflationary environment. As already discussed, the pass-through to prices is likely to be low in a low inflationary environment. Choudhri and Hakura (2006) examine whether the inflationary environment matters for ERPT to consumer prices. Using quarterly data for 71 countries from 1979 to 2000, they divide the countries into three inflation-categories (low, moderate and high) and use single equation estimation to estimate ERPT to consumer prices. For low, moderate and high inflation countries, the unweighted average ERPT to consumer prices are 0.16 percent, 0.35 percent and 0.56 percent, respectively, in the long run (5 years) in response to 1 percent depreciation of the nominal exchange rate. Slavov (2008) examines whether monetary integration reduces ERPT for a broad panel of 101 countries over the period 1976–2006 using single equation estimation method. He used a common currency dummy for 32 countries with 33 different episodes of common currency and found a strong reduction in ERPT in EMU member countries since the launch of the Euro. Mishkin and Schmidt-Hebbel (2007), employ panel VAR technique on a group of inflation targeting countries along with a group of nontargeters. They found that inflation targeting countries have a smaller inflation response to exchange rate shocks.

2.3. Data

We use quarterly data of the selected countries from 1974 to 2016. This time period is split into two group based on the regime change: inflation targeting or adoption of Euro. The starting point of inflation targeting for Australia, Canada, New Zealand, Sweden and UK are 1993Q1, 1991Q1, 1990Q1, 1993Q1 and 1992Q1, respectively. Finland, France, Germany, Italy, Netherlands and Spain adopted the Euro as their currency since 1999Q1. The only exception is USA among our selected countries, as USA neither targets inflation nor has joined any currency union. However, since the inception of Euro, there is a competition for the USD to sustain its global currency status. Based on this, we consider that the regime has changed in USA since 1999Q1.

All variables except the interest rate are expressed in logarithmic form. We use the Census X13 procedure for seasonal adjustment of the quarterly data for all variables except the interest rate (IR). IR used is the short-term interest rate, policy rate, discount rate or money market interest rate, depending upon data availability. CPI represents consumer prices. Nominal Effective Exchange Rate (NEER) is the trade-weighted average of the nominal exchange rates, defined as the amount of foreign currency required to buy one unit of the domestic currency. Hence, an increase in NEER implies an appreciation of the domestic currency. The details of the variables and sources of the data are provided in Appendix 2.1.

To decide on the order of integration of the variables, we employ two unit root tests: (1) Augmented Dickey-Fuller (ADF) test, and (2) Dickey-Fuller test with generalized least square de-trending (DF-GLS). In the DF-GLS test, the time series is transformed via a generalized least squares (GLS) regression before performing the ADF-type test. Both tests are performed with (1) a constant term and (2) a constant and linear time trend. Lag selection is based on the general-to-specific procedure. This procedure -- also known as the t-test selection criteria -- selects the number of lags, starting with a pre-specified maximum lags, for which the last included lag has a marginal significance level less than the pre-specified level of significance. We set the maximum lag at 9 and the significance level at 10 percent. Table 2.1 shows the summary of the results using the ADF and DF-GLS tests and Appendix 2.2 shows the details of the test results.

Before the regime change, CPI is $I(2)$ while IR and NEER are $I(1)$ for most countries. After the regime change, unit root behaviour is somewhat diverse. CPI is $I(1)$ for all countries except New Zealand. IR is trend-stationary (TS) and NEER is $I(0)$ for the countries that adopted the Euro. For the remaining countries, IR is either TS or $I(1)$ and NEER is either $I(1)$ or $I(0)$.

2.4. Causality Analysis

We start the causality analysis with the Granger causality test. We examine both unconditional and conditional causality between NEER and CPI, and between NEER and IR. The stationary form of each variable is used to form a 3-variable VAR model with CPI, IR and NEER. Granger causality results are sensitive to the number of lags included in the model. The traditional information criteria (e.g., Akaike, Schwarz, Hannan-Quinn, etc.) selects 1 or 2 lags for most cases. However, CPI can take a while to respond to exchange rate or interest rate changes. Similarly, there can be a delayed response of IR to NEER or CPI changes. To allow for that possibility, we restrict the number of lags between 4 and 6; and use Akaike information criterion (AIC) to choose the best fit within this lag range. Then we use the VAR stability test and serial correlation test to ensure that the model is stable and the residuals are not serially correlated. If the model does not pass these two tests, we choose the next best fit based on AIC in the specified range for lags. The summary of the results is presented in Table 2.2 and the details of the Granger causality test is presented in Table 2.3a and Table 2.3b.

Tables 2.3a and 2.3b show the p-value of the Granger causality test before the regime change and after the regime change, respectively. These tables show that the unconditional and conditional tests produce similar results. From the summary Table 2.2, we can see that the number of rejections for the null hypothesis that "NEER does not Granger cause CPI" and "NEER does not Granger cause IR" decreased considerably after the regime change. On the contrary, the number of rejections for the hypothesis that "IR does not Granger cause NEER" increased after the regime change. These results indicate that the central banks are using pro-active monetary policy throughout the period and even more so after the regime change. Also, after the regime change, NEER does not explain the movements of CPI in the countries that adopted the Euro. We consider this evidence as supporting local currency pricing for these countries.

Granger causality tests have two limitations: (1) it can examine causality only at horizon 1 and (2) it is possible that neither or both null hypothesis is rejected. To overcome these limitations, we now estimate the causality measure following Dufour and Taamouti (2010), which examines the strength of the direction of causality at multiple horizons. We use the 3-variable VAR specification where all variables are in stationary form and select the same number of lags used in the Granger causality test. We estimate conditional causality measures up to horizon 12 (for 3 years) and construct 95% bootstrap confidence intervals. We found that the test results are sensitive to the model specification, i.e., differencing of the variables and number of lags chosen. This restricts us from comparisons across countries as the order of integration of the variables and number of selected lags differ across countries. However, depending on the model specification, the value of the causality measure changes but the strength of the direction of causality does not change for a country. This is a considerable advantage over the standard Granger causality test where the test results can completely change depending on the chosen lags.

The results are presented graphically in Figures 2.1 to 2.4. We have kept the vertical scale the same in both figures to facilitate comparison within a country. The causality measure are represented by the solid line and 95% Bootstrap confidence band is represented by the dashed lines. Figures 2.1a-b present the causality measures between CPI and NEER before the regime change while Figures 2.2a-b present the causality measure between CPI and NEER after the regime change. Similarly, Figures 2.3a-b and 2.4a-b present the causality measures between IR and NEER before the regime change and after the regime change, respectively. The summary of the results is provided in Table 2.4. We observe in these figures that causality measures have higher values at shorter horizons (up to 3-4 quarters) and decrease considerably as the horizon increases. The lower bound of the confidence band becomes zero after 4 to 6 quarters for most cases. There are some instances where the strength of the direction of causality is similar. For example, after the regime change in Canada, the measure of causality from CPI to NEER is similar to that of NEER to CPI.

Table 2.4 also shows an interesting pattern. After the regime change, the direction of causality from NEER to CPI is stronger for inflation targeting countries while the direction of causality from CPI to NEER is stronger for the countries in a currency union. This implies that, for the countries in the currency union, CPI affects the exchange rate via the trade balance while exchange rate fluctuations do not affect the domestic price level. This can be interpreted as an example of the local currency pricing, increased trade among the countries in the currency union and/or success in keeping a low inflationary environment in the region. As in the Granger causality test, the direction of causality from IR to NEER is stronger both before and after the regime change. Again, after the regime change, the direction of causality from IR to NEER is stronger for countries in the currency union while the results are mixed for the inflation targeting countries. Our results imply that the central banks in the inflation targeting countries have to react by altering the interest rate to keep the price level at the desired level. On the contrary, the ECB does not need to change the interest rate to keep the price level stable and exchange rate shock is absorbed via exporters' PTM behaviour or by switching to local substitutes.

2.5. Pass-Through Behaviour

2.5.1. Econometric Methodology

In this section, we discuss the procedure used for estimating the VAR model and deriving the IRFs. In doing so, we will also discuss the difference between structural equations and reduced form equations, how to recover the structural shocks from the reduced form variance-covariance matrix, and also the significance of the ordering of the variables. Since we will be using the Cholesky decomposition of the residual covariance matrix, the ordering of these variables is crucial as it affects the IRFs.

A structural VAR of lag length *p,* in general form, can be written as:

$$AY_t = B_1 Y_{t-1} + B_2 Y_{t-2} + \cdots + B_p Y_{t-p} + u_t \tag{1}$$

where $E(uu') = \Sigma_u$. The structural shocks are mutually uncorrelated which implies that Σ_u is diagonal and shocks are equal to the number of variables. Furthermore, the diagonal elements can be normalized to 1 without loss of generality as long as the diagonal elements of A remain unrestricted. Now, pre-multiplying both side by A^{-1}, we get the reduced form VAR as:

$$Y_t = C_1 Y_{t-1} + C_2 Y_{t-2} + \cdots + C_p Y_{t-p} + \varepsilon_t \tag{2}$$

where $C_i = A^{-1} B_i$ for $i = 1, 2, \dots, p$; and $\varepsilon_t = A^{-1} u_t$; $E(\varepsilon\varepsilon') = \Sigma_\varepsilon = A^{-1}\Sigma_u (A^{-1})' = A^{-1}(A^{-1})'$ using $\Sigma_u = I$.

This VAR in equation (2), known as reduced form VAR, can be efficiently estimated by OLS even though ε_t may be contemporaneously correlated. Thus, the coefficient matrix C_i $(i = 1,2, \dots, p)$ and the reduced form covariance matrix Σ_ε can be estimated from (2). This information is sufficient for forecasting. However, for structural analysis, we need to recover the elements of matrix A. Cholesky decomposition can be used for that purpose. Using this decomposition, we can recover matrix A from Σ_ε and C_i with no prior knowledge of Σ_u.

Σ_ε is a symmetric matrix where the diagonal elements show variances of the shocks and the off-diagonal elements show the covariance of the shocks. It contains $(n^2 + n)/2$ distinct elements, so that, the maximum number of elements that we can uniquely identify in matrix A is also $(n^2 + n)/2$. We have to set the remaining $(n^2 - n)/2$ elements to zero. A lower triangular matrix is sufficient for that by defining P such that $PP' = \Sigma_\varepsilon$. It follows immediately that $A^{-1} = P$. In this decomposition, the ordering of the endogenous variables becomes important as it defines which contemporaneous effect is assumed to be zero. Since the Cholesky decomposition imposes restriction on contemporaneous effects, the estimated structural shocks may differ from actual structural shocks. Hence, in empirical literature, different orderings of the variables are often examined to verify the robustness of the results.

With an ordering of the variables $Y = (IR, NEER, CPI)'$, the relationship between the reduced form VAR residuals and structural VAR residuals can be written as:

$$\begin{bmatrix} \varepsilon_t^{IR} \\ \varepsilon_t^{NEER} \\ \varepsilon_t^{CPI} \end{bmatrix} = \begin{bmatrix} a_{11} & 0 & 0 \\ a_{21} & a_{22} & 0 \\ a_{31} & a_{32} & a_{33} \end{bmatrix} \begin{bmatrix} u_t^{IR} \\ u_t^{NEER} \\ u_t^{CPI} \end{bmatrix} \tag{3}$$

To sum up, we are estimating the reduced form equations in (2) and recovering the structural shocks using (3).

2.5.2. Ordering of the Variables

From the above discussion, it is evident that our estimating equations can vary with the ordering of the variables. However, there is no unique way to determine the ordering of the variables, so that the common practice in statistical analysis is to use the Granger causality test for selecting the ordering of the variables or use prior information from the economic literature to set $(n^2 - n)/2$ contemporaneous effects to zero.

Mihailov (2009) performed the Granger causality test for the variables included in his study for the USA, Germany and Japan. He did not find a unique ordering based on the Granger causality test for each country. Based on the test results, he used four 'most likely' orderings for estimation and presented a range for each pass-through coefficient. The variables included in his analysis are: CPI, import price index (MPI), export price index (XPI), NEER and monetary policy variable (M1); with the most likely ordering: (1) $Y = (M1, NEER, MPI, XPI, CPI)'$ (2) $NEER \to M1 \to MPI \to XPI \to CPI$; (3) $CPI \to M1 \to NEER \to MPI \to XPI$; and (4) $CPI \to NEER \to M1 \to MPI \to XPI$.

In the absence of conclusive results from the Granger causality tests, one has to rely on other empirical studies and some assumptions that best suit the purpose of the study. Along these lines, a number of studies (e.g., McCarthy (2000, 2007), Hahn (2003), Faruqee (2006) etc.) have used pricing along the distribution chain (i.e., transmission of exchange rate shock to consumer prices via import prices and producer prices) to identify the extent of the exchange rate shocks to import prices, producer prices and consumer prices simultaneously. These studies, in general, include three types of variables: domestic aggregate prices, aggregate demand and supply side variables; and monetary policy variables along with the nominal exchange rate. Domestic prices include import prices, export prices, producer prices (PPI) and consumer prices. Demand and supply shocks are represented by wages, output (or output gap) and oil prices. Monetary variables include the central bank's target interest rate and the growth of money supply (with money defined as M1 or M2).

McCarthy (2000, 2007) uses 8 variables with the ordering of the endogenous variables as: $Y = (Oil, gap, NEER, MPI, PPI, CPI, IR, M2)'$. This study does not provide any specific reason for the choice of this ordering. In Hahn (2003), the ordering of the endogenous variables is: $Y = (Oil, IR, gap, NEER, MPI, PPI, CPI)'$. Oil price is the most exogenous variable and ordered first. Due to the lagged availability of GDP data, it assumed that monetary shocks affect GDP (or output gap) contemporaneously and not vice versa. Monetary policy shocks also have a contemporaneous effect on the exchange rate. The output gap is ordered above the exchange rate to allow for the contemporaneous effect of demand shock on the exchange rate. The last three variables are ordered along the chain of production. Faruqee (2006) includes wage and export prices in the endogenous variables with the ordering $Y = (NEER, wage, MPI, XPI, PPI, CPI)'$. He orders the exchange rate first on the basis of the assumption that exchange rate innovations are caused by exogenous asset market disturbances. Next in ordering is wages, which is a major determinant of the cost of production. This is followed by the pricing chain of production. Ito and Sato (2008) estimate ERPT to CPI for East Asian countries and use the ordering: $Y = (Oil, gap, M2, NEER, CPI)'$. Oil price and the output gap represent supply and demand shocks, respectively. The monetary policy variable is ordered ahead of the exchange rate since monetary authorities tend to puts more emphasis on domestic targets (e.g., inflation) and the exchange rate is taken to fluctuate in response to monetary policy. All these studies perform robustness checks with different orderings of the variables.

In our case, we have a three-variable VAR and there are $3! = 6$ possible orderings among these three variables. As discussed earlier in this section, in VAR, each variable is a function of past values of all variables and current values of other variables. We need to impose restrictions only on the contemporaneous relations. For our selected countries, we did not get a unique ordering based on Granger causality tests. Hence, using information from the prior literature to derive the ordering of the variables, we set: $Y = (IR, NEER, CPI)'$. We assume that central banks have a pre-determined monetary target and that monetary policy contemporaneously affects NEER and CPI. The Granger causality test and causality measure also suggest that the causality runs from IR to CPI. Information on CPI becomes available with a delay and often prices are sticky. Hence, CPI is ordered after NEER. This ordering is similar to, considering the variables in this model, that in Hahn (2003), Ito and Sato (2008) and one of the possible ordering of Mihailov (2009). We also examine two other orderings to verify the robustness of our results. These are: $Y = (IR, CPI, NEER)'$ and $Y = (CPI, NEER, IR)'$.

2.5.3. Results

A VAR model can be estimated in three different forms: (1) estimating the model in levels, ignoring the nonstationarity of variables; (2) differencing any nonstationary series before estimating the model; and (3) using the cointegrated model if all the variables are $I(1)$. In our case, for each country, all the variables do not have the same order of integration; so that we do not look for cointegration in the model. Further, a temporary shock in the exchange rate conceptually will not have a long-term effect on domestic prices; so that, even if all variables have the same order of integration, the cointegration framework may not be appropriate in this case. If the model is estimated in the level form, ignoring nonstationarity, the parameters are estimated consistently. Furthermore, even if the true model is a VAR in differences, the differenced model produces no gain in asymptotic efficiency (for details, see Hamilton (1994), pp 651-653). Hence, we estimate our model in level form.

As mentioned earlier, we estimate a 3-variable VAR in this chapter. We have not included any demand or supply side variable here. Hence, there can be omitted variable bias in the estimates. However, we have a small sample period in each regime and we want to use lags of at least one year to capture lagged effects. For each country, we include a constant and the chosen number of lags, following the same lag selection criteria discussed in the causality analysis section. The estimated models are stable (i.e., the characteristic roots lie outside the unit circle), and hence, have a vector moving average (VMA) representation. We derive the impulse response of CPI and IR to NEER shocks from this VMA representation. Impulse responses are estimated over a horizon of 3 years (12 quarters) and the responses are standardized to correspond to a 1 percent shock in the variable of interest.

Impulse responses[3] trace out the response of current and future values of a variable to a one-unit (or to a one-standard deviation) increase in the current value of one of the VAR errors

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[3] Given our structural model: $A(L)Y_t = u_t$, with $E(u_t) = 0$ and $E(u_t u_t') = \Sigma_u$. With Cholesky decomposition, the covariance stationary VAR has the following $VMA(\infty)$ representation: $Y_t = \mu + \Phi_0 \eta_t + \Phi_1 \eta_{t-1} + \cdots$.
The impulse response to the orthogonal shock η_{jt} is $\frac{\partial y_{i,t+h}}{\partial \eta_{j,t}} = \frac{\partial y_{i,t}}{\partial \eta_{j,t-h}} = \phi_{ij}^h$, $i, j = 1, \dots, n$; $h > 0$, where ϕ_{ij}^h is the (i, j)-th element of Φ_h.

under the assumption that this error returns to zero in subsequent periods and that all other errors are equal to zero. We estimate the impulse responses of CPI and IR over the twelve-quarter horizon for a Cholesky one standard deviation increase in the exchange rate. Then, the results are standardized to represent impulse responses to a 1 percent shock to the exchange rate. Figure 2.5a-b displays the responses of the CPI to a 1 percent increase (appreciation of the domestic currency) in the exchange rate, before the regime change in the left panel and after the regime change in the right panel. The solid line in each graph is the estimated response while two sets of dashed lines denote 16% - 84% lower and upper bounds of the confidence band (this is the robust version of the 1 standard error confidence band) and 2.5% - 97.5% lower and upper bounds of the confidence band (this is the robust version of the 2 standard error confidence band), obtained through Monte Carlo integration. Figure 2.6a-b presents the response of IR to a 1 percent appreciation of the exchange rate.

Based on Figures 2.5a-b and 2.6a-b, very few responses are statistically different from zero since the 2 standard error confidence interval includes zero. Hence, it will not provide scope for comparisons across countries. For the remainder of this section, we will focus on the 1 standard error confidence interval. We draw our conclusions based on up to 4 to 6 horizons unless there is a clear evidence of a more delayed response. Conclusions drawn from the responses in these figures are summarized in Table 2.5. Before the regime change, the responses of CPI to the NEER shock are statistically significant for Australia, Canada, Germany, Italy, Netherlands, New Zealand and USA, and have the expected negative signs. For the remaining five countries, the responses are statistically insignificant and show wide variations in terms of the sign of the responses. The responses of IR to a NEER shock are statistically significant and are positive for Australia, Finland, Germany, Sweden and UK while the responses are negative for Netherlands, New Zealand and USA. If an exchange rate appreciation creates a current account deficit, countries may need to raise the interest rate to achieve a surplus in the capital account and restore the balance of payments. In this regard, the responses for Netherlands, New Zealand and USA are unexpected, though plausible in the context of sterilizing monetary policy intervention.

After the regime change, the differences in the responses of inflation targeting countries and countries using the Euro become more visible. For the countries with the Euro, the response of CPI to a NEER shock is statistically insignificant -- and has the unexpected positive sign. On the other hand, for the inflation targeting countries, the response of CPI to a NEER appreciation is statistically significant with the expected sign for 4 countries. The responses of IR to a NEER appreciation are positive for both sets of countries, though the responses are not statistically significant for three inflation targeting countries.

Since we divided our sample period into two groups and our sample periods are different from those in previous studies, we do not expect our results to be completely identical with those of other studies. However, some comparison with the findings of other studies may be illustrative. The sample period of McCarthy (2007) is comparable to our sample period before the regime change. Their graphs of the impulse responses show a similar pattern for the common countries discussed in this chapter. Mihailov (2009) reported that the ERPT to CPI after one year of the shock (depreciation) for USA and Germany is -0.01% to 0.00% and 0.09% to 0.13%, respectively, for the period 1979-2002, depending on the ordering of the variables. Our corresponding pass-through estimates are somewhat higher at this horizon for both countries. Choudhri and Hakura (2006), using quarterly data from 1979 to 2000, use single equation method to estimate the ERPT to consumer prices. For the 12 countries of this chapter, they found the pass-through to lie in the range of 0.02 percent (in UK and USA) to 0.27 percent (in New Zealand) at the 1-year horizon. For Finland, the pass-through elasticity has the wrong sign (-0.02 percent). In our analysis, the pass-through estimates are higher and more countries (including Finland) have an unexpected response at the 1-year horizon. For the countries with expected sign of responses, before the regime change, the pass-through elasticities lie between 0.04 percent to 0.36 percent at 1-year horizon. After the regime change, again for the countries with expected sign of responses, the pass-through elasticity has decreased considerably and the decrease has not occurred proportionally to all countries. For example, the pass-through elasticity was highest in New Zealand before the regime change (similar to the Choudhri and Hakura study) while it becomes insignificant after the regime change. Based on 1 standard error confidence interval, in An and Wang (2012), the pass-through estimates are statistically significant for Canada, Spain and Sweden; and statistically insignificant for Finland, Italy, UK and USA. For these seven countries, our results are similar to those in the An and Wang study only for Canada, Finland and UK.

We also examined the responses of CPI and IR to NEER shock with different orderings of the variables, as mentioned in section 2.5.2. Although the pass-through coefficients change to some extent, they do not alter the basic results from the benchmark ordering. These results are not reported here in order to conserve space.

2.6. Conclusions

This chapter examines causality between (1) the exchange rate and consumer prices and (2) the exchange rate and the short-term interest rate; and estimates the response of consumer prices and the interest rate to exchange rate shocks for 12 OECD countries for the period 1974 to 2016. Based on the regime change (inflation targeting or adoption of the Euro as the currency), we divide the time period into two groups and measure the causality and pass-through behaviour. This allows us to compare whether the response and causality changed for each country after the regime change and whether the Euro countries behaved differently to exchange rate shocks compared with the inflation targeting countries.

In this chapter, we used a forecast based causality measure proposed by Dufour and Taamouti (2010) to measure the strength of causality at multiple horizons. To the best of our knowledge, this is the first paper to use this method to examine causality between (1) the exchange rate and the consumer price level, and (2) the exchange rate and the interest rate. Using this method, we have found that the direction of causality from CPI to the exchange rate has more strength in the countries that adopted the Euro while for inflation targeting countries the direction of causality from the exchange rate to CPI has more strength. The strength of the direction of causality changed for some of the countries after both types of regime changes.

We also estimate the response of consumer prices and interest rates to exchange rate shocks, using a recursive vector autoregression (VAR) model. This adds another contribution to the literature as it provides additional evidence to the debate on whether the pass-through is decreasing across countries. A comparison of the two sets of countries provides further insights on the differences in pass-through behaviour among the inflation targeting countries and the Euro countries. We found that the ERPT to CPI is not statistically different from zero for the countries that adopted the Euro as their currency. Before the adoption of the Euro, the response was statistically different from zero for some of the countries. On the other hand, the responses of the CPI to the exchange rate changes do not exhibit any distinctive pattern for the inflation targeting countries after the regime change. Interest rates increased in response to exchange rate appreciations for all countries after the regime change while they showed considerable cross-country variation before the regime change.

To sum up, we find that the causality and pass-through behaviour shows a distinctive pattern between our two sets of countries after the regime change while the results were somewhat mixed before the regime change. For countries that adopted the Euro, exchange rate fluctuations do not explain the movements in the interest rate and consumer prices in the statistical causality sense. Rather the direction of causality from IR to NEER and from CPI to NEER is stronger compared to their respective counterparts. Also, the ERPT to consumer prices is not statistically different from zero in these countries. For the inflation targeting countries, the direction of causality from NEER to CPI is stronger compared to the direction of causality from CPI to NEER. Further, the ERPT to CPI is statistically significant with the expected sign of the responses in four out of the six countries. In general, the ERPT to consumer prices has decreased.

Table 2.1: Summary of Unit Root Tests

	Before Regime Change			After Regime Change		
	CPI	IR	NEER	CPI	IR	NEER
AUS	I(2)	I(1)	I(1)	I(1)	I(1)	I(1)
CAN	I(2)	I(1)	I(1)	I(1)	I(1)	I(1)
FIN	I(2)	I(1)	I(1)	I(1)	TS	I(0)
FRA	I(2)	I(1)	I(1)	I(1)	TS	I(0)
GER	I(2)	I(0)	I(1)	I(1)	TS	I(0)
ITA	I(2)	I(1)	I(1)	I(1)	TS	I(0)
NET	I(2)	I(1)	I(1)	I(1)	TS	I(0)
NZL	I(1)	I(1)	I(1)	I(2)	I(1)	I(0)
SPA	I(2)	I(1)	I(1)	I(1)	TS	I(0)
SWE	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
UK	I(2)	I(0)	I(1)	I(1)	TS	I(1)
USA	I(2)	TS	I(1)	I(1)	TS	I(1)

Table 2.2: Summary of Granger Causality Test (→ stands for Granger causes)

	Before Regime Change	After Regime Change
NEER→CPI	Canada, **Finland**, France, Germany, **Netherlands**, **New Zealand**, UK	Sweden, **UK**, USA
CPI→NEER	**Finland**, **Netherlands**, **New Zealand**, Spain	Canada, France, Italy, Spain, **UK**
NEER→IR	New Zealand, Spain, Sweden, UK, **USA**	**UK**
IR→NEER	Australia, Canada, France, Germany, Italy, Netherlands, **USA**	Australia, Finland, France, Germany, Italy, Netherlands, Spain, Sweden, **UK**

Note: Countries showing bi-directional causality are in bold font.

Table 2.3a: Unconditional and Conditional Granger Causality Test (Before Regime Change)
(p-values are reported in the table)

	$NEER \nrightarrow CPI$	$CPI \nrightarrow NEER$	$NEER \nrightarrow IR$	$IR \nrightarrow NEER$	$NEER \nrightarrow CPI \vert IR$	$CPI \nrightarrow NEER \vert IR$	$NEER \nrightarrow IR \vert CPI$	$IR \nrightarrow NEER \vert CPI$
AUS	0.83 (5)	0.92 (5)	0.46 (5)	**0.04** (5)	0.54 (6)	0.55 (6)	0.33 (6)	**0.01** (6)
CAN	0.21 (5)	0.95 (5)	0.87 (4)	**0.10** (4)	**0.05** (4)	0.71 (4)	0.88 (4)	**0.08** (4)
FIN	**0.04** (4)	**0.07** (4)	0.11 (4)	0.33 (4)	**0.05** (4)	**0.04** (4)	0.25 (4)	0.19 (4)
FRA	**0.06** (5)	0.62 (5)	0.76 (5)	0.28 (5)	**0.07** (5)	0.24 (5)	0.67 (5)	**0.08** (5)
GER	0.13 (4)	0.88 (4)	0.75 (5)	**0.03** (5)	**0.06** (5)	0.78 (5)	0.93 (5)	**0.03** (5)
ITA	0.75 (5)	0.60 (5)	0.98 (5)	0.12 (5)	0.79 (4)	0.28 (4)	0.96 (4)	**0.05** (4)
NET	0.14 (5)	**0.02** (5)	0.13 (5)	**0.00** (5)	**0.00** (5)	**0.03** (5)	0.29 (5)	**0.00** (5)
NZL	**0.00** (6)	**0.08** (6)	**0.05** (5)	0.36 (5)	**0.00** (6)	**0.05** (6)	0.15 (6)	0.29 (6)
SPA	0.60 (5)	**0.01** (5)	**0.02** (5)	0.82 (5)	0.59 (5)	**0.01** (5)	**0.04** (5)	0.75 (5)
SWE	0.74 (4)	0.67 (4)	**0.04** (4)	0.48 (4)	0.97 (4)	0.80 (4)	**0.05** (4)	0.61 (4)
UK	0.82 (6)	0.13 (6)	**0.00** (5)	**0.05** (5)	**0.05** (5)	0.49 (5)	**0.01** (5)	0.24 (5)
USA	0.21 (4)	0.29 (4)	**0.04** (5)	**0.06** (5)	0.28 (5)	0.36 (5)	**0.09** (5)	**0.08** (5)

Table 2.3b: Unconditional and Conditional Granger Causality Test (After Regime Change)
(p-values are reported in the table)

	$NEER \nrightarrow CPI$	$CPI \nrightarrow NEER$	$NEER \nrightarrow IR$	$IR \nrightarrow NEER$	$NEER \nrightarrow CPI\|IR$	$CPI \nrightarrow NEER\|IR$	$NEER \nrightarrow IR\|CPI$	$IR \nrightarrow NEER\|CPI$
AUS	0.98 (4)	0.29 (4)	0.87 (4)	**0.08** (4)	0.36 (4)	0.82 (4)	0.75 (4)	0.33 (4)
CAN	0.17 (4)	0.11 (4)	0.59 (4)	0.96 (4)	0.21 (4)	**0.09** (4)	0.44 (4)	0.86 (4)
FIN	0.18 (5)	0.14 (5)	0.29 (6)	**0.02** (6)	0.26 (6)	0.21 (6)	0.64 (6)	**0.03** (6)
FRA	0.48 (4)	**0.09** (4)	0.14 (6)	**0.01** (6)	0.83 (6)	0.18 (6)	0.12 (6)	**0.02** (6)
GER	0.98 (4)	0.40 (4)	0.21 (6)	**0.02** (6)	0.99 (6)	0.19 (6)	0.23 (6)	**0.01** (6)
ITA	0.41 (5)	**0.05** (5)	0.18 (6)	**0.01** (6)	0.73 (6)	**0.01** (6)	0.27 (6)	**0.00** (6)
NET	0.16 (4)	0.99 (4)	0.16 (6)	**0.01** (6)	0.59 (6)	0.27 (6)	0.19 (6)	**0.01** (6)
NZL	0.30 (6)	0.36 (6)	0.15 (6)	0.82 (6)	0.38 (6)	0.43 (6)	0.15 (6)	0.73 (6)
SPA	0.12 (5)	**0.07** (5)	0.11 (6)	**0.04** (6)	0.41 (6)	0.13 (6)	0.21 (6)	**0.05** (6)
SWE	**0.05** (4)	0.68 (4)	0.11 (4)	**0.07** (4)	0.28 (4)	0.27 (4)	0.40 (4)	**0.01** (4)
UK	0.32 (6)	**0.08** (6)	**0.09** (6)	**0.01** (6)	**0.05** (6)	0.42 (6)	0.34 (6)	**0.08** (6)
USA	0.18 (4)	0.21 (4)	0.88 (5)	0.17 (5)	**0.09** (5)	0.43 (5)	0.71 (5)	0.24 (5)

Table 2.4: Summary of Causality Measure (Based on Multiple Horizons)

	Before Regime Change	After Regime Change
NEER→CPI	(Australia), Canada, France, Germany, New Zealand, UK, (USA)	(Australia), (Canada)[a], (New Zealand), Sweden, (UK), USA
CPI→NEER	(Finland), Italy, Netherlands, Spain, (Sweden)	(Finland), France, Germany, Italy, (Netherlands), Spain
NEER→IR	New Zealand, Spain, Sweden, UK	(Canada), New Zealand, Sweden[a], UK
IR→NEER	Australia, Canada, (Finland), France, Germany, Italy, Netherlands, (USA)	Australia, Finland, France, Germany, Italy, Netherlands, Spain, USA

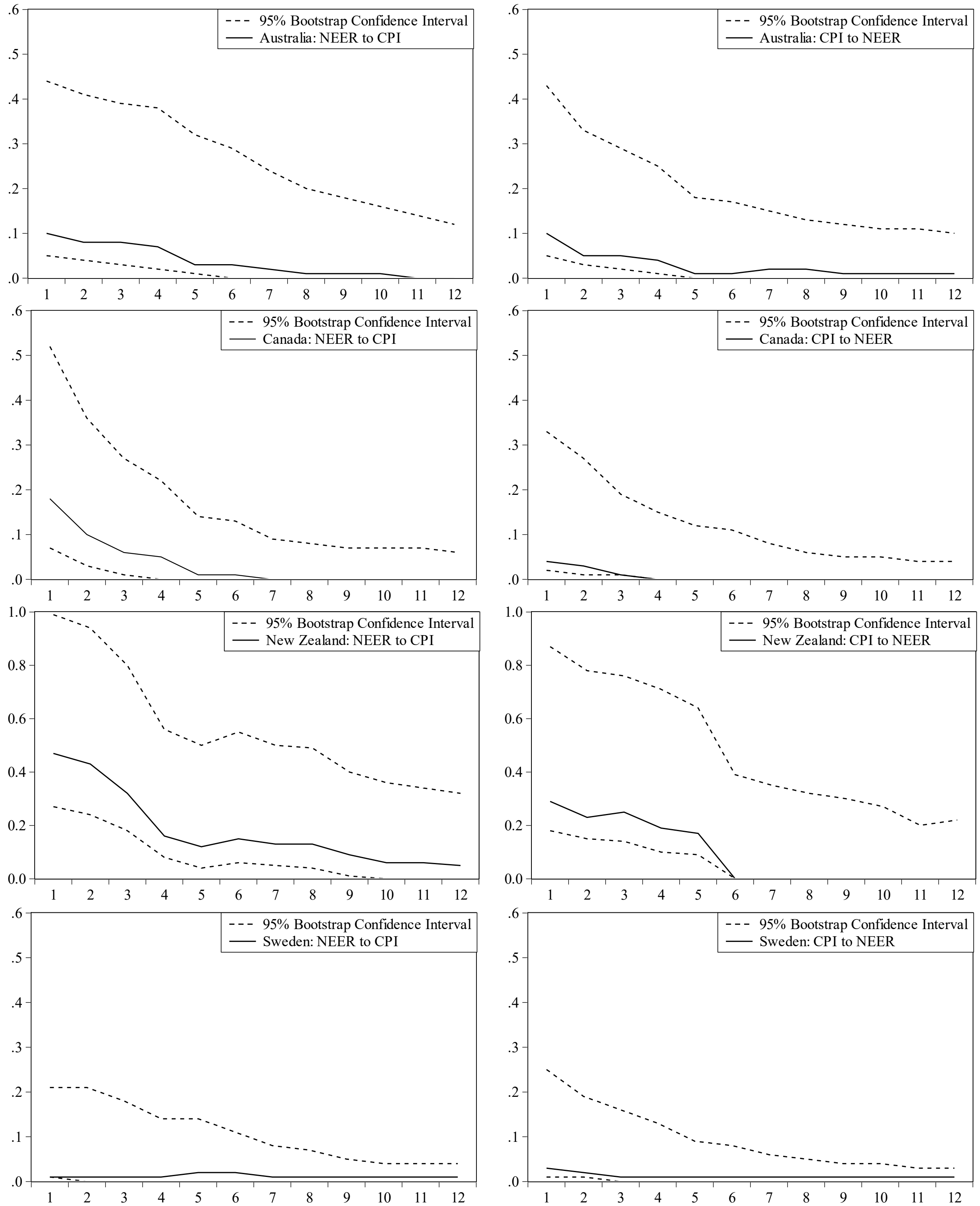
Note: Countries in parentheses means the difference in causal measure is very little. [a] Causality measure is higher after the first horizon

Table 2.5: Response of CPI and IR to 1% Appreciation of NEER

	Before Regime Change		After Regime Change	
	Response of CPI	Response of IR	Response of CPI	Response of IR
AUS	Significant, Negative	Significant, Positive	Significant, Negative	Insignificant, Positive
CAN	Significant, Negative	Insignificant, Negative	Significant, Positive	Insignificant, Positive
FIN	Insignificant, Positive	Significant, Positive	Insignificant, Positive	Significant, Positive
FRA	Insignificant, Positive	Insignificant, Negative	Insignificant, Positive	Significant, Positive
GER	Significant, Negative	Significant, Positive*	Insignificant, Positive	Significant, Positive
ITA	Significant, Negative	Insignificant, Positive	Insignificant, Positive	Significant, Positive
NET	Significant, Negative	Significant, Negative	Insignificant, Negative	Significant, Positive
NZL	Significant, Negative	Significant, Negative	Insignificant, Negative	Significant, Positive
SPA	Insignificant, Negative	Insignificant, Negative	Insignificant, Positive	Significant, Positive
SWE	Insignificant, Positive	Significant, Positive	Significant, Negative*	Insignificant, Positive
UK	Insignificant, Positive	Significant, Positive	Significant, Negative	Significant, Positive
USA	Significant, Negative	Significant, Negative	Significant, Negative	Significant, Positive

* *Delayed Response*

Figure 2.1a: Causality Measures between CPI and NEER (Before Regime Change)



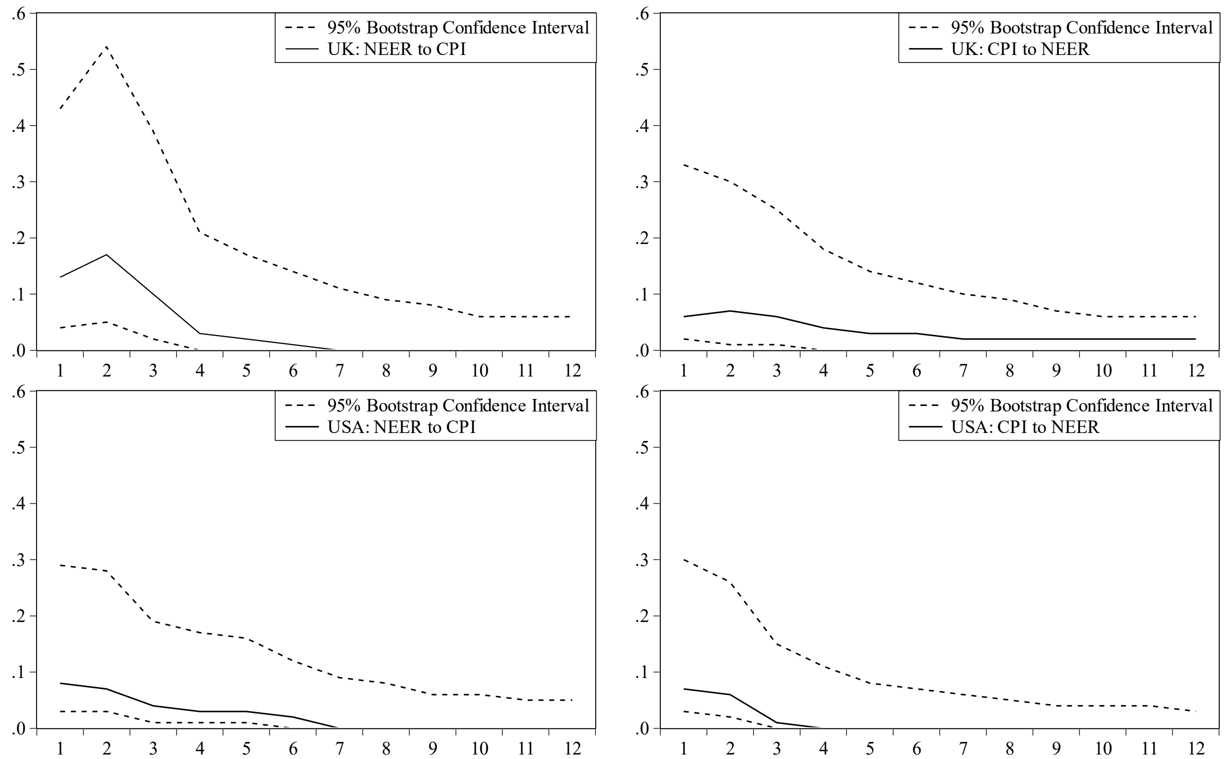
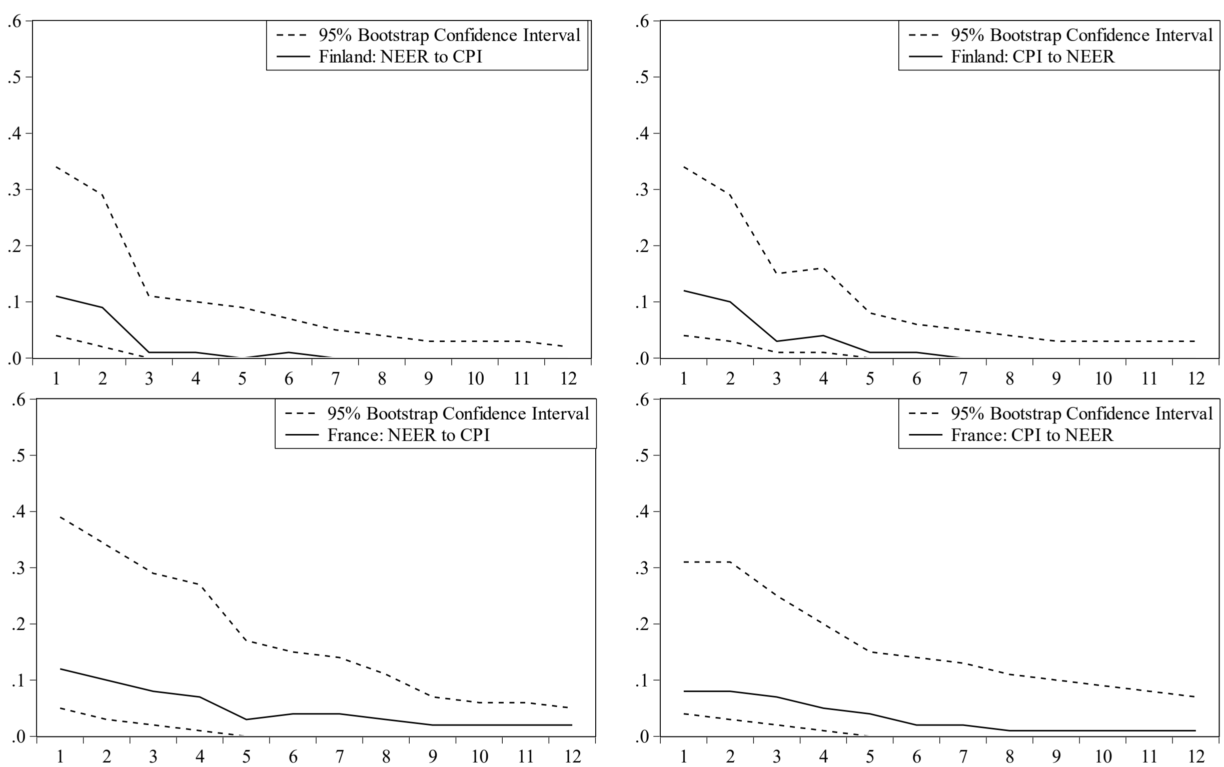


Figure 2.1b: Causality Measures between CPI and NEER (Before Regime Change)



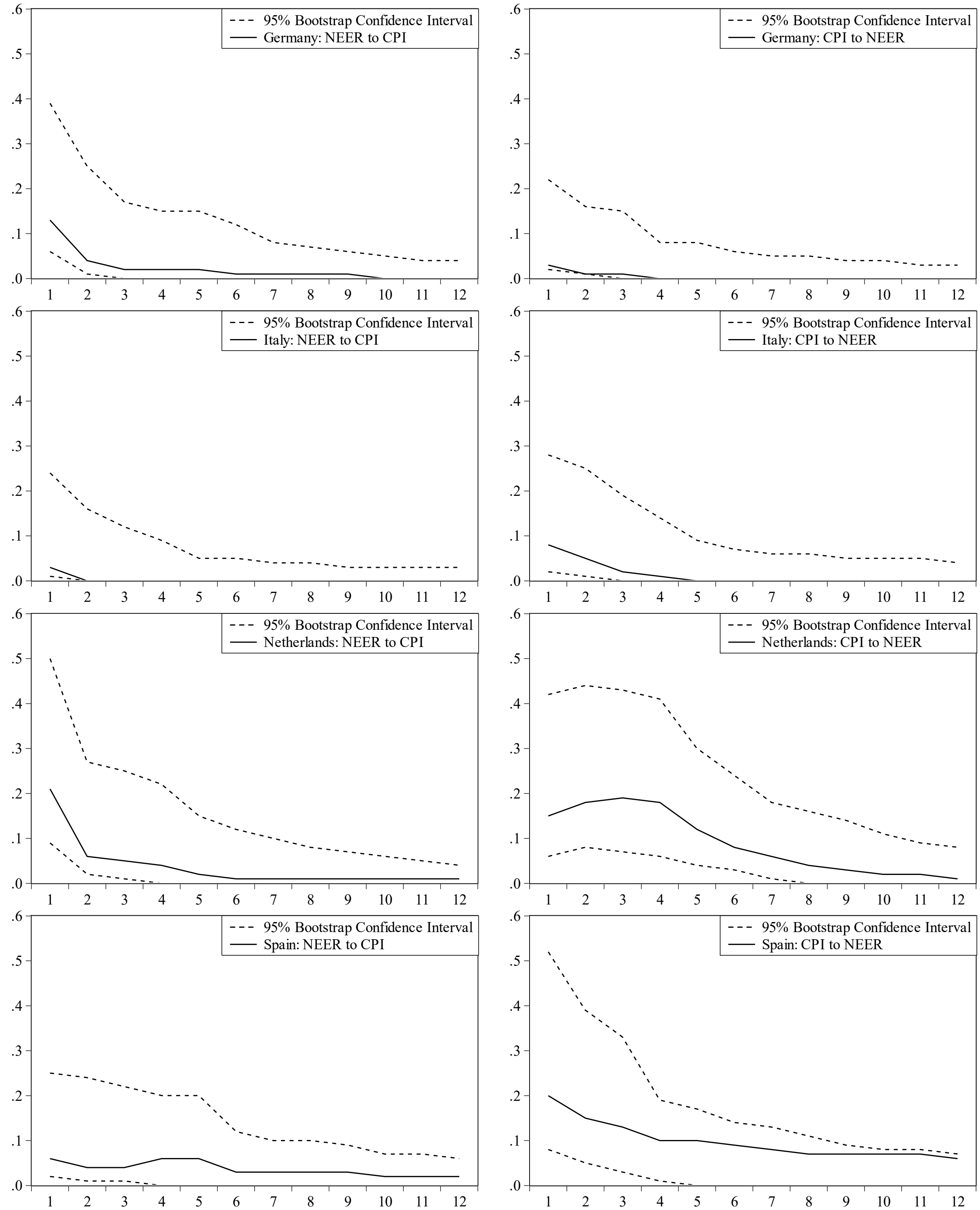
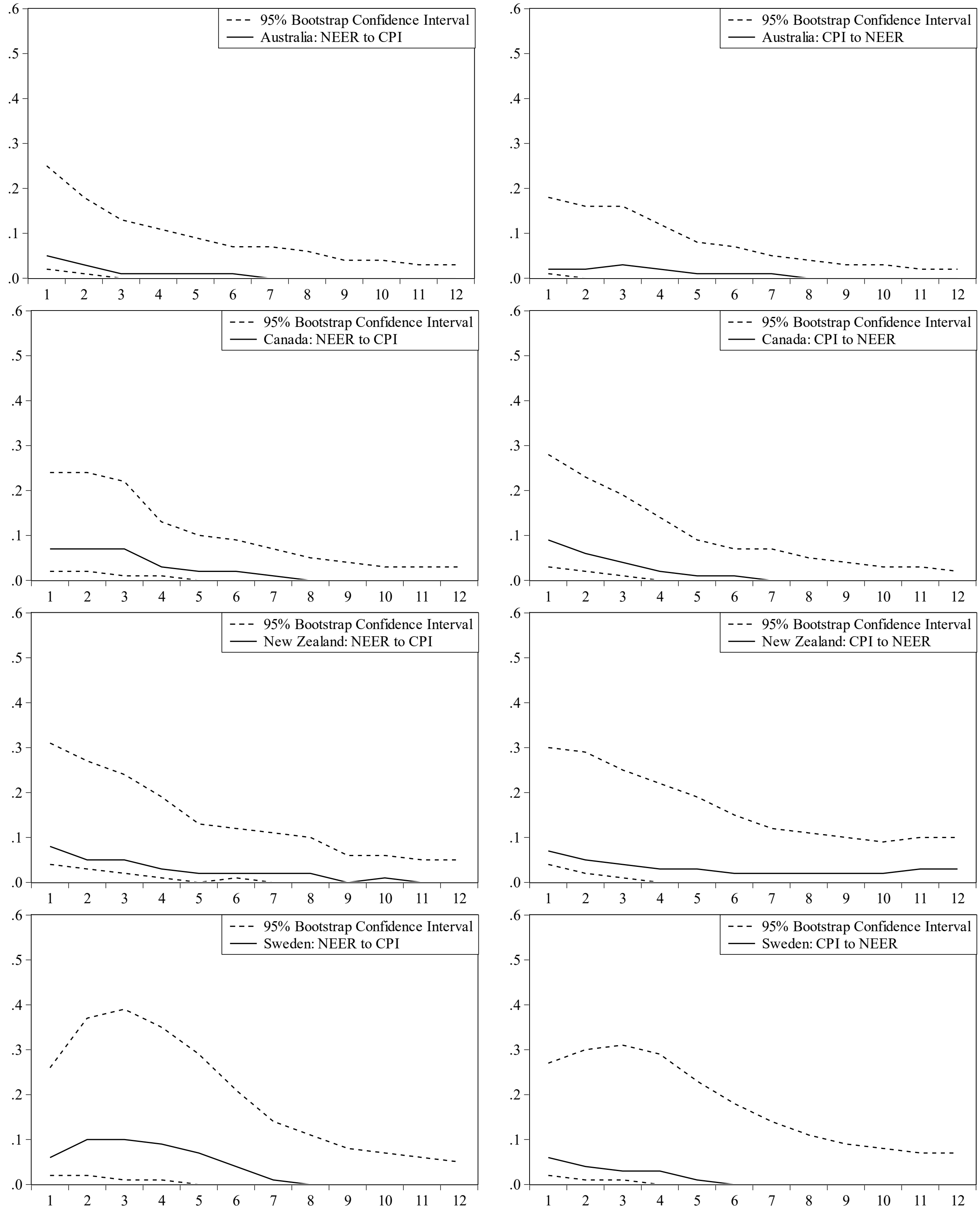

95% Bootstrap Confidence Interval
Germany: NEER to CPI
95% Bootstrap Confidence Interval
Germany: CPI to NEER
95% Bootstrap Confidence Interval
Italy: NEER to CPI
95% Bootstrap Confidence Interval
Italy: CPI to NEER
95% Bootstrap Confidence Interval
Netherlands: NEER to CPI
95% Bootstrap Confidence Interval
Netherlands: CPI to NEER
95% Bootstrap Confidence Interval
Spain: NEER to CPI
95% Bootstrap Confidence Interval
Spain: CPI to NEER
.0
.1
.2
.3
.4
.5
.6
1 2 3 4 5 6 7 8 9 10 11 12

Figure 2.2a: Causality Measures between CPI and NEER (After Regime Change)



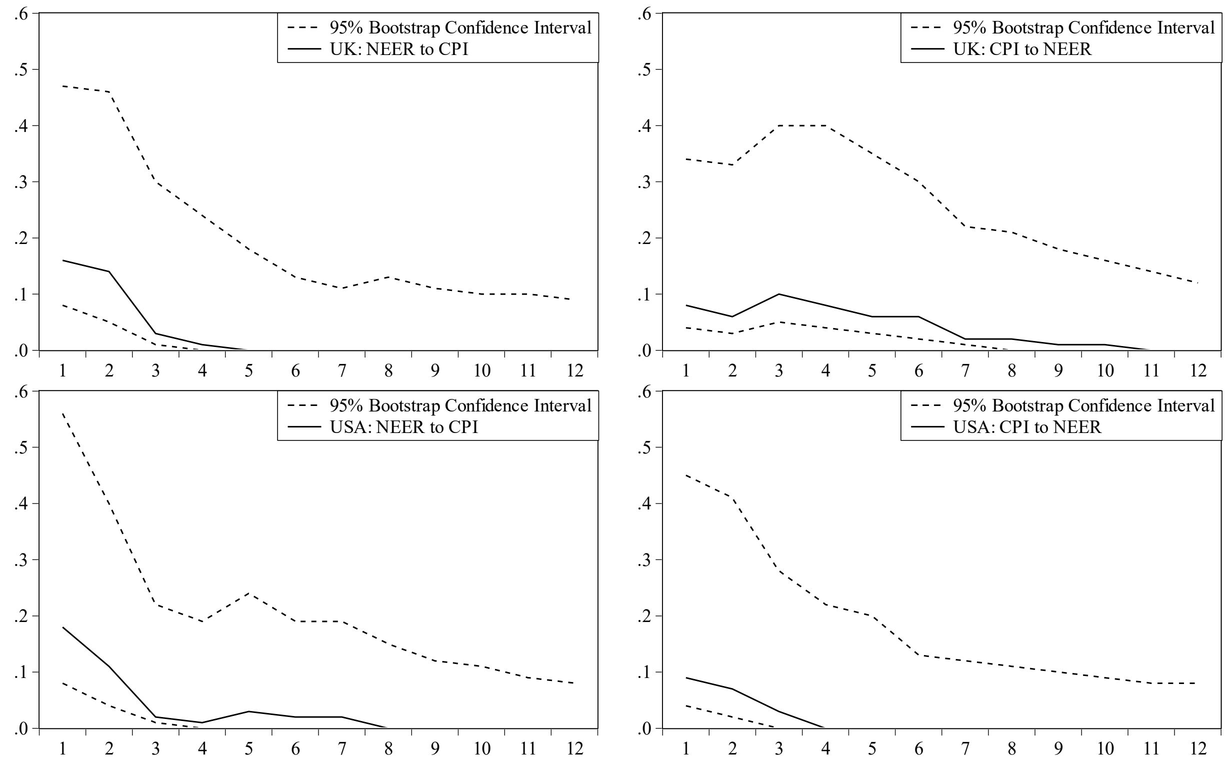
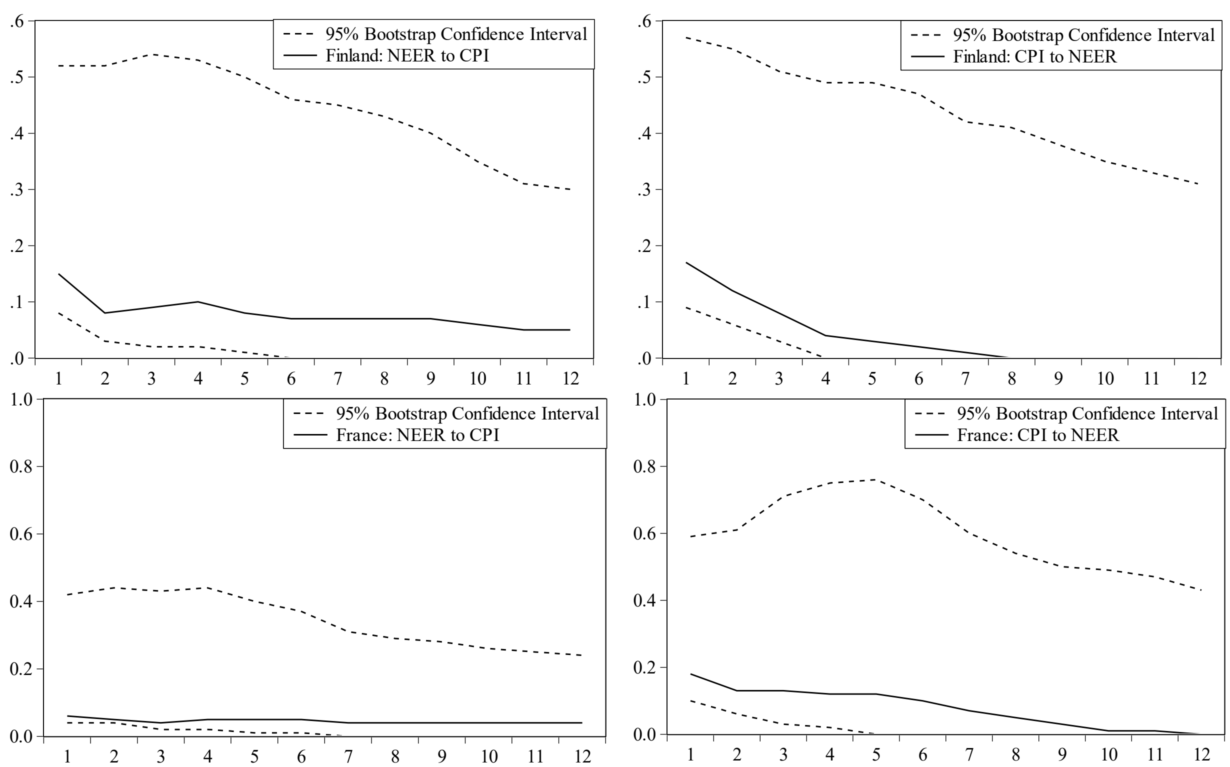


Figure 2.2b: Causality Measures between CPI and NEER (After Regime Change)



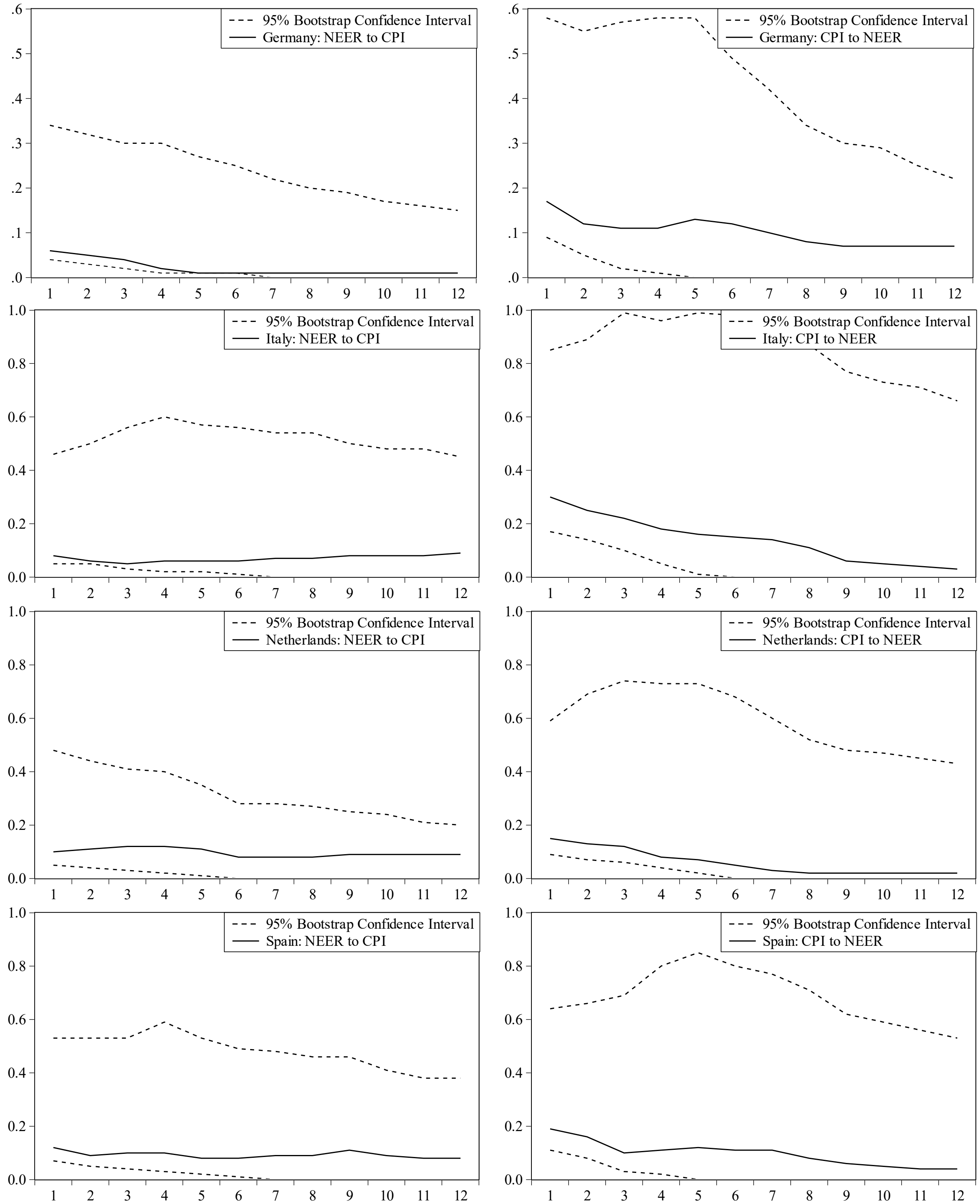
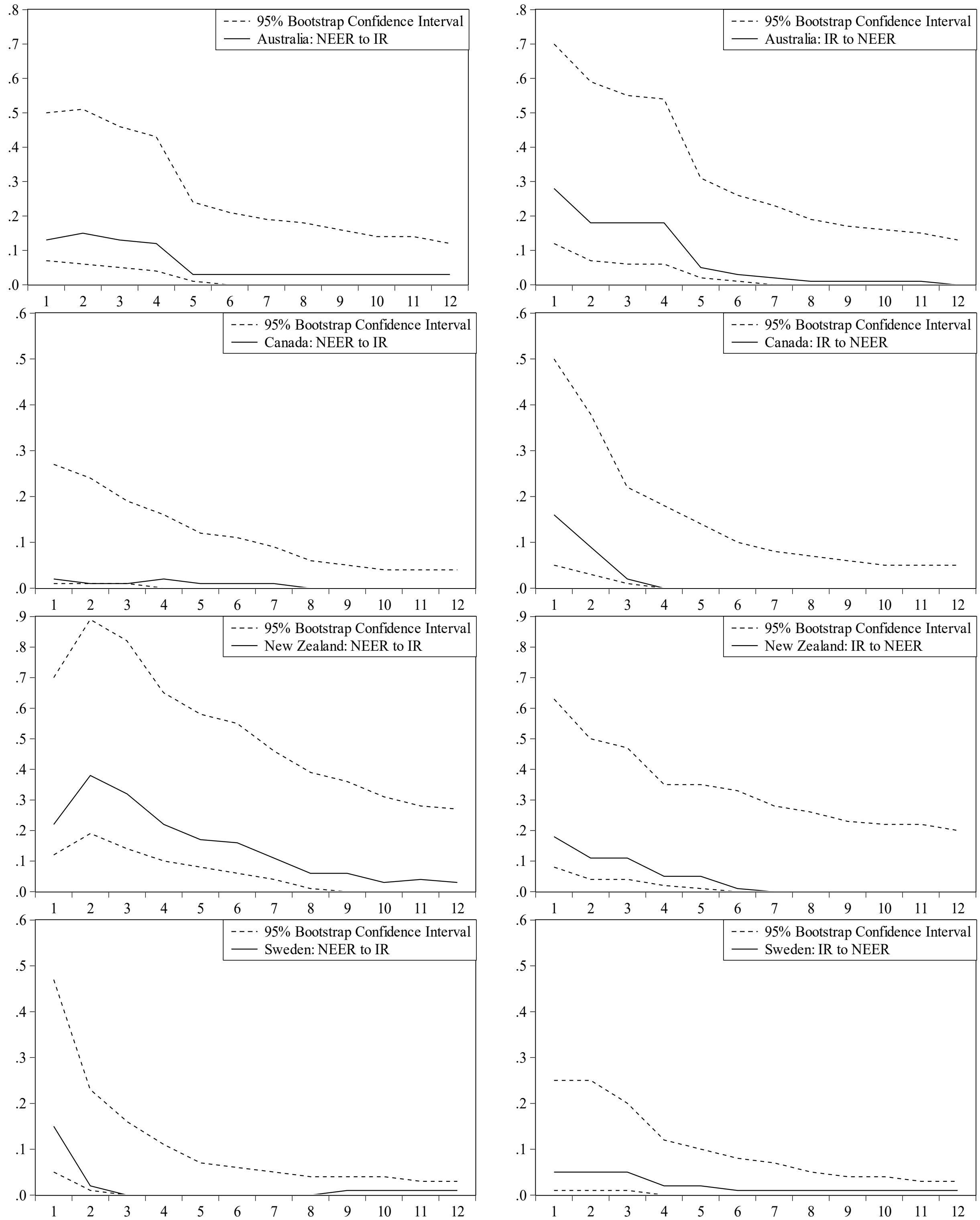

95% Bootstrap Confidence Interval
Germany: NEER to CPI
95% Bootstrap Confidence Interval
Germany: CPI to NEER
95% Bootstrap Confidence Interval
Italy: NEER to CPI
95% Bootstrap Confidence Interval
Italy: CPI to NEER
95% Bootstrap Confidence Interval
Netherlands: NEER to CPI
95% Bootstrap Confidence Interval
Netherlands: CPI to NEER
95% Bootstrap Confidence Interval
Spain: NEER to CPI
95% Bootstrap Confidence Interval
Spain: CPI to NEER

Figure 2.3a: Causality Measures between IR and NEER (Before Regime Change)



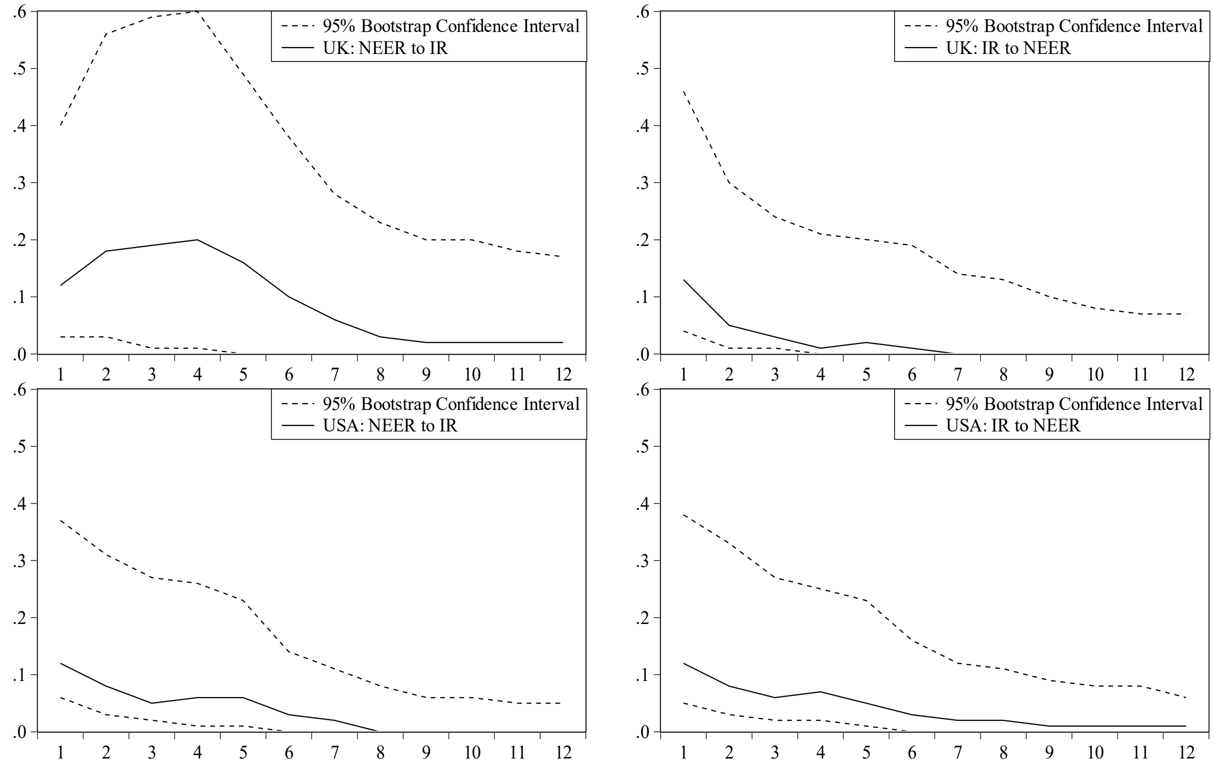
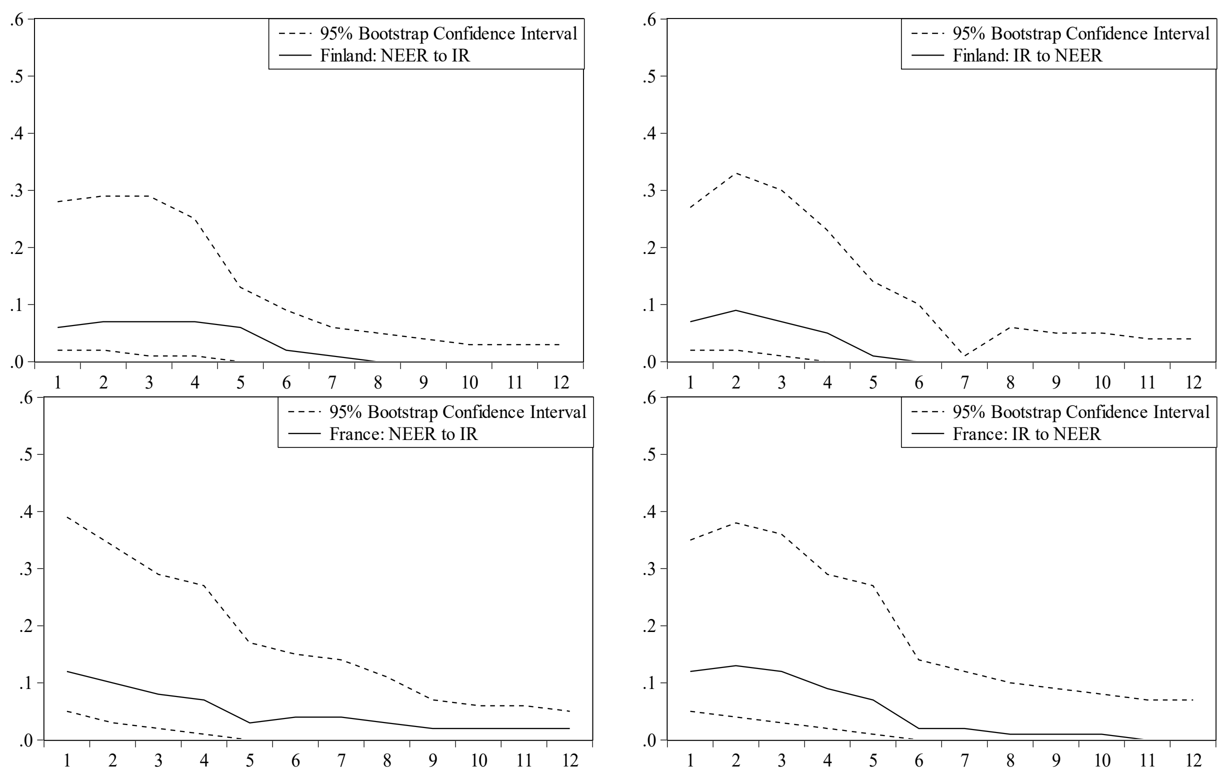


Figure 2.3b: Causality Measures between IR and NEER (Before Regime Change)



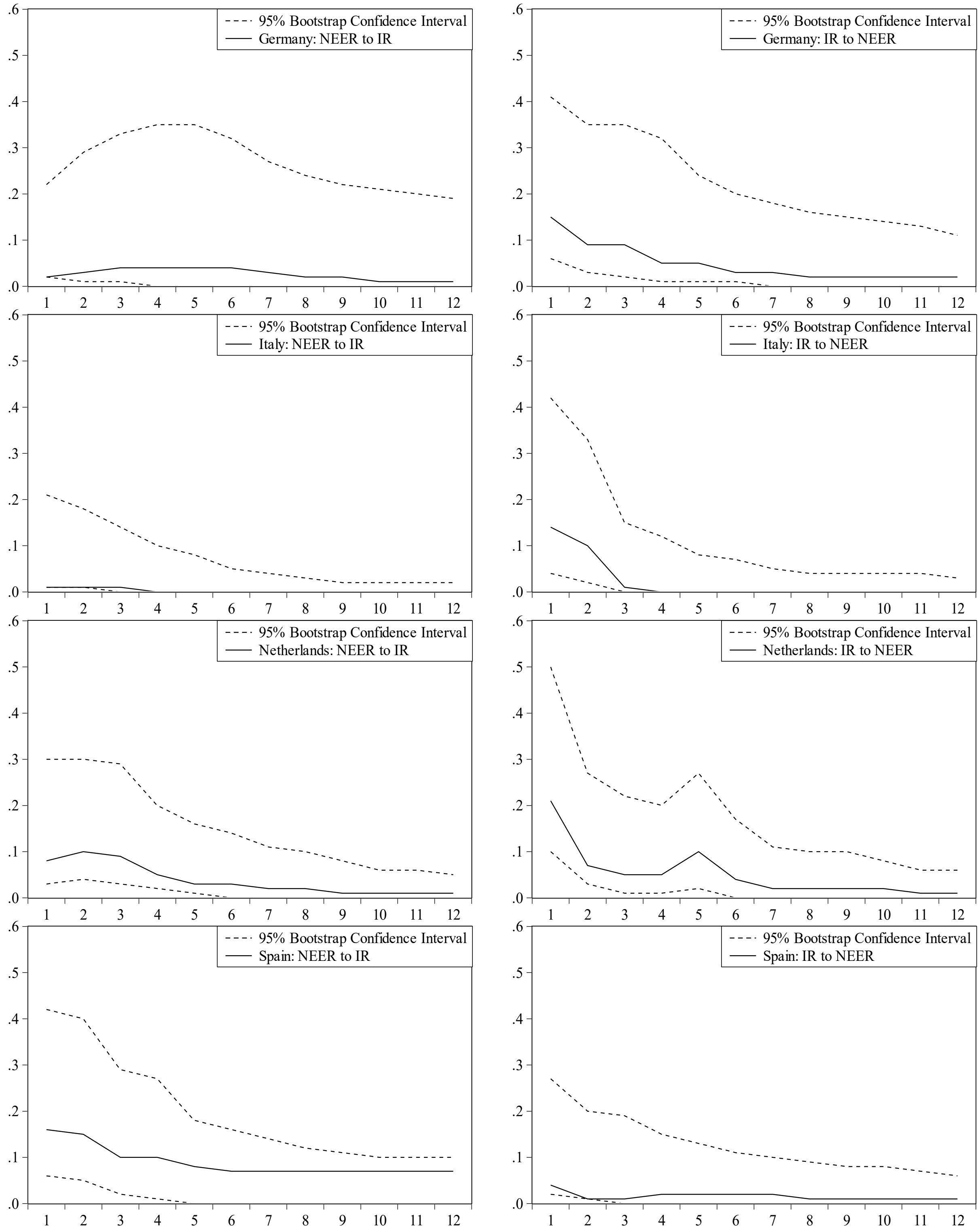
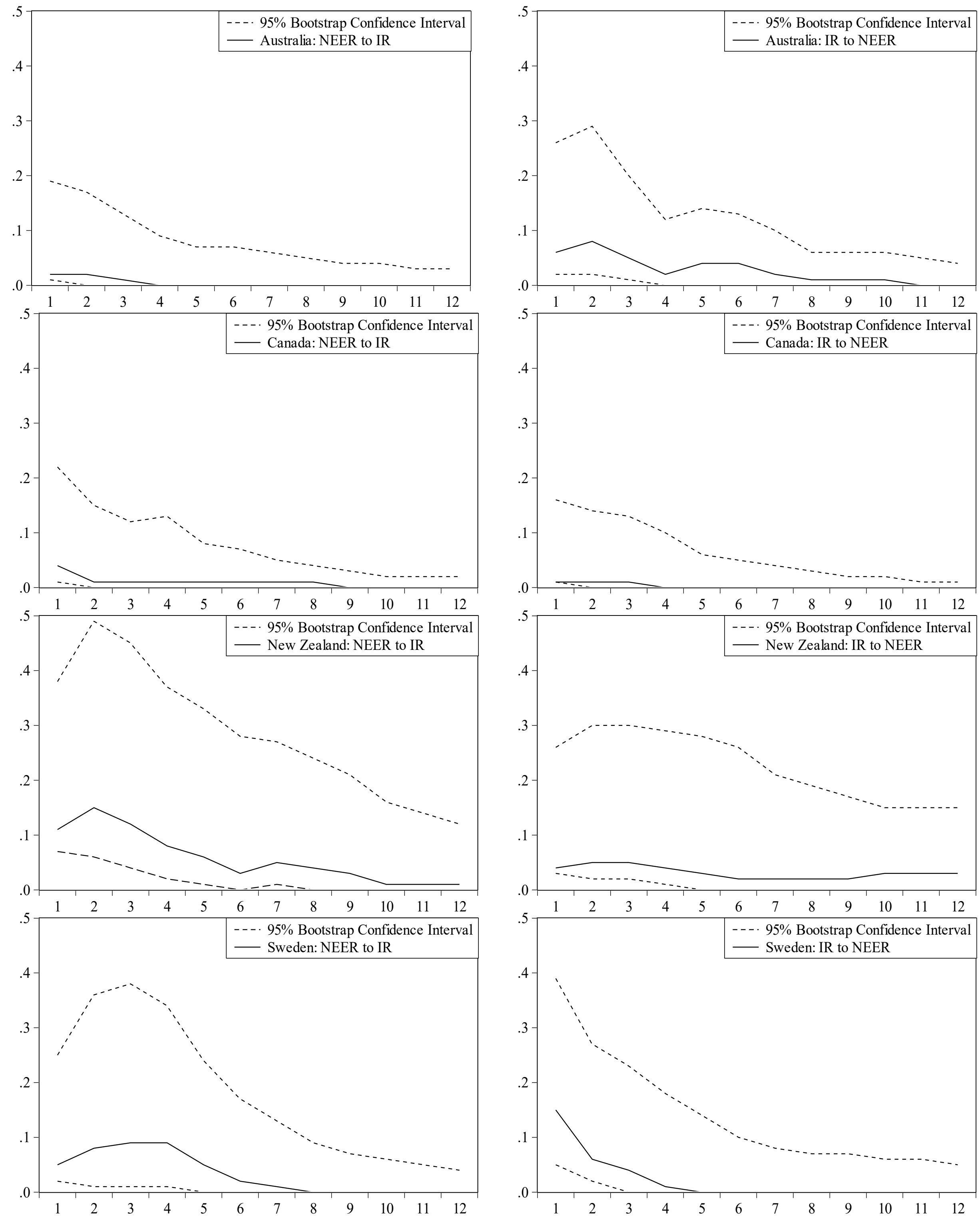

95% Bootstrap Confidence Interval
Germany: NEER to IR
95% Bootstrap Confidence Interval
Germany: IR to NEER
95% Bootstrap Confidence Interval
Italy: NEER to IR
95% Bootstrap Confidence Interval
Italy: IR to NEER
95% Bootstrap Confidence Interval
Netherlands: NEER to IR
95% Bootstrap Confidence Interval
Netherlands: IR to NEER
95% Bootstrap Confidence Interval
Spain: NEER to IR
95% Bootstrap Confidence Interval
Spain: IR to NEER
.0
.1
.2
.3
.4
.5
.6
1 2 3 4 5 6 7 8 9 10 11 12

Figure 2.4a: Causality Measures between IR and NEER (After Regime Change)



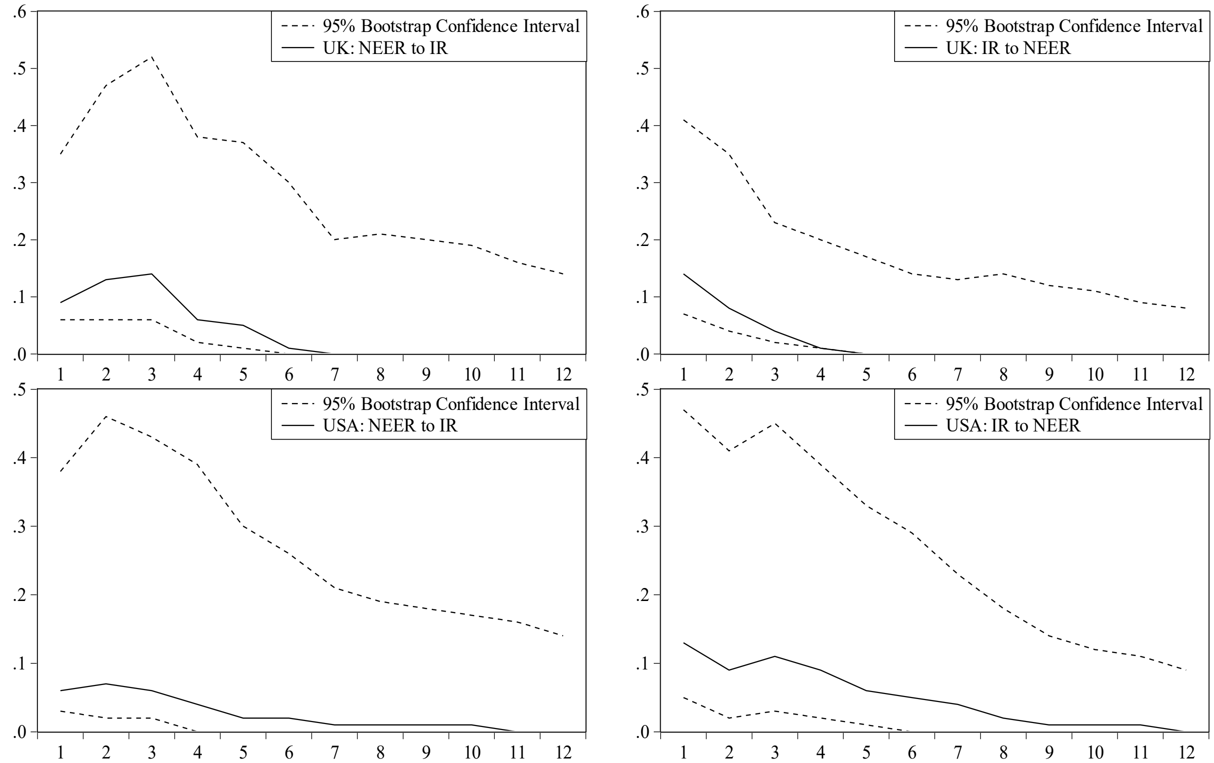
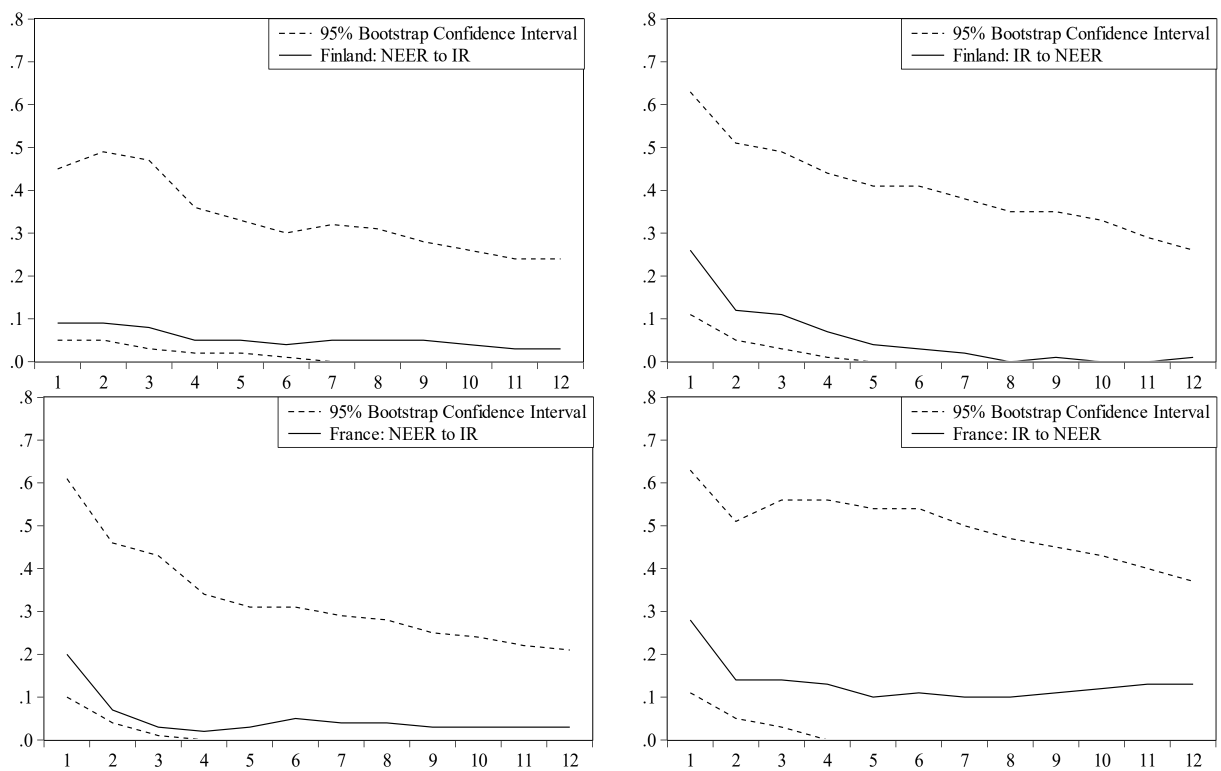


Figure 2.4b: Causality Measures between IR and NEER (After Regime Change)



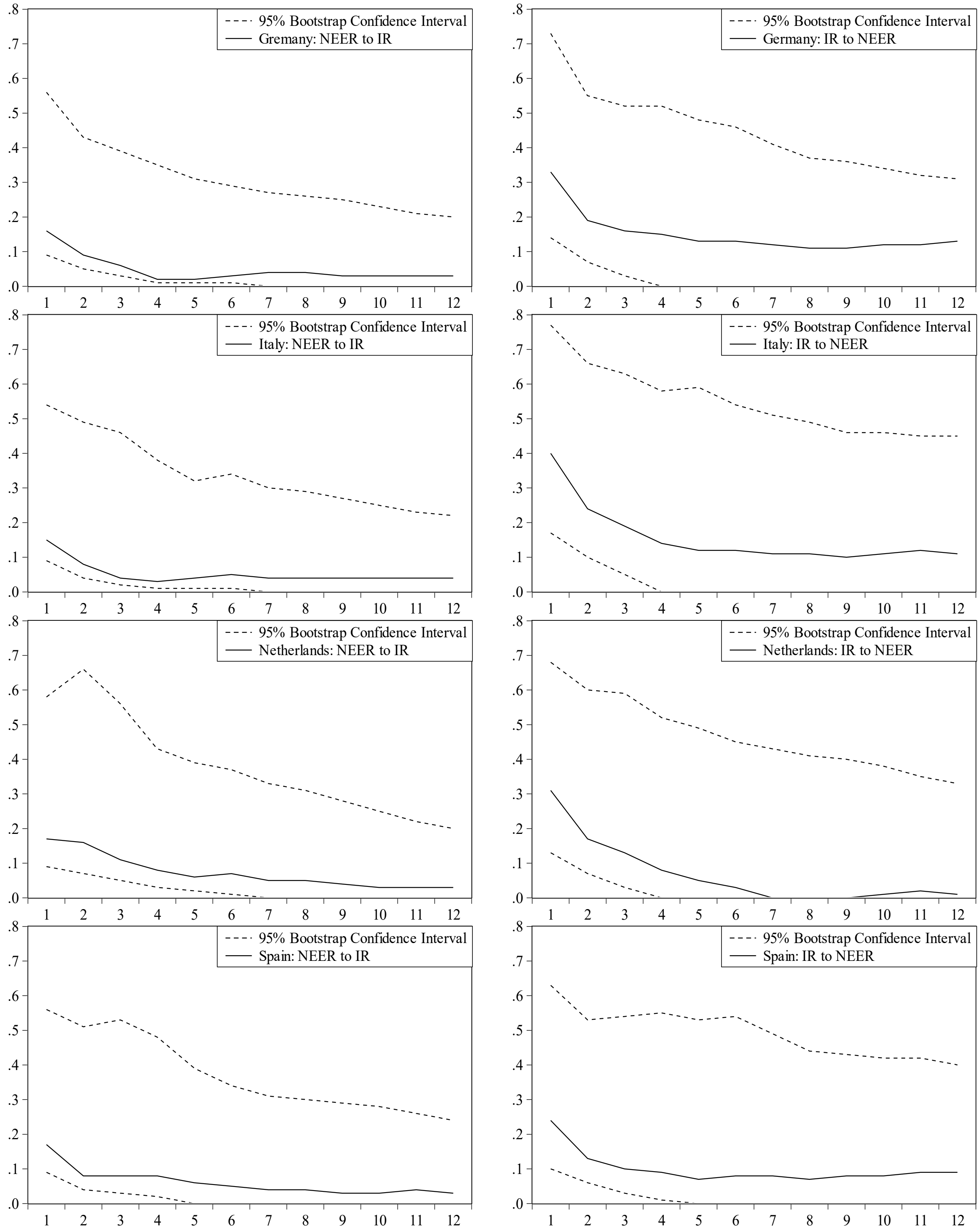
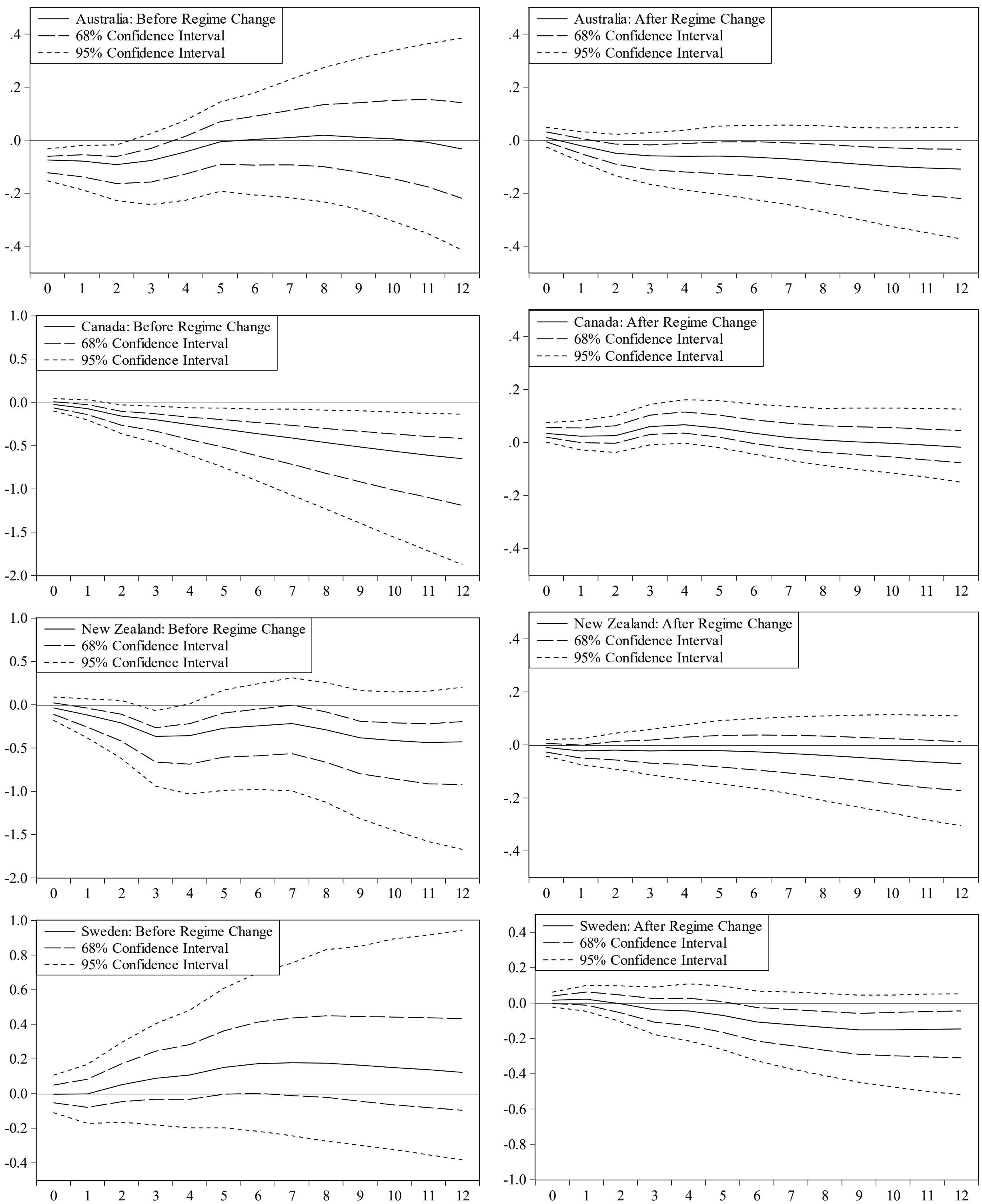

95% Bootstrap Confidence Interval
Gremany: NEER to IR
95% Bootstrap Confidence Interval
Germany: IR to NEER
95% Bootstrap Confidence Interval
Italy: NEER to IR
95% Bootstrap Confidence Interval
Italy: IR to NEER
95% Bootstrap Confidence Interval
Netherlands: NEER to IR
95% Bootstrap Confidence Interval
Netherlands: IR to NEER
95% Bootstrap Confidence Interval
Spain: NEER to IR
95% Bootstrap Confidence Interval
Spain: IR to NEER

Figure 2.5a: Response of CPI to 1% Appreciation of NEER



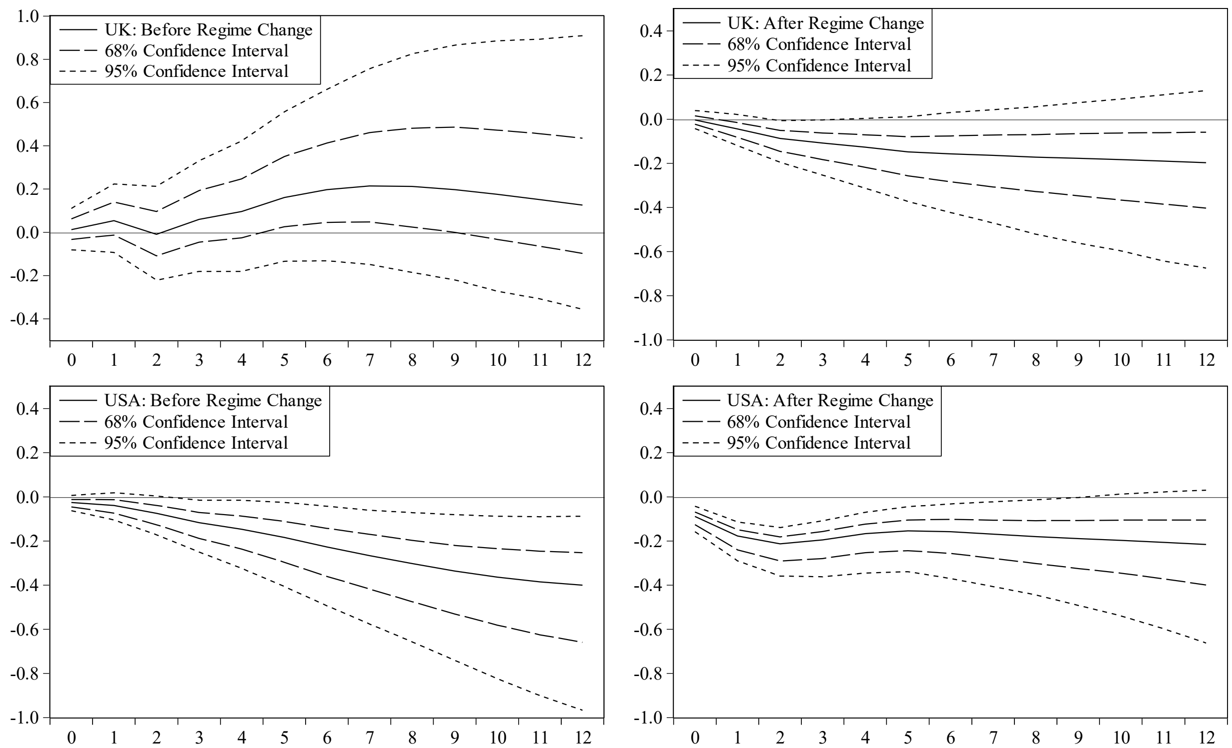
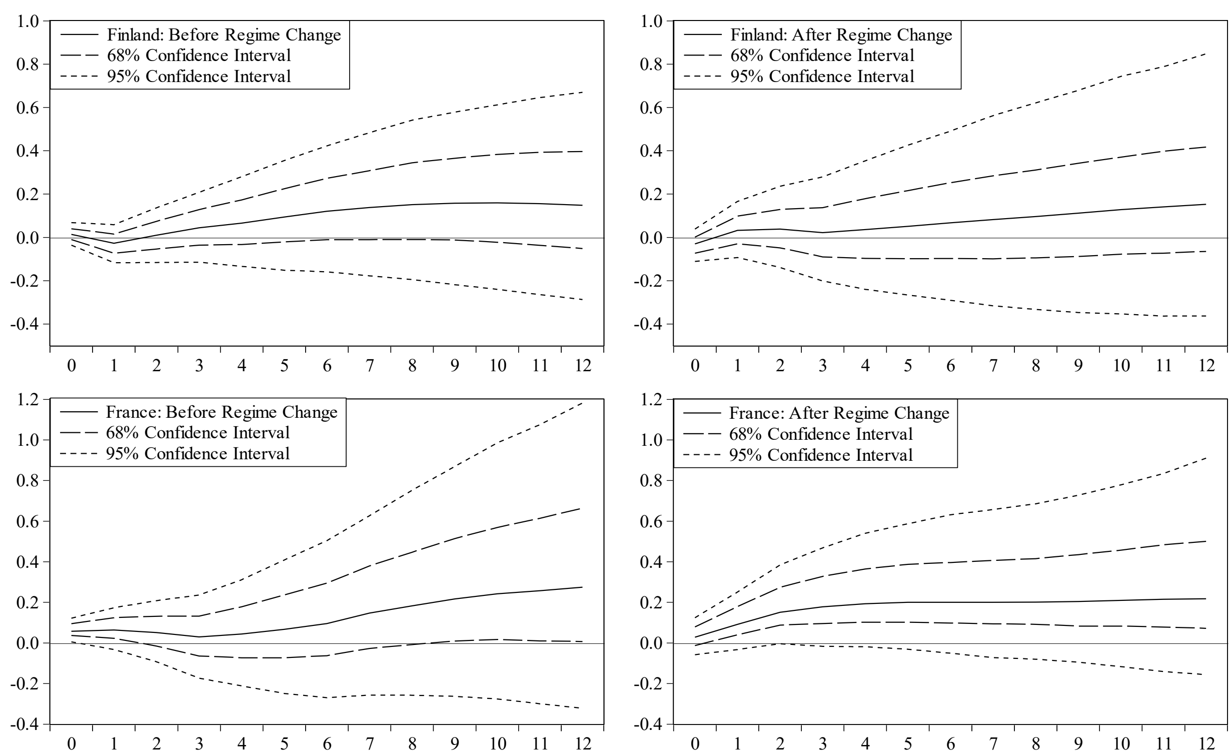


Figure 2.5b: Response of CPI to 1% Appreciation of NEER



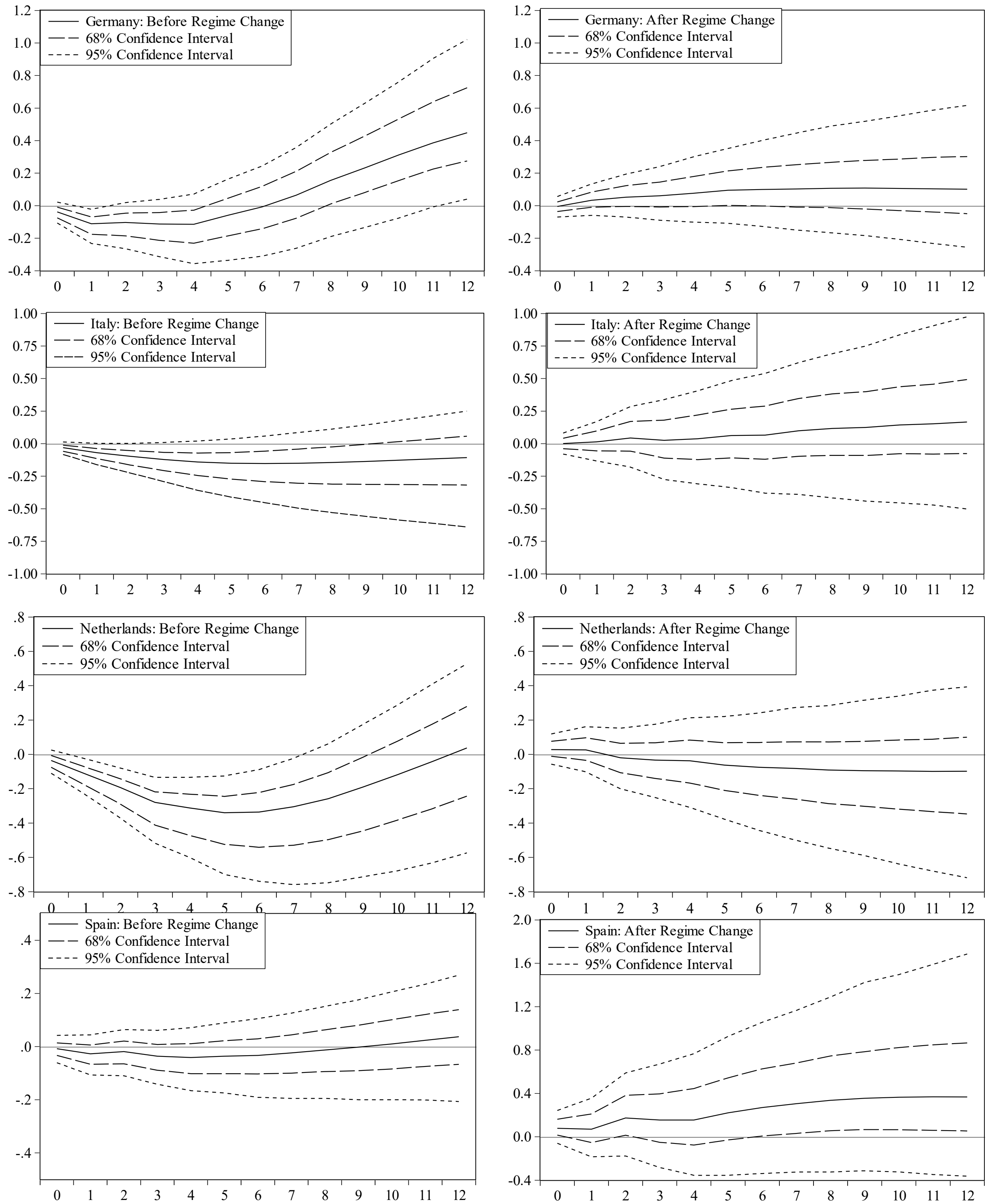
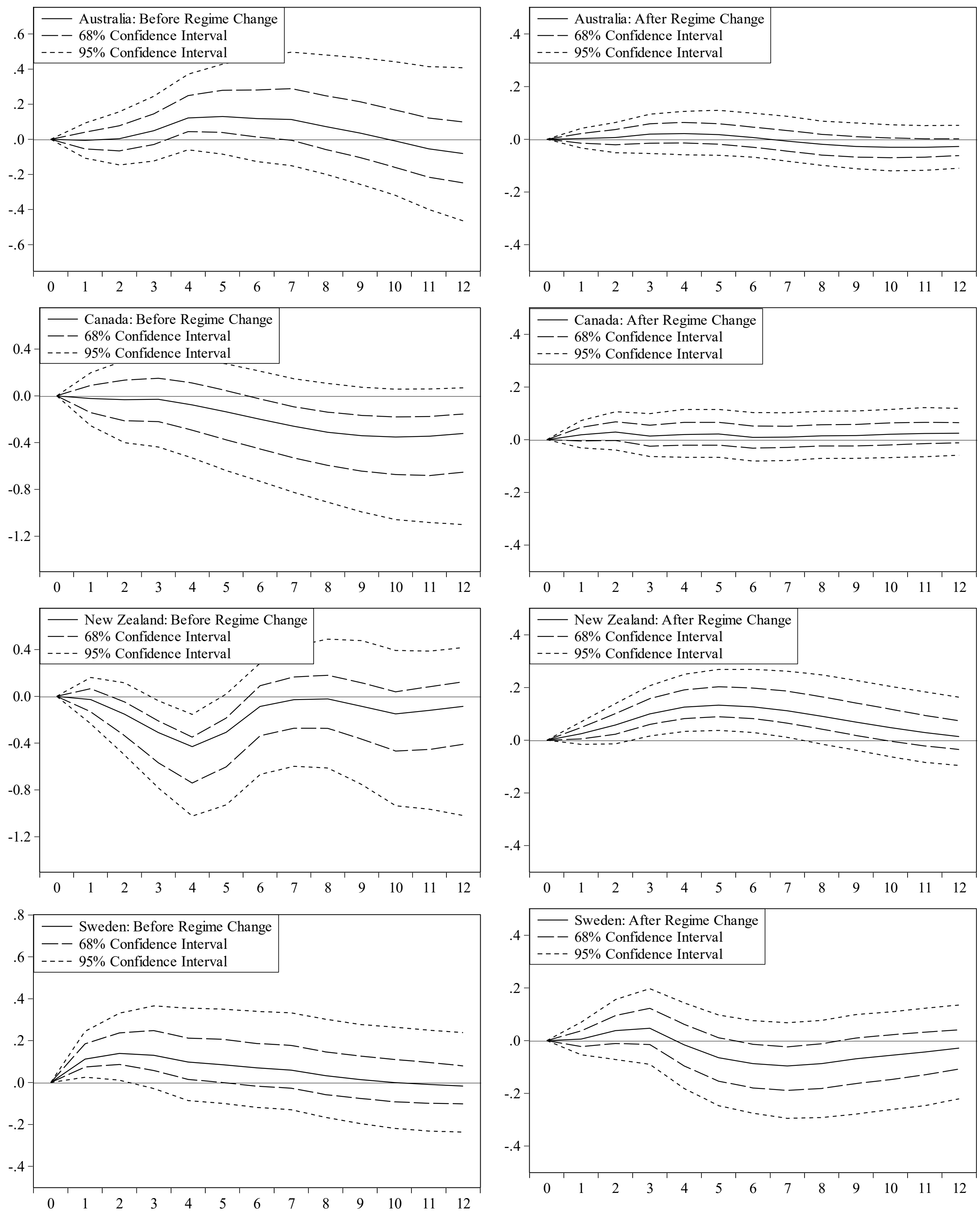

Germany: Before Regime Change
68% Confidence Interval
95% Confidence Interval
Germany: After Regime Change
68% Confidence Interval
95% Confidence Interval
Italy: Before Regime Change
68% Confidence Interval
95% Confidence Interval
Italy: After Regime Change
68% Confidence Interval
95% Confidence Interval
Netherlands: Before Regime Change
68% Confidence Interval
95% Confidence Interval
Netherlands: After Regime Change
68% Confidence Interval
95% Confidence Interval
Spain: Before Regime Change
68% Confidence Interval
95% Confidence Interval
Spain: After Regime Change
68% Confidence Interval
95% Confidence Interval

Figure 2.6a: Response of IR to 1% Appreciation of NEER



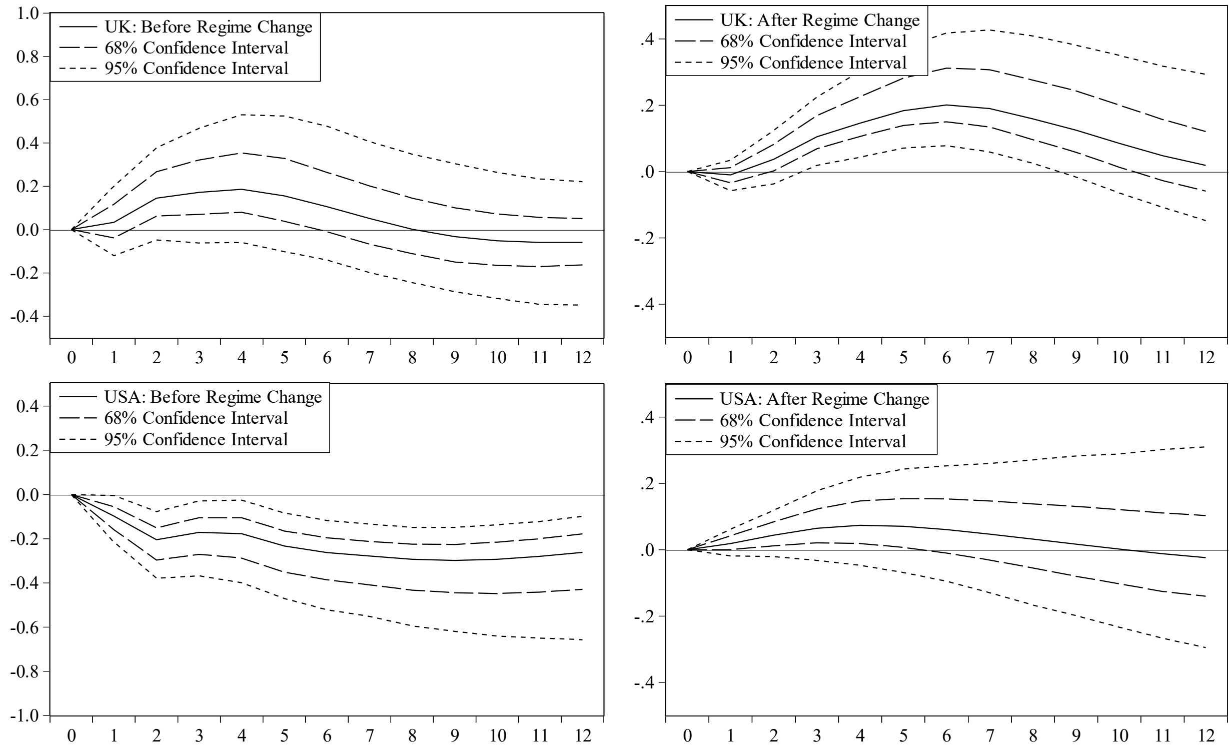
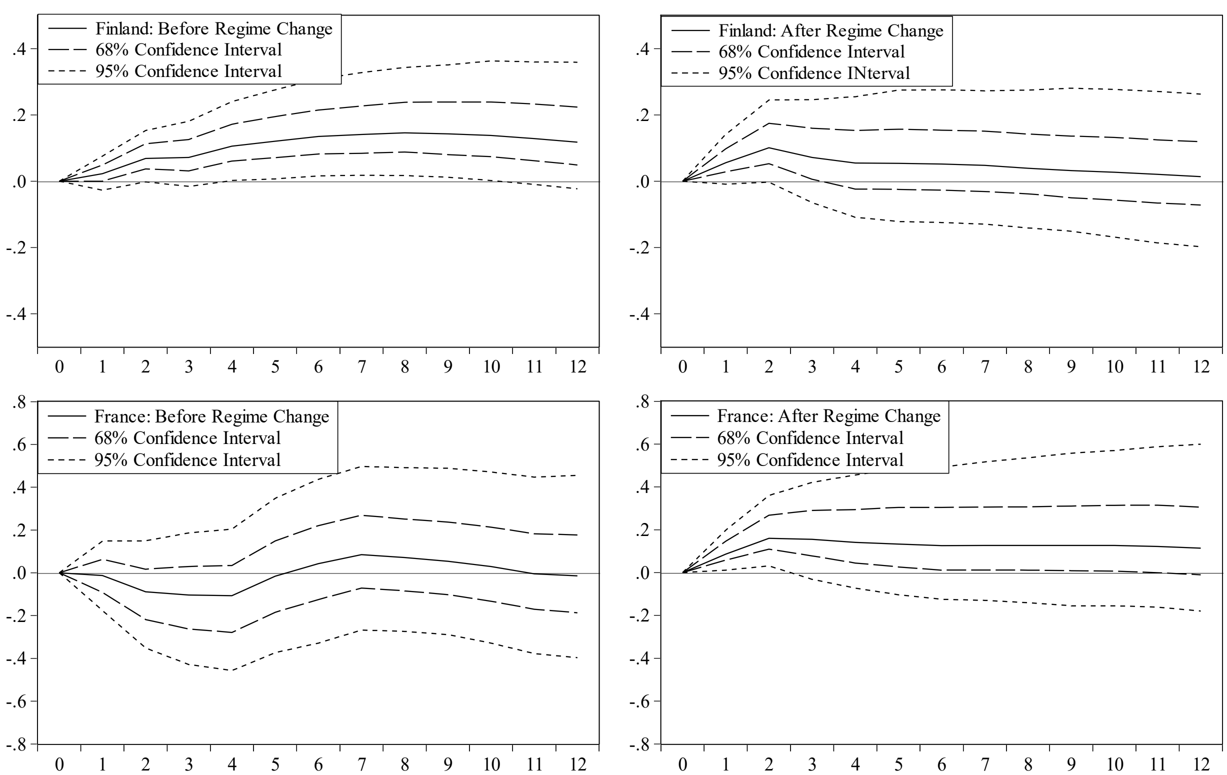
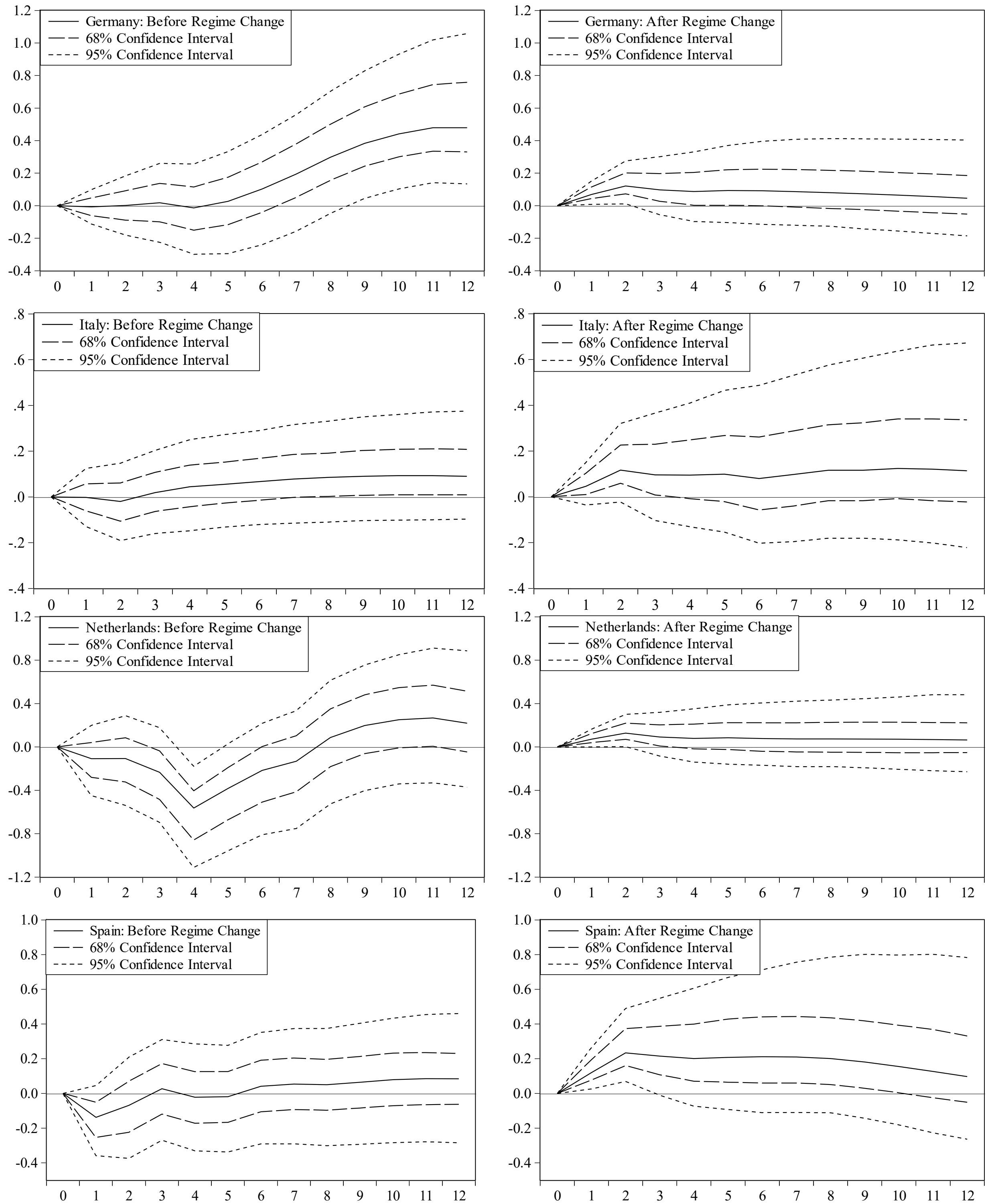


Figure 2.6b: Response of IR to 1% Appreciation of NEER




Germany: Before Regime Change
68% Confidence Interval
95% Confidence Interval
Germany: After Regime Change
68% Confidence Interval
95% Confidence Interval
Italy: Before Regime Change
68% Confidence Interval
95% Confidence Interval
Italy: After Regime Change
68% Confidence Interval
95% Confidence Interval
Netherlands: Before Regime Change
68% Confidence Interval
95% Confidence Interval
Netherlands: After Regime Change
68% Confidence Interval
95% Confidence Interval
Spain: Before Regime Change
68% Confidence Interval
95% Confidence Interval
Spain: After Regime Change
68% Confidence Interval
95% Confidence Interval

Appendix 2.1: Source and Description of Data

	CPI	NEER	Interest Rate (IR)
AUS	IMF-IFS	IMF-IFS	Policy Rate IMF-IFS, 1974-2016
CAN	IMF-IFS	IMF-IFS	Bank Rate Statistics Canada, 1974-2016
FIN	IMF-IFS	IMF-IFS	Discount Rate IMF-IFS, 1974-1998 Short-term Interest Rate OECD, 1999-2016
FRA	IMF-IFS	IMF-IFS	3-Month T-Bill IMF-IFS, 1974-1998 Short-term Interest Rate OECD, 1999-2016
GER	OECD	IMF-IFS	Short-term Interest Rate OECD, 1974-2016
ITA	IMF-IFS	IMF-IFS	Short-term Interest Rate OECD, 1978-2016
NET	IMF-IFS	IMF-IFS	Money Market Interest Rate IMF-IFS, 1974-1998 Short-term Interest Rate OECD, 1999-2016
NZL	IMF-IFS	IMF-IFS	3-Month Bank Bills OECD, 1974-2016
SPA	IMF-IFS	IMF-IFS	Short-term Interest Rate OECD, 1977-2016
SWE	IMF-IFS	IMF-IFS	Interbank Rate OECD, 1974-2016
UK	OECD	IMF-IFS	Policy Rate IMF-IFS, 1974-2016
USA	IMF-IFS	IMF-IFS	Federal Funds Rate FRED, 1974-2016

IMF-IFS: International Financial statistics data from the International Monetary Fund, available at http://data.imf.org/?sk=4C514D48-B6BA-49ED-8AB9-52B0C1A0179B&sId=1390030341854

OECD: Organization for Economic Cooperation and Development, available at http://stats.oecd.org/

Statistics Canada, available at http://www5.statcan.gc.ca/cansim/a08

FRED: Federal Reserve Economic Data, Federal Reserve Bank of St. Louis, available at https://fred.stlouisfed.org/series/FEDFUNDS

Appendix 2.2: Unit Root Test Statistics (Null Hypothesis: The variable has a unit root)
Lag selection: Ttest (with a minimum of 1 lag)

Variable	Before Regime Change				After Regime Change			
	ADF Test		DF-GLS		ADF Test		DF-GLS	
	Constant	Constant & Trend	Constant	Constant & Trend	Constant	Constant & Trend	Constant	Constant & Trend
AUS-CPI	-2.73^{c}	0.15	-0.12	-0.52	-0.62	-2.38	1.35	-2.35
ΔAUS-CPI	-1.60	-3.39^{c}	-0.11	-2.13	-4.89^{a}	-4.88^{a}	-4.11^{a}	-4.59^{a}
ΔΔAUS-CPI	-11.91^{a}	-11.81^{a}	-1.37	-2.82				
AUS-IR	-2.14	-1.76	-1.52	-2.07	-0.33	-1.56	-0.38	-1.43
ΔAUS-IR	-3.13^{b}	-4.17^{a}	-1.72^{c}	-3.26^{b}	-5.15^{a}	-5.18^{a}	-5.05^{a}	-5.26^{a}
AUS-NEER	-1.83	-2.86	-0.38	-2.65	-2.10	-2.64	-0.69	-2.66
ΔAUS-NEER	-3.06^{b}	-3.06	-2.28^{b}	-2.88^{c}	-5.97^{a}	-5.93^{a}	-5.99^{a}	-5.99^{a}
CAN-CPI	-2.52	-1.10	0.56	-1.89	-0.13	-2.28	0.83	-1.23
ΔCAN-CPI	-2.13	-2.84	-1.07	-2.97^{c}	-4.19^{a}	-4.17^{a}	-1.48	-2.63
ΔΔCAN-CPI	-7.37^{a}	-7.35^{a}	-2.10^{b}	-6.79^{a}				
CAN-IR	-2.50	-2.51	-1.94^{c}	-2.44	-2.60^{c}	-3.24^{c}	0.45	-1.58
ΔCAN-IR	-3.48^{b}	-3.45^{c}	-3.41^{a}	-5.55^{a}	-3.75^{a}	-3.66^{b}	-0.63	-1.56
CAN-NEER	-1.96	-2.40	-0.84	-2.45	-1.31	-1.71	-1.25	-1.79
ΔCAN-NEER	-3.33^{b}	-3.40^{c}	-3.34^{a}	-3.44^{b}	-3.48^{b}	-3.48^{b}	-1.30	-2.89^{c}
FIN-CPI	-2.10	-1.06	0.49	-0.43	-0.75	-2.71	0.53	-2.81
ΔFIN-CPI	-2.29	-4.78^{a}	-0.15	-2.49	-3.50^{b}	-3.47^{c}	-2.88^{a}	-3.40^{b}
ΔΔFIN-CPI	-4.05^{a}		-1.83^{c}	-2.77^{c}				
FIN-IR	-0.53	-2.38	-0.05	-2.36	-1.40	-3.56^{b}	-1.10	-3.11^{c}
ΔFIN-IR	-5.55^{a}	-5.61^{a}	-5.56^{a}	-5.63^{a}	-4.34^{a}		-3.44^{a}	
FIN-NEER	-1.97	-2.38	-0.71	-2.31	-2.81^{c}	-2.51	-2.28^{b}	-2.92^{c}
ΔFIN-NEER	-4.47^{a}	-4.48^{a}	-4.02^{a}	-4.27^{a}	-3.33^{b}	-3.55^{b}		-2.03
FRA-CPI	-2.26	-1.68	-0.87	-1.41	-2.41	0.65	-0.14	-1.19
ΔFRA-CPI	-0.72	-2.87	0.37	-3.03^{c}	-2.27	-3.79^{b}	-2.20^{b}	-2.78
ΔΔFRA-CPI	-3.36^{b}	-3.34^{c}	-0.34	-1.22				
FRA-IR	-2.00	-2.62	-1.31	-2.66	-1.40	-3.56^{b}	-1.10	-3.11^{c}
ΔFRA-IR	-5.69^{a}	-5.81^{a}	-6.60^{a}	-5.92^{a}	-4.34^{a}		-3.44^{a}	
FRA-NEER	-1.85	-1.52	-1.34	-1.39	-2.79^{c}	-2.09	-2.54^{b}	-2.66
ΔFRA-NEER	-2.58^{c}	-2.94	-0.49	-1.97	-2.87^{c}	-5.03^{a}		-1.78
GER-CPI	-1.39	-1.84	0.36	-1.66	-1.32	-1.06	0.74	-1.50
ΔGER-CPI	-2.48	-2.70	-1.52	-2.61	-3.79^{a}	-4.00^{b}	-2.76^{a}	-3.93^{a}
ΔΔGER-CPI	-7.12^{a}	-7.08^{a}	-2.53^{b}	-7.16^{b}				
GER-IR	-2.80^{c}	-2.97	-1.64^{c}	-3.01^{c}	-1.40	-3.56^{b}	-1.10	-3.11^{c}
ΔGER-IR	-4.03^{a}	-3.99^{b}	-1.19	-2.12	-4.34^{a}		-3.44^{a}	
GER-NEER	-1.93	-2.39	0.90	-1.51	-2.95^{b}	-2.27	-2.44^{b}	-2.72
ΔGER-NEER	-5.41^{a}	-5.62^{a}	-0.54	-1.81		-3.63^{b}		-1.89
ITA-CPI	-1.55	-2.12	0.24	-2.05	-2.09	0.83	-0.21	-1.21
ΔITA-CPI	-2.83^{c}	-1.99	-0.72	-2.51	-3.21^{b}	-3.62^{b}	-2.76^{a}	-3.41^{b}
ΔΔITA-CPI	-3.76^{a}	-4.21^{a}	-1.35	-1.93				
ITA-IR	-0.52	-3.20^{c}	-0.72	-1.65	-1.40	-3.56^{b}	-1.10	-3.11^{c}
ΔITA-IR	-5.82^{a}	-6.16^{a}	-5.83^{a}	-6.11^{a}	-4.34^{a}		-3.44^{a}	
ITA-NEER	-2.07	-2.56	-0.61	-2.28	-3.31^{b}	-2.52	-1.96^{b}	-2.77
ΔITA-NEER	-4.13^{a}	-4.21^{a}	-2.63^{a}	-4.28^{a}		-5.32^{a}		-1.81
NET-CPI	-2.11	-2.81	0.83	-1.21	-1.87	-2.60	0.43	-1.81
ΔNET-CPI	-2.49	-2.61	-0.74	-2.11	-3.66^{a}	-4.26^{a}	-2.68^{a}	-4.09^{a}

ΔΔNET-CPI	-10.02[a]	-10.13[a]	-9.99[a]	-10.09[a]				
NET-IR	-2.49	-2.69	-1.37	-2.53	-1.40	-3.56[b]	-1.10	-3.11[c]
ΔNET-IR	-3.14[b]	-9.01[a]	-1.57	-2.23	-4.34[a]		-3.44[a]	
NET-NEER	-1.97	-2.29	0.46	-1.88	-2.81[c]	-2.15	-2.49[b]	-2.74
ΔNET-NEER	-4.64[a]	-4.94[a]	-2.32[b]	-4.73[a]		-4.97[a]		-1.98
NZL-CPI	-2.23	-1.64	0.16	-1.77	-0.63	-2.06	0.21	-2.12
ΔNZL-CPI	-2.96[b]	-3.41[c]	-2.73[a]	-3.11[c]	-2.39	-2.38	-0.39	-1.83
ΔΔNZL-CPI					-6.41[a]	-6.40[a]	-0.21	-1.44
NZL-IR	-1.84	-2.08	-0.76	-2.31	-2.26	-3.41[c]	-0.34	-2.37
ΔNZL-IR	-3.79[a]	-4.12[b]	-3.57[a]	-4.13[a]	-5.65[a]		-5.60[a]	-5.69[a]
NZL-NEER	-2.12	-2.71	0.58	-2.17	-2.38	-3.59[b]	-2.30[b]	-3.38[b]
ΔNZL-NEER	-6.35[a]	-6.70[a]	-2.68[a]	-6.34[a]	-2.78[c]			
SPA-CPI	-2.28	-1.83	-0.56	-0.50	-2.20	0.89	-0.02	-0.74
ΔSPA-CPI	-1.38	-1.75	0.48	-1.76	-3.45[b]	-4.54[a]	-3.48[a]	-4.01[a]
ΔΔSPA-CPI	-4.90[a]	-4.96[a]	-0.43	-2.04				
SPA-IR	0.14	-3.03	-0.30	-1.49	-1.40	-3.56[b]	-1.10	-3.11[c]
ΔSPA-IR	-5.28[a]	-5.36[a]	-5.33[a]	-5.44[a]	-4.34[a]		-3.44[a]	
SPA-NEER	-1.51	-1.63	-0.17	-1.42	-2.56	-2.05	-2.27[b]	-1.77
ΔSPA-NEER	-4.54[a]	-5.51[a]	-4.19[a]	-4.41[a]	-3.11[b]	-3.23[c]		-2.27
SWE-CPI	-1.66	-1.10	0.32	-1.10	-0.46	-2.72	1.35	-2.77
ΔSWE-CPI	-2.48	-2.93	-2.30[b]	-2.57	-4.06[a]	-4.00[b]	-3.90[a]	-4.04[a]
SWE-IR	-2.57	-3.04	-0.59	-2.83[c]	-2.22	-3.45[c]	0.47	-1.25
ΔSWE-IR	-4.36[a]	-4.41[a]	-3.05[a]	-4.19[a]	-4.87[a]	-4.93[a]	-0.56	-1.89
SWE-NEER	-0.81	-2.17	0.14	-2.03	-3.25[b]	-3.23[c]	-2.75[a]	-3.26[b]
ΔSWE-NEER	-4.70[a]	-4.67[a]	-3.76[a]	-3.99[a]				
UK-CPI	-2.27	-2.18	0.57	-1.12	-0.20	-2.86	0.29	-2.22
ΔUK-CPI	-1.88	-2.27	-0.63	-2.17	-2.17	-2.15	-1.71[c]	-2.05
ΔΔUK-CPI	-3.43[b]	-3.42[c]	-0.88	-1.97				
UK-IR	-3.15[b]	-3.12	-3.03[a]	-3.13[c]	-2.14	-3.95[b]	-0.28	-3.00[c]
ΔUK-IR		-4.53[a]		-4.29[a]	-5.82[a]		-4.72[a]	
UK-NEER	-2.51	-3.17[c]	-0.19	-2.50	-1.04	-1.56	-0.80	-1.64
ΔUK-NEER	-3.52[b]	-3.41[c]	-2.05[b]	-3.50[b]	-3.79[a]	-3.88[b]	-3.72[a]	-3.93[a]
USA-CPI	-2.19	-2.18	-0.09	-0.91	-1.86	-1.44	0.67	-1.13
ΔUSA-CPI	-1.91	-3.58[b]	-1.00	-3.65[a]	-3.14[b]	-3.48[b]	-3.06[a]	-5.65[a]
ΔΔUSA-CPI	-3.35[b]		-3.08[a]					
USA-IR	-2.36	-3.26[c]	-1.73[c]	-3.11[b]	-2.26	-3.91[b]	-1.41	-3.68[b]
ΔUSA-IR	-3.98[a]	-4.02[b]	-1.64[c]		-3.58[a]		-3.52[a]	
USA-NEER	-0.78	-2.82	1.07	-2.78[c]	-1.76	-0.85	-1.57	-1.68
ΔUSA-NEER	-4.73[a]	-4.69[a]	-2.11[b]	-3.90[a]	-2.10	-2.56	-2.08[b]	-2.36

Critical Values for ADF Statistic:

With constant, the critical values are -3.53, -2.91 and -2.59 for 1%, 5% and 10% level of significance, respectively; and with constant and trend, the critical values are -4.10, -3.48 and -3.17 for 1%, 5% and 10% level of significance, respectively.

Critical Values for DF-GLS Statistic:

With a constant, the critical values are -2.60, -1.95 and -1.61 for the 1%, 5% and 10% levels of significance, respectively; and with constant and trend, the critical values are -3.72, -3.14 and -2.85 for 1%, 5% and 10% levels of significance, respectively.

[a,b,c] represent the level of significance at 1%, 5% and 10%, respectively.

Chapter 3

Are the Responses of Consumer Prices Asymmetric to Exchange Rate Appreciations and Depreciations?

3.1. Introduction

The objective of this chapter is to examine potential asymmetric adjustments of consumer prices to exchange rate appreciations and depreciations. The response of prices to changes in the exchange rate is commonly known as exchange rate pass-through (ERPT). This is an important concern for central banks for two reasons: first, a low ERPT can provide relatively greater independence for domestic monetary policies, and second, ERPT to domestic prices determines the real exchange rate which affects price competitiveness in the global market and, hence, trade balances. For monetary policy to have the desired outcome, it is necessary to take into account of the potential directional asymmetries, i.e., whether the effect of an exchange rate appreciation is exactly the opposite of a depreciation of the same size.

There is a large literature that estimates the degree of ERPT to the aggregate price level such as import prices, export prices, producer prices and consumer prices. However, most of these studies employ linear models in the estimation process. If a 1 percent depreciation of exchange rate leads to 1 percent increase in prices, the pass-through is known as complete; if it is less than one, the pass-through is incomplete or partial. Although complete ERPT is possible for import prices, ERPT to consumer prices is expected to be incomplete and lower than import prices as imported goods are only one component of all the goods and services sold (or an intermediate good in the production process) in the economy. Several studies (see Goldberg and Knetter (1997), Campa and Goldberg (2005) and Bussière, Chiaie and Peltonen (2014) among others) show that ERPT is incomplete for import prices and ERPT to consumer prices is lower than to import prices (see McCarthy (2007), and Ito and Sato (2008) among others). For example, Campa and Goldberg (2005) find that the average pass-through to import prices for a sample of 23 OECD countries is 46 percent in the short run and 64 percent in the long run. Choudhri, Faruqee and Hakura (2005) find that the ERPT to consumer prices is in the range of 0.04 percent to 0.20 percent for non-US

G-7 countries. However, these results can be questioned as they do not accommodate potential asymmetries into the model.

The economic literature has been aware of asymmetric price adjustments at the firm level since the 1980s. An example of pricing strategy of foreign firms is provided in Goldberg and Knetter (1997). Between January 1994 and April 1995, the Japanese Yen appreciated by 34 percent against the US Dollar (USD). In this period, the price of Toyota Celica ST Coupe increased by less than 2 percent and the retail price of a large-screen SONY Trinitron decreased by 15 percent. Thus, in order to maintain (or increase) their market share, Toyota and SONY decreased their export prices in Yen by lowering their mark-ups (or profit margins). This type of pricing strategy led to the analysis of asymmetric price adjustments using firm-level or industry-level data on the export and import prices of some developed countries. However, studies with aggregate prices, such as producer or consumer prices, started quite recently and are very few in number.

Asymmetries can be introduced to a linear model through dummy variable(s) or decomposing the required variable based on certain conditions. This type of nonlinearity is often used with different variants of single equation method to analyze asymmetries. The corresponding slope-based tests are useful to examine the presence of asymmetry, but they are not informative about the degree of asymmetry in responses at multiple horizons to positive and negative structural shocks. A dynamic structural vector autoregression (VAR) model is more useful in this regard as it allows estimation of asymmetric impulse responses; and permits examination of the degree of asymmetries in these responses. Kilian and Vigfusson (2011) propose a procedure to estimate these impulse responses from a nonlinear VAR model and to examine the statistical significance of the asymmetry in responses.

The aim of this chapter is to test for potential asymmetries in the responses of consumer prices to exchange rate appreciations and depreciations for 12 Asian countries, located in East, Southeast and South Asia. The countries analyzed in this chapter are Bangladesh (BGD), China (CHN), Indonesia (IDN), India (IND), Japan (JPN), Republic of Korea (KOR), Malaysia (MYS), Pakistan (PAK), the Philippines (PHL), Singapore (SGP), Taiwan (TWN) and Thailand (THA). Based on 2016 data, almost 55 percent of the world population live in this geographic region and almost 50 percent of the world population live in the selected 12 countries of this chapter.[4] Still, there is hardly any study that focuses on the countries of this region (except Japan) to examine asymmetric price adjustments to exchange rate appreciations and depreciations. This selection of the countries is also quite diverse since some are developing countries, some are emerging countries and others are high income countries. This diversity will allow us to explore whether the extent of pass-through and directional asymmetry varies across countries that are at different stages of economic development and have a different macroeconomic environment. Thus, this chapter contributes to the empirical evidence on directional asymmetries, which is currently limited to a few developed countries.[5] To the best of our knowledge, this is also the first paper to examine this issue by an impulse response-based test.

Among these 12 countries, Indonesia, Korea, Philippines and Thailand started to target inflation in the early 2000s. The exchange rate regime of only Japan and Korea remained unchanged and freely floating throughout the time period under investigation. The remaining 10 countries went through some regime changes. A de facto classification of the exchange rate regime by the IMF is presented in Appendix 3.1. It shows that the monetary authority in some of these 10 countries started to allow free float of the exchange rate in the latter periods while others (for example, Bangladesh, China, Malaysia, Pakistan, Singapore and Taiwan) still intervene in the foreign exchange market.

We use quarterly data from 1994 to 2016 for our estimation. Consumer prices and the exchange rate are represented respectively by consumer price index (CPI) and nominal effective exchange rate (NEER). NEER is the trade-weighted average of the nominal exchange rate, defined as the amount of foreign currency required to buy one unit of the domestic currency. Hence, an increase in NEER implies an appreciation of the domestic currency. This time period is chosen as the NEER data is available for 6 of our selected countries only from 1994 or later. The NEER data are downloaded from the International Financial Statistics (IFS) of International Monetary Fund

[4] Population data for all countries, except Taiwan, are taken from the World Bank website to compute the share of these region (and countries) in world population. For Taiwan, population data is taken from the National Statistics, Republic of China (Taiwan) website.

[5] Asymmetries in the price adjustment can also arise from the size of the exchange rate change since a large exchange rate shock can have a disproportionately greater effect on prices compared to a smaller shock, known as size asymmetries. In this chapter, we focus only on the directional asymmetry and leave the size asymmetries for future research.

(IMF) website (for China, Japan, Malaysia, Pakistan, Philippines and Singapore), Bank for International Settlements (BIS) (for Indonesia, India, Korea, Thailand and Taiwan) and from Darvas (2012) for Bangladesh. The data on CPI and nominal exchange rate vis-à-vis USD for all countries, except Taiwan, are downloaded from the IMF-IFS. For Taiwan, the data on the nominal exchange rate and CPI are taken from the website of the Central Bank of the Republic of China and the website of National Statistics of the Republic of China, respectively. The details on the sources of the data are presented in Appendix 3.2.

An overview of our main results is as follows. We have found evidence of asymmetric responses of CPI to NEER appreciations and depreciations in 6 (China, Indonesia, Korea, Malaysia, Pakistan and Philippines) out of the 12 countries. Surprisingly, in 4 countries (Malaysia, Philippines, Singapore and Taiwan) either an appreciation of the exchange rate increases the price level and/or a depreciation decreases the price level, which is just the opposite of theoretical expectation. Further, we do not find any pattern that depreciations are passed-through to the price level more than appreciations. Our findings contrast with those of Delatte and Lopez-Villavicencio (2012) in this regard. Among the countries that have (theoretically) expected responses, there are wide variations in the degree of pass-through, especially when exchange rate depreciates. Our results are robust to changes in model specification but are sensitive to the measure of the exchange rate, e.g., when we replace NEER by USD-based nominal exchange rate. These results have important policy implications, especially for the countries that have theoretically unexpected responses.

The organization of this chapter is as follows. Section 3.2 presents the arguments for asymmetric responses of consumer prices to exchange rate appreciations and depreciations. Section 3.3 provides empirical evidence on the asymmetric responses found in the contemporary literature. A description of the econometric methodology is presented in Section 3.4. Section 3.5 explains the results obtained from the baseline estimation and Section 3.6 examines the robustness of the baseline results. Finally, Section 3.7 summarizes the main results and discusses the scope for future research.

3.2. Theoretical Arguments for Asymmetric Pass-through

Exchange rate movements directly affect import prices. They affect consumer prices via changes in import prices, both as final goods and as inputs to domestic goods. If the ERPT to import prices is higher, the ERPT to consumer prices is also likely to be higher. The degree of ERPT to import prices and consumer prices largely depends on foreign exporters' pricing strategies and distribution costs in the local market, which itself depend on a number of microeconomic and macroeconomic factors. The response of domestic prices can be asymmetric in response to exchange rate appreciations and depreciations in the local market. This asymmetry can arise from foreign exporters' reaction to exchange rate changes or from factors originating from the local economy. In addition, depending on these factors, an appreciation of the local currency can lead to a higher or lower ERPT than a depreciation. In this section, we provide a brief review of the factors that can cause asymmetric responses in import and consumer prices in the local market.

Pricing Strategies of Foreign Exporters

In imperfectly competitive markets, firms can charge a mark-up over their marginal cost and can adjust this mark-up in response to changes in the exchange rate or marginal cost; thus paving the way for local currency pricing or pricing-to-market (PTM), a term coined by Krugman (1987). The extent of the price adjustment by foreign firms depends on the degree of product substitutability between local and foreign goods and the relative share of foreign firms in the local market (see Dornbusch (1987), Feenstra, Gagnon, and Knetter (1996)). If foreign exporters want to keep market share constant (or increase) in the local market, they can decrease the mark-up when the local currency depreciates. On the other hand, when the local currency appreciates, foreign exporters can retain their mark-ups (or increase mark-up, but less than the extent of change in the exchange rate) and allow the prices in the local currency to fall. Therefore, an appreciation of the local currency can have a different degree of pass-through than a depreciation.

Domestic prices can also behave differently based on the size and duration (temporary vs. permanent) of the exchange rate change and the frequency of price adjustment by foreign exporters. The size asymmetries in the response of foreign exporters can arise from 'menu costs' associated with changing prices. Firms absorb unexpected small and temporary exchange rate fluctuations in their markup and adjust prices in the local market only if the fluctuation is large (see Krugman (1987), Dixit (1989)). Froot and Klemperer (1989) show that the size and direction of the pass-through depend on whether the exchange rate change is perceived to be temporary or permanent. A temporary appreciation of the local currency increases foreign exporters' current profit, relative to the future one, in their own currency. This makes investment in the share of local market less attractive, causing the foreign exporters' profit margin to grow. A permanent appreciation, on the other hand, causes the current and future costs of foreign firms to fall in the local currency. This leads foreign exporters to compete for market share and reduce prices in the local currency. Pass-through will also be higher if foreign exporters adjust prices more frequently in the local market (see Gopinath and Itskhoki (2010)).

The preceding comments imply that asymmetric price adjustments can arise from foreign exporters' ability to adjust their mark-ups when needed. The next question is: do they have a sufficiently large mark-ups? Melitz (2003) develops an open economy dynamic industry model with heterogeneous firms. The model shows that the more productive firms enter the export market while the less productive firms produce only for the domestic market. Bernard, Redding and Schott (2011) extends this analysis to multiple-product multiple-destination firms. Similar to Melitz, the low productivity firms produce only for the domestic market, while the high productivity firms export as well as producing for the domestic market. They also find that a decline in trade costs raise the productivity of the exporting firms by causing them to stop production of their least-successful products. Berman, Martin and Mayer (2012) find that only a small number of highly productive and large firms has a high share in total exports. Heterogeneous demand elasticities and distribution costs of these firms enable them to generate heterogeneous PTM in different destination markets. Mayer, Melitz, and Ottaviano (2014) also find that, in a model of multi-product firms, tougher competition in the export market leads exporting firms to sell only the best performing products in the destination market. Therefore, firms that operate in the export market tend to be highly productive, large in size and sell their best performing products; and thus, enabling them to adjust mark-ups to achieve their desired market share.

Distribution Chain in the Local Market

Foreign exporters do not sell products directly to the consumer. There are costs of advertising, distributing and retail selling which must be paid in the local currency. Burstein, Neves and Rebelo (2003) estimate that such distribution costs represent more than 40 percent of retail price in USA and create a natural wedge between retail prices in different countries. For some products, a local firm can import products directly from the foreign manufacturer and own the distribution chain. In these cases, the local firm can also adjust its distribution margin in response to changes in the exchange rate. Goldberg and Campa (2010) find that a nominal exchange rate depreciation of 1 percent results in a roughly 0.35 percent decrease in the distribution margin. Berman et al. (2012), using firm-level export data for France, find that PTM is observed more for final consumer goods compared to intermediate goods; and for sectors with higher distribution costs.

Binding Quantity Constraints

Binding quantity constraints can arise if foreign firms cannot increase exports when the local currency appreciates. This can be attributed to the production capacity constraints of foreign exporters or trade restrictions in the local market. For example, Japanese automobile manufacturers faced trade restrictions in the form of quotas in the 1980s in the USA. When the USD appreciated against the Yen in the early 1980s, the Japanese manufacturers could not increase exports to the USA since the quota has already been reached. In the presence of such binding constraints, the degree of ERPT to import prices is higher for a depreciation compared to an appreciation of the local currency.

Production Switching

When the local currency depreciates, it can affect the costs of both the local and foreign firms. For this change in the exchange rate, foreign firms may import production inputs from their trading partners at lower cost which in turn lowers their cost of production. In this situation, it is possible for the foreign exporters to lower export prices while keeping the profit margin unchanged. On the other hand, when the local currency appreciates, foreign exporters can switch away from imported inputs. In this case, production costs may not increase but there is scope for foreign exporters to increase their profit margin. Similar production switching can also take place in the local market. In this context, Burstein, Eichenbaum, and Rebelo (2005) claim that low ERPT to consumer prices is partially due to this ‘flight from quality’ to lower quality local substitutes in response to a depreciation of the exchange rate.

Goldberg and Campa (2010) report that imported inputs account for 10 percent to 48 percent of the production cost of tradable goods and 3 percent to 35 percent of the production cost of non-tradable goods for 21 industrialized countries. They find that integrated production (imported goods used as inputs in domestic production) is the dominant channel, compared to direct consumption of imported goods, for explaining the exchange rate or import price pass-through to consumer prices.

Macroeconomic Performance and Policies

Along with the microeconomic factors, macroeconomics factors like exchange rate volatility and inflationary environment also play major roles in pass-through behavior. High volatility in the exchange rate may make foreign exporters and/or local distributors to adjust mark-up instead of adjusting prices frequently, thus reducing pass-through (see Mann (1986), Taylor (2000)). Also, countries with low and stable inflation tend to have low ERPT (see Mishkin and Schmidt-Hebbel (2007), Slavov (2008), Frankel, Parsley and Wei (2012) etc.).

Asymmetry can also arise from the monetary policy pursued in response to exchange rate changes. A depreciation of the local currency can be accompanied by monetary tightening to maintain a low and stable inflation. Empirical evidence from Gagnon and Ihrig (2004) supports this argument. On the other hand, an appreciation of the local currency is supportive to low inflation and will not require a monetary policy response. However, if the central bank targets inflation, it may use expansionary monetary policy to offset the effects of an appreciation.

3.3. Empirical Literature

Theoretical literature has been aware of asymmetric price responses to exchange rate appreciations and depreciations at least since Dornbusch (1987) and Krugman (1987). However, surprisingly, there has been a rather limited empirical literature on the possibility of asymmetric price adjustments even though estimation of the ERPT to prices, using a linear model, is in abundance. Among the studies that do analyze asymmetry in price responses to exchange rate appreciations and depreciations, most use disaggregated industry-level data (firm-level data in some cases) and import prices. There is only a handful of studies that focus on asymmetric ERPT to aggregate price level (such as, consumer prices) though this has great significance for monetary policies. The empirical evidence from this literature can, at best, be described as mixed: some studies confirm the existence of asymmetries while others do not. Even among the studies that support the existence of asymmetries, there is no agreement on the type of asymmetry: some studies find appreciations to cause higher ERPT than depreciations while others report the opposite.

Earlier studies (e.g., Mann (1986) for US exports and imports; and Kadiyali (1997) for US the photographic print film industry) examine the effects of exchange rate movements on prices by dividing the sample period into two subgroups, based on sustained appreciations or depreciations of the exchange rate. The results from these two studies suggest that ERPT to prices is higher under a sustained depreciation of the USD.

Marston (1990) investigates the variations in the ratio of export to domestic prices to changes in the real exchange rate for 17 products in the transport and electrical machinery industries. Their sample period covers 1980 to 1987 and, to accommodate asymmetry, their estimation includes a slope dummy variable of the real exchange rate, which takes the value 1 from February 1985 until the end of the sample period and 0 otherwise. He finds the coefficient of the slope dummy to be statistically significant for 5 products. Athukorala and Menon (1994) follows a methodology similar to that of Marston (1990) to estimate the effects of exchange rate changes on Japanese export prices for total manufacturing along with its 7 sub-categories of two-digit International Standard Industrial Classification (ISIC) industries for the period 1980 to 1992. They find that the intercept and slope dummy is statistically insignificant in all industries; and hence, reject the view of asymmetric ERPT. Following similar specification and allowing slope dummy variable to account for asymmetry, Yang (2007) also finds limited evidence of asymmetry in US aggregate import prices and its 98 disaggregated industries level.

Knetter (1994) examines asymmetries using German and Japanese seven-digit industry exports. He separates exchange rate changes into two variables, one includes the values of exporter's currency appreciations and the other includes those of depreciations. The hypothesis that the slope of these two variables is equal is rejected only for 2 out of 32 industries. Gil-Pareja (2000) follows a specification similar to Knetter (1994) to investigate asymmetric adjustment in export prices of seven European countries with highly disaggregated product categories (8-digits level). The hypothesis that the slope of the variables for exchange rate appreciation and depreciations is equal is rejected for 10 of the 115 source country-product pairs. By comparison, Pollard and Coughlin (2004) find relatively more evidence of asymmetric price adjustments in US import prices for 9 two-digit ISIC manufacturing industries along with the total manufacturing industry.[6] They find that firms in 5 industries respond asymmetrically to appreciations and

[6] They also analyze 20 industries at three-digit level and find asymmetric behaviour in half of the industries.

depreciations, and the direction of asymmetry varies across industries. Moreover, they find that manufacturing as a whole shows no evidence of asymmetry. To capture the size asymmetry, they include two slope dummy variables that separate large and small changes in the exchange rate. They find that pass-through is positively related to the size of the exchange rate change for most industries. In addition, when the slope dummies for both direction and size effects are included, they find that the size effect is more dominant.

Webber (2000) investigates asymmetric ERPT to import prices for 8 countries across the Asia-Pacific region using the cointegration technique. Asymmetry is introduced into the model by decomposing the exchange rate variable into the accumulated sum of the appreciations and depreciations. In 5 of their countries, out of the 7 countries where asymmetry is found, ERPT to import prices is higher when the local currency depreciates.

Peltzman (2000) studies 242 markets for consumer and producer goods to examine asymmetric adjustment in output prices in response to input cost changes.[7] To accommodate asymmetry, he includes a slope dummy variable which takes the value of 1 when the change in input cost is positive and 0 otherwise. In two-thirds of the markets, product prices responds faster to increases in input costs than to decreases. The asymmetric response to the cost shock is substantial and this difference is sustained for at least 5 to 8 months. Although the change in input costs may not be necessarily due to exchange rate fluctuations, the results are important as it provide strong support for asymmetric price adjustment. Frankel et al. (2012) accommodate asymmetry in a similar manner to examine directional and size asymmetries of ERPT to product prices of eight brand commodities in 76 developing countries. They find strong evidence of directional and size asymmetries. They cannot reject the hypothesis that appreciation is not passed through at all, which they interpret as evidence of downward price rigidity.

There are only a handful of studies that examine the asymmetric adjustment of CPI to exchange rate appreciations and depreciations. Mihaljek and Klau (2008) examine directional and size asymmetries in CPI for 14 emerging countries by including intercept dummies for the periods of appreciation and depreciation. If the coefficients of these two dummy variables are significantly different, this will provide support for asymmetric effects. They find mixed results: for some

[7] Constructs a single input cost with data from input-output table.

countries appreciation has a stronger effect, while for some other countries depreciation has a stronger effect; there are also industries for which the difference is not significant. However, results from Razafimahefa (2012) show that, for a panel of sub-Saharan African (SSA) countries, ERPT to CPI is higher when the local currency depreciates. To accommodate asymmetry, he includes a slope dummy to account for depreciations, in the estimating equation along with the exchange rate. Caselli and Roitman (2016) examine asymmetric adjustments in CPI to exchange rate fluctuation for a panel of 28 emerging economies. Asymmetry is introduced into the model through intercept and slope dummies for the depreciation episodes. They find significant evidence of asymmetric price adjustment in the first 8 months after the initial shock.

Delatte and López-Villavicencio (2012) investigate directional asymmetry in ERPT to consumer prices for Germany, Japan, the UK and the USA. They decompose exchange rate into the accumulated sum of the appreciations and depreciations and follow the model of Shin, Yu and Greenwood-Nimmo (2011, 2014). They find evidence of both long-run and short-run asymmetry; and the long-run adjustment of consumer prices can be somewhat different than the short-run one. Their results also show that exchange rate depreciation causes higher pass-through to consumer prices than appreciations and they attribute this to a weak competition structure.

In addition to searching for asymmetry in the price adjustment process, several studies have focused on nonlinear and regime-dependent relations between the exchange rate and consumer prices. Campa, Mínguez and Barriel (2008) use the error correction model and examine nonlinearity of three different forms: non-proportional adjustment, asymmetric adjustment and the existence of threshold below which no adjustment takes place. They introduce nonlinearity through the error correction term (nonlinear adjustment towards long-run equilibrium). They find evidence of three different nonlinearities, with considerable variation across industries and countries, for import prices of 15 EU countries. Using a logistic smooth transition model, Nogueira and Leon-Ledesma (2011) find evidence of nonlinearity in ERPT to CPI for Mexico; and also, pass-through is higher in periods of macroeconomic instability. Bussière (2013) examines nonlinearity and asymmetry in the response of export and import prices to exchange rate changes in the G7 countries. Nonlinearities and/or asymmetries are characterized by augmenting a standard linear model with polynomial terms of the exchange rate, with slope dummy variables and with threshold variables in different specifications. The results provide evidence of nonlinearities and asymmetries with a considerable cross-country variation in terms of direction of asymmetry and magnitude of nonlinearity. Donayre and Panovska (2016), using a Threshold VAR model, find evidence of regime-dependent ERPT to import, producer and consumer prices for Canada and Mexico.

To conclude, while a variety of approaches are employed to analyze the asymmetric responses of prices to exchange rate changes, the empirical evidence is still mixed in this regard. Additionally, such analysis is still lacking for most of our sample of countries. This chapter attempts to fill this lacuna.

3.4. Econometric Methodology

It is clear from the preceding review of the empirical literature that asymmetric price adjustments can be introduced into the estimated model through additional dummy variables or by decomposing the exchange rates into two separate variables that account for episodes of appreciations and depreciations. A standard model with asymmetry can be written as (since most models are estimated in difference form due to non-stationarity of the variables):

$$\Delta y_t = \beta_0 + \sum_{i=1}^{p} \beta_{1,i} \Delta y_{t-i} + \sum_{i=0}^{p} \beta_{2,i} x_{t-i}^{+} + \sum_{i=0}^{p} \beta_{3,i} x_{t-i}^{-} + \varepsilon_t \qquad (1)$$

or, equivalently:

$$\Delta y_t = \beta_0 + \sum_{i=1}^{p} \beta_{1,i} \Delta y_{t-i} + \sum_{i=0}^{p} \beta_{2,i} \Delta x_{t-i} + \sum_{i=0}^{p} \gamma_i \tilde{x}_{t-i} + \varepsilon_t \qquad (2)$$

where $x_t^{+} = I_t^{+} \Delta x_t$, $x_t^{-} = I_t^{-} \Delta x_t$, and $I_t^{+} = \begin{cases} 1 \; if \; \Delta x_t > 0 \\ 0 \; otherwise \end{cases}$ and $I_t^{-} = \begin{cases} 1 \; if \; \Delta x_t < 0 \\ 0 \; otherwise \end{cases}$

In equation (2), $\tilde{x}_t$ can be either x_t^{+} or x_t^{-}. These x_t^{+} and x_t^{-} can be defined in a number of alternative ways. An alternative, often used in the empirical literature, is the use of censored variables. To illustrate the use of censored variables, let $x_t^{c,+} = max(0, x_t - x_t^{*})$ and $x_t^{c,-} = min(0, x_t - x_t^{*})$, where x_t^{*} can be the last period's value, the highest value over a certain period of time (e.g., 1 year) or a certain threshold value. Hamilton (2003) and Kilian and Vigfusson (2011) use the censored variable form to examine asymmetric response of US GDP to oil price changes. A time series variable can also be decomposed as: $x_t = x_0 + x_t^{++} + x_t^{--}$, where $x_t^{++} = \sum_{i=1}^{T} max(0, \Delta x_i)$ and $x_t^{--} = \sum_{i=1}^{T} min(0, \Delta x_i)$. Thus, x_t^{++} and x_t^{--} are partial sum processes that account for the positive and negative growth of the variable, respectively. Webber (2000), Shin et al (2011, 2014) and Delatte and López-Villavicencio (2012), among others, use this partial sum decomposition in their studies.

The null hypothesis of symmetry, in equation (1), can be tested by the two standard Wald test: (i) $\beta_{2,i} = -\beta_{3,i}$, for all $i = 0,1, \dots, p$ or (ii) $\sum_{i=0}^{p} \beta_{2,i} = -\sum_{i=0}^{p} \beta_{3,i}$. Similarly, for equation (2), the hypothesis of symmetry can be tested with a Wald/F test by testing $\gamma_i = 0$ for all $i = 0,1, \dots, p$. The specification in equations (1) and (2) can explain only short-run adjustments as they are estimated in the difference form. Recently, Shin et al. (2011, 2014) proposed a nonlinear autoregressive distributed lag (NARDL) model, which has a dynamic error correction representation, to analyze short- and long-run asymmetry. Within our bivariate framework, the model can be expresses as:

$$\Delta y_t = \beta_0 + \alpha_1 y_{t-1} + \alpha_2 x_{t-1}^{++} + \alpha_3 x_{t-1}^{--} + \sum_{i=1}^{p} \beta_{1,i} \Delta y_{t-i} + \sum_{i=0}^{p} \left(\beta_{2,i} \Delta x_{t-i}^{++} + \beta_{3,i} \Delta x_{t-i}^{--} \right) + \varepsilon_t$$

In this specification, short-run asymmetry can be tested similar to equation (1). Additionally, long-run asymmetry can be tested with a Wald test, $H_0: \alpha_2/\alpha_1 = -\alpha_3/\alpha_1$. Shin et al. (2011, 2014) also computed dynamic multipliers, using the slope parameters, to show the dynamic adjustment from the short run to the long run. Delatte and López-Villavicencio (2012) follow this procedure to examine asymmetric ERPT to consumer prices.

As already mentioned in the introduction, slope-based tests are not informative regarding the extent of asymmetry in impulse responses to structural shocks. Since impulse response functions are nonlinear functions of the slope parameters and the size of the structural shock, the response can show statistically significant asymmetries even when the slope-based test fails to reject the symmetry hypothesis (or vice versa). Some early studies (for example, Bernanke, Gertler, and Watson (1997); Leduc and Sill (2004)) included censored variables into the VAR model and estimated the impulse response functions. However, impulse responses in the VAR models with censored variables are dependent on the history of observations (see, e.g., Gallant, Rossi, and Tauchen (1993), Koop, Pesaran, and Potter (1996)) whereas responses in the linear model are independent of the history of the observations. This point was ignored in those early studies, so that impulse responses were not estimated properly and had asymptotic biases. To correct for this, Kilian and Vigfusson (2011) develop a procedure for computing structural impulse responses from the VAR model, that includes censored variable(s), using Monte Carlo integration over all possible paths of the data. Since this procedure is preferable to its alternatives, this chapter examines asymmetry of impulse responses derived from VAR framework following this procedure.

Based on the review of theoretical and empirical literature, discussed in the previous two sections, it is evident that exchange rate appreciations and depreciations can affect CPI differently. To examine asymmetric price adjustments, we use the following model allowing for exchange rate appreciations and depreciations to have an effect, but to different extents:

$$\Delta x_t = \beta_{10} + \sum_{i=1}^{p} \beta_{11,i} \Delta x_{t-i} + \sum_{i=1}^{p} \beta_{12,i} \Delta y_{t-i} + \varepsilon_{1,t} \tag{3}$$

$$\Delta y_t = \beta_{20} + \sum_{i=0}^{p} \beta_{21,i} \Delta x_{t-i} + \sum_{i=1}^{p} \beta_{22,i} \Delta y_{t-i} + \sum_{i=0}^{p} \gamma_{21,i} \tilde{x}_{t-i} + \varepsilon_{2,t} \tag{4}$$

Let x and y stand for NEER and CPI, respectively. This model can be estimated using standard OLS regressions as the residuals are uncorrelated. An additional advantage, as pointed by Kilian and Vigfusson (2011), is that the dynamic responses can be estimated consistently even if the true adjustment process is symmetric. The only disadvantage of estimating this model is that the estimated parameters are not efficient asymptotically.

The unconditional impulse responses, the average of impulse responses across all histories, to both exchange rate appreciations and depreciations can be calculated from the parameters obtained by estimating equations (3) and (4). Then, a Wald test can be used to test the null hypothesis of symmetry, using:

$$H_0: I_y(h, \delta) = -I_y(h, -\delta) \tag{5}$$

or, $H_0: I_y(h, \delta) + I_y(h, -\delta) = 0$ for $h = 0,1,2, \ldots, H$

This hypothesis tests whether the response of CPI to an appreciation of the exchange rate of size δ is equal to the negative of the response of CPI to a depreciation of the exchange rate of same size at horizon h. This Wald test has an asymptotic $\chi^2_{\mathrm{H}+1}$ distribution. The variance-covariance matrix of the sum of impulse responses can be estimated using bootstrap simulation. Using this test, Kilian and Vigfusson (2011) found very little evidence of asymmetric responses of GDP in the USA to positive and negative oil price shocks. Also, Serletis and Istiak (2016), following this procedure, find that the responses of real GDP to positive and negative money supply shocks are symmetric in the USA.

3.5. Results

We estimate a bivariate VAR model, as in (3) and (4), as the inclusion of additional macroeconomic variables neither affects the econometric points of interest nor is required for consistently estimating the response of CPI to an exchange rate shock (for details, see Kilian and Vigfusson (2011)). The identifying assumption that CPI has no contemporaneous effect on exchange rate is based on the latest empirical literature (see, McCarthy (2007), Faruqee (2006), Ito and Sato (2008) etc.). We use the Census X13 procedure for seasonal adjustment of the quarterly data for both variables. For our baseline estimation, the model is estimated with both variables in first-difference form[8] and 6 lags of the VAR model, i.e., $p = 6$ in equations (3) and (4). The slope dummy variable in equation (4) is defined as $\tilde{x}_t = x_t^- = I_t^- \Delta x_t$, where $I_t^- = \begin{cases} 1 \; if \; \Delta x_t < 0 \\ 0 \; otherwise \end{cases}$. Given the specification of our model $\tilde{x}_t$ can also be expressed in the form of a censored variable as $\tilde{x}_t = x_t^{c,-} = min(0, \Delta x_t)$. We, then, estimate the model by OLS and calculate the cumulated impulse response functions for 12 quarters (i.e., $h = 12$) as we are interested in the response of CPI, rather than their growth rates, to an NEER shock.

The unconditional cumulated impulse responses are calculated by averaging 1000 cumulated impulse responses over 250 histories each. We, first, compute the response of CPI to 1 standard deviation positive NEER shock (an appreciation of the local currency), $I_y(h, \delta)$, and the negative of the responses of CPI to a negative NEER shock (a depreciation of the local currency), $-I_y(h, -\delta)$. We, then, normalize these responses to get the response of CPI to a 1 percent NEER shock. This gives a better idea about the extent of responses and it allows comparison of responses across countries. The unconditional cumulated impulse responses are presented graphically in Figure 3.1, where we also present the impulse responses from a linear model for the sake of comparison.

Figure 3.1 shows that responses from the nonlinear model are quite different than the impulse responses calculated from the linear model, especially for Indonesia, Malaysia, Pakistan and Philippines. For the remaining 8 countries, the responses from the linear model lie between the responses of the CPI to a positive NEER shock and the negative of the responses of the CPI to

[8] Both the variables are integrated of order 1, based on the ADF and DF-GLS unit root tests. We are not reporting the details of the unit root tests in order to conserve space.

a negative NEER shock. Figure 3.1 presents mixed results on whether appreciation or depreciation of the exchange rate are passed-through more to consumer prices. For China, currency appreciation has a stronger effect on CPI while for Indonesia and India currency depreciation has a stronger effect at all horizons. In Bangladesh and Japan, currency appreciation has a stronger effect on short horizons (in short-run) and depreciation has a stronger effect as the horizon increases. For Korea, Pakistan and Thailand, the pattern is just the opposite. Delatte and Lopez-Villavicencio (2012) found that depreciations are more strongly passed-through to consumer prices than are appreciations in the four countries (Japan, Germany, UK and USA) of their study. Our results are different from their findings but are consistent with the results from the disaggregated data discussed in section 3.3. With disaggregated data, for some products (industries), a depreciation is passed-through to prices more while for others the pass-through to prices is higher when the currency appreciates. We also observe large cross-country variations in the responses of CPI to currency appreciations and depreciations. For depreciations, the ERPT to CPI is higher for China, Indonesia, India and Pakistan compared to that for Bangladesh, Japan, Korea, Philippines and Thailand. The variation in CPI responses is somewhat lower across countries when their currency appreciates.

For Malaysia, Philippines, Singapore and Taiwan, the impulse responses of CPI to NEER appreciations and/or NEER depreciations is opposite to what expected theoretically, i.e., the CPI increases when the local currency appreciates and/or decreases in response to a depreciation. In theory, an appreciation of the local currency will reduce the prices of imported goods unless the foreign producers keep prices in the local market constant by increasing their mark-ups. Some of these imported goods are consumed as final goods while others are used as inputs by local firms. If imported goods become relatively cheaper, the cost of local productions will also decrease. Also, an appreciation of local currency deteriorates net exports. This results in an increase in the interest rate (decrease in the domestic money supply) unless the monetary authority intervenes. Therefore, an appreciation reduces consumer prices directly since imported goods become cheaper and indirectly by lowering the cost of domestic production and increasing the interest rate. A currency depreciation has the opposite effect.

For Philippines, the CPI increases in response to both appreciations and depreciations. This can be due to the country's inflation targeting policy since 2002 as it may use an expansionary monetary policy to offset the downward pressure on price level, arising from the currency appreciations. However, for Malaysia, Singapore and Taiwan, the CPI increases when the exchange rate appreciates and decreases in response to depreciation.[9] This can be due to misspecification of the model (for example, omitted variable bias) or underlying macroeconomic factors. With the linear model, CPI increases in response to currency appreciation only for Singapore and Taiwan. Using linear models, Choudhri and Hakura (2006) and Ca'Zorzi, Hahn and Sánchez (2007) also found similar evidence for Singapore. Also, the omitted variable bias is less likely to change the sign of responses in all horizons. Thus, we suspect that this pattern of responses is due to the macroeconomic policies of these three countries. Although Malaysia, Singapore and Taiwan do not officially target inflation, their inflation rates are very low and stable in our period of investigation (see Figure 3.2). A de facto classification by the IMF of the exchange rate regimes is presented in Appendix 3.1. Although the classification of the exchange rate regime by the IMF is not available for Taiwan, the country is mostly under a managed float as evident from the website of the Central Bank of the Republic of China (Taiwan).[10] Malaysia and Singapore had fixed peg arrangements and/or managed float during this period. This implies active intervention from policymakers in the foreign exchange market. These three countries also have sizable current account surplus over most of the time period under investigation (see Figure 3.2).[11] Hence, these three countries do not have balance of payments difficulties, which allowed them to use fiscal and monetary policies more freely to achieve their desired target in the foreign exchange market. Other than these three countries, Indonesia, China, Japan, Korea and Thailand also had current account surpluses in most of the period. However, the extent of the surplus in these countries is relatively much lower.

We now use the Wald test to examine the hypothesis of symmetric impulse responses, as presented in equation (5). The p-values of the Wald test are reported in Table 3.1. To perform this test, we use the responses of CPI to 1 standard deviation appreciation of NEER and to 1 standard deviation depreciation of NEER. Based on this test, we find evidence of asymmetry in half of the countries under investigation in this chapter: China, Indonesia, Korea, Malaysia, Pakistan and Philippines. Generally, more evidence of asymmetry is observed in horizons 3-9 compared to

[9] In Taiwan, CPI starts to increase after a year or so when exchange rate depreciates.
[10] http://www.cbc.gov.tw/ct.asp?xItem=856&CtNode=480&mp=2
[11] These three countries also had balance of payments surpluses over most of the period. However, we do not plot this into Figure 3.2 to keep the graph clearer.

horizons 0-2 and 10-12. From Figure 3.1, we can see that the immediate response of CPI to exchange rate appreciations and depreciations are very low for most countries. Also, over longer horizons, the responses tend to be more symmetric for most countries. The evidence of asymmetry seems to be independent of the stage of economic development and the size of the economy.

3.6. Robustness

3.6.1. Different Number of Lags

The baseline model in the previous section is estimated with 6 lags. We, now, examine sensitivity of our results to different lag order of the VAR model (i.e., $p = 4$, 5, 7 and 8) while keeping other specification of the baseline estimation unchanged. The impulse response functions with these different lags are mostly similar to the responses with 6 lags. Hence, we only report the results obtained from Wald test on the null hypothesis in (5). Further, the test results with lags 4 and 5 are similar so that Table 3.2 reports only the p-values of the test using the VAR with 4 lags. The VAR model with 7 lags also produces similar test results to the one with 6 lags so that we do not report its results separately in order to conserve space. Table 3.3 presents the test results derived from the VAR model with 8 lags. With the VAR with 4 lags, the evidence of asymmetry is somewhat lower compared to the baseline results. There is evidence of asymmetry for 5 countries (responses of Korea are no longer asymmetric) and with relatively fewer number of horizons. With 8 lags in the VAR model, we get evidence of asymmetry in 7 countries (the additional country is Singapore) and also at more horizons.

3.6.2. Size of the Shock

In this nonlinear model, the response of CPI to NEER shock can depend on the size of the shock. Hence, we examine asymmetry in the presence of a larger shock, a 2 standard deviation change in NEER, while keeping the other specification of the baseline case unchanged. Again, we examine the null hypothesis in (5) and the results of the test are reported in Table 3.4. The evidence of asymmetry is almost similar to the baseline case. The only difference is that there is a weak evidence (p-value lies between 0.05 to 0.10 in only two horizons) of asymmetry for Bangladesh while there is a weak evidence of asymmetry for Taiwan in the baseline case. Since our estimated model does not have any component that is dependent on the size of exchange rate change, this is a test of sensitivity to the size of the shock and is not a test for size asymmetries.

3.6.3. A Different Measure of the Censored Variable

We, now, replace $\tilde{x}_t = x_t^{c,-} = min(0, \Delta x_t)$ in (4) with $\tilde{x}_t = x_t^{c,+} = max(0, \Delta x_t)$ while keeping the other specifications of the baseline case the same. Theoretically, this change should not alter our results. However, due to the nonlinear nature of the model, the dependence of impulse responses on histories of the observation and the simulation-based construction of the variance-covariance matrix of the impulse responses for the Wald test, there can be some variations in the results. The objective of replacing $x_t^{c,-}$ with $x_t^{c,+}$ is to examine whether these variations are statistically significant. Table 3.5 presents the results with $x_t^{c,+}$. The results are similar to the baseline case, though with the difference that asymmetry has been observed in fewer horizon for some countries. We examine further whether the responses, derived using different measures of the censored variable, are equal for both positive and negative shocks. The hypothesis that the responses are equal is rejected for only Indonesia and Korea at a few horizons for positive (appreciation) shock, as shown in Table 3.6. With negative (depreciation) shock, the null hypothesis is rejected only for Korea and in fewer horizons. These results are not reported in Table format to conserve space. Also, when we examine the responses for Indonesia and Korea, where the hypothesis of equality of responses under different specifications of the censored variable is rejected, the difference is not noticeable at two decimal points. Hence, our baseline results are robust to the specification of the censored variable.

3.6.4. Results from the Slope-based Test

We estimate equation (2) and test the null hypothesis that $\gamma_i = 0$ for all $i = 0,1, \ldots, p$. A rejection of the null will provide evidence against symmetry. A standard F-test is used to test the null and the p-values of the test are reported in Table 3.7. We examine our results with different lags and different measures of the censored variable. Results with $x_t^{c,-}$ and $x_t^{c,+}$ as censored variable produce identical results. Hence, we only report results with $x_t^{c,+}$ as the censored variable and with different lags in estimating (2). The null hypothesis of symmetry is rejected for Bangladesh, China, Indonesia, Malaysia, Pakistan, Pakistan and Thailand. This result is quite different from the impulse response based tests since there is additional evidence of asymmetry for Bangladesh and Thailand while there is no evidence of asymmetry for Korea and Philippines. This result is consistent with the claim in Section 3.4 (econometric methodology) that the responses can show statistically significant asymmetries even when the slope-based test fails to reject the symmetry hypothesis (or vice versa).

3.6.5. Bivariate Model with CPI and USD-based Exchange Rate

For most of our countries examined in this chapter, foreign transactions take place in USD, they hold foreign exchange reserve in USD, a high portion of their assets and liabilities are denominated in USD and, in some of the countries, the local currency was pegged against the USD for a considerable period of time. Hence, it is important to examine the response of CPI to USD-based nominal exchange rate appreciations and depreciations. We estimate a bivariate model, as in equations (3) and (4), with seasonally adjusted data, 6 lags, the censored variable $\tilde{x}_t = x_t^{c,-} = min(0, \Delta x_t)$ and 1 standard deviation exchange rate shock. With these specifications, we examine the null hypothesis in (5) using the Wald test and found evidence of asymmetry in 9 out of 12 countries. For China, Indonesia, India, Korea, Malaysia and Philippines, there is evidence of asymmetric responses at almost all horizons. On the contrary, the responses are asymmetric only at a few horizons for Pakistan, Singapore and Thailand. We also examine the null hypothesis in (5) with impulse responses derived from the model with different number of lags, size of the shocks and measure of censored variable. However, the results are mostly similar and we do not report them to conserve space.

Our next concern is the extent of the impulse responses across countries. Similar to the baseline case, we calculate the responses of CPI to 1 standard deviation positive and negative exchange rate shocks. The responses are then normalized to represent the response of CPI to 1 percent appreciation and depreciation of the local currency. The results are reported in Figure 3.3, which also plots the responses from the linear model under the same specifications without the censored variable. We notice that for most of the countries the impulse responses have signs opposite to what was expected theoretically. The responses have the expected signs for appreciations and depreciations only for Indonesia, Japan, Korea and Pakistan (for Pakistan, when the exchange rate depreciates, the CPI starts decreasing after 6 quarters). For the rest of the countries, the CPI increases when currency appreciates and/or decreases when the currency depreciates against the USD. This implies that policymakers respond more strongly to changes in their local currency against the USD than to any other currency. The discussion on such cases with

NEER is also applicable here. Again, the responses of fiscal and monetary policies to USD-based exchange rate appreciations and depreciations are worth further examination.

3.7. Conclusions

To sum up, in this chapter, we have examined the response of consumer prices to exchange rate appreciations and depreciations. All the previous studies on this topic use slope-based tests to examine the degree of asymmetry. While this test is useful, it does not provide insights on the responses of prices to exchange rate shocks. Also, the impulse responses, derived from a structural VAR model, of prices to exchange rate appreciations and depreciations can be asymmetric, even when the slope-based test provide evidences against asymmetry. Therefore, impulse response based test are more informative about the asymmetric adjustments of prices to exchange rate changes. Hence, in this chapter, we use an impulse response based test, proposed by Kilian and Vigfusson (2011), to examine whether the price responses are asymmetric to exchange rate appreciations and depreciations. This is the first paper to examine this topic using this impulse response based test.

We found evidence of asymmetric responses of CPI to NEER appreciations and depreciations in half of the countries under investigation. The results for the baseline case is robust to changes in model specifications. Unlike Delatte and Lopez-Villavicencio (2012), there is no clear evidence on whether the pass-through is higher when the currency depreciates. The evidence on asymmetry is sensitive to the measure of exchange rate. There is greater evidence of asymmetric responses with USD-based exchange rate compare to the trade-weighted exchange rate, NEER. Also, with the USD-based exchange rate there are more instances where consumer prices increase when the local currency appreciates and/or decreases when it depreciates. We predict that this theoretically unexpected responses may arise from the policy intervention in the foreign exchange market. Further, the policy response is stronger when the exchange rate changes against the USD. We intend to explore this channel in future research. Further, this chapter limited the analysis only to directional asymmetry. An examination for size asymmetry will provide additional insights on this topic and we also leave such examination for future research.

Table 3.1: p-values of the Wald Test for $\boldsymbol{H_0: I_y(h,\delta) = -I_y(h,-\delta)}$

(Bivariate VAR with NEER and CPI, 6 lags, $\tilde{x}_t = x_t^{c,-} = min(0, \Delta x_t)$, $I_y(.)$ = Response of CPI, h = Horizon, δ = 1 Standard Deviation Exchange Rate Shock)

Horizon	0	1	2	3	4	5	6	7	8	9	10	11	12
BGD	0.28	0.19	0.12	0.10	0.17	0.25	0.35	0.44	0.53	0.54	0.63	0.71	0.77
CHN	0.52	**0.00**	**0.01**	**0.02**	**0.02**	**0.02**	**0.03**	**0.05**	**0.07**	**0.09**	0.13	0.17	0.22
IDN	**0.05**	0.13	**0.01**	**0.02**	**0.03**	**0.06**	**0.09**	0.14	0.20	0.20	0.34	0.41	0.49
IND	0.92	0.99	0.84	0.68	0.25	0.15	0.22	0.29	0.38	0.46	0.55	0.64	0.71
JPN	0.18	0.11	0.17	0.22	0.17	0.20	0.20	0.28	0.35	0.43	0.47	0.52	0.60
KOR	0.44	0.22	0.19	**0.04**	**0.05**	**0.08**	**0.02**	**0.04**	**0.04**	**0.05**	**0.07**	0.10	**0.06**
MYS	0.54	**0.01**	**0.03**	**0.02**	**0.04**	**0.03**	**0.03**	**0.03**	**0.05**	**0.07**	0.10	0.11	0.15
PAK	**0.09**	0.25	0.33	**0.01**	**0.03**	**0.04**	**0.07**	**0.10**	0.14	0.20	0.23	0.28	0.33
PHL	0.15	**0.01**	**0.04**	**0.00**	**0.00**	**0.01**	**0.00**	**0.01**	**0.01**	**0.02**	**0.02**	**0.03**	**0.05**
SGP	0.21	0.30	0.48	0.57	0.48	0.36	0.28	0.36	0.44	0.53	0.60	0.68	0.75
THA	0.30	0.59	0.59	0.54	0.12	0.14	0.21	0.26	0.35	0.41	0.50	0.59	0.67
TWN	0.36	**0.08**	**0.08**	0.15	0.14	0.20	0.28	0.37	0.47	0.55	0.64	0.69	0.76

Table 3.2: p-values of the Wald Test for $\boldsymbol{H_0: I_y(h,\delta) = -I_y(h,-\delta)}$

(Bivariate VAR with NEER and CPI, **4 lags**, $\tilde{x}_t = x_t^{c,-} = min(0, \Delta x_t)$, $I_y(.)$ = Response of CPI, h = Horizon, δ = 1 Standard Deviation Exchange Rate Shock)

Horizon	0	1	2	3	4	5	6	7	8	9	10	11	12
BGD	0.48	0.35	0.23	0.32	0.45	0.50	0.68	0.78	0.85	0.88	0.92	0.95	0.96
CHN	0.25	**0.00**	**0.01**	**0.01**	**0.03**	**0.03**	**0.05**	**0.08**	0.11	0.16	0.21	0.22	0.28
IDN	**0.03**	0.11	**0.02**	**0.01**	**0.02**	**0.03**	**0.06**	**0.07**	0.10	0.15	0.20	0.24	0.30
IND	0.87	0.98	0.96	0.84	0.51	0.61	0.71	0.79	0.86	0.91	0.94	0.96	0.97
JPN	0.29	0.13	0.18	0.29	0.41	0.53	0.65	0.71	0.79	0.83	0.86	0.89	0.93
KOR	0.64	0.29	0.20	0.26	0.30	0.39	0.46	0.54	0.63	0.72	0.74	0.80	0.85
MYS	0.25	**0.05**	0.11	**0.03**	**0.04**	**0.07**	0.11	0.15	0.22	0.25	0.27	0.33	0.41
PAK	0.24	0.44	0.47	0.10	0.15	0.23	0.30	0.40	0.50	0.59	0.68	0.73	0.77
PHL	0.18	**0.04**	**0.09**	**0.02**	**0.03**	**0.05**	**0.07**	0.11	0.14	0.19	0.25	0.32	0.38
SGP	0.35	0.51	0.62	0.68	0.57	0.68	0.78	0.83	0.85	0.89	0.92	0.95	0.97
THA	0.24	0.50	0.43	0.22	**0.07**	0.12	0.16	0.23	0.31	0.39	0.43	0.51	0.57
TWN	0.29	0.11	0.12	0.19	0.29	0.40	0.52	0.63	0.71	0.79	0.85	0.89	0.93

Table 3.3: p-values of the Wald Test for H_0: $I_y(h, \delta) = -I_y(h, -\delta)$

(Bivariate VAR with NEER and CPI, **8 lags**, $\tilde{x}_t = x_t^{c,-} = min(0, \Delta x_t)$, $I_y(.)$ = Response of CPI, h = Horizon, δ = 1 Standard Deviation Exchange Rate Shock)

Horizon	0	1	2	3	4	5	6	7	8	9	10	11	12
BGD	0.17	0.27	0.20	0.12	0.20	0.30	0.28	0.35	0.40	0.49	0.57	0.66	0.73
CHN	0.87	**0.00**	**0.01**	**0.02**	**0.02**	**0.01**	**0.02**	**0.01**	**0.01**	**0.01**	**0.01**	**0.02**	**0.03**
IDN	**0.01**	**0.05**	**0.00**	**0.00**	**0.01**	**0.01**	**0.01**	**0.02**	**0.03**	**0.04**	**0.06**	**0.06**	**0.08**
IND	0.81	0.90	0.74	0.66	0.57	0.32	0.42	0.37	0.43	0.49	0.58	0.67	0.72
JPN	0.24	0.14	0.20	0.30	0.39	0.41	0.52	0.63	0.72	0.79	0.85	0.89	0.92
KOR	0.38	0.46	0.11	**0.05**	**0.05**	**0.06**	**0.08**	0.11	0.15	0.20	0.27	0.34	0.36
MYS	0.85	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.02**	**0.03**
PAK	**0.00**	**0.01**	**0.02**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**
PHL	**0.02**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**
SGP	0.14	**0.05**	0.10	0.11	**0.07**	**0.04**	**0.06**	**0.08**	**0.07**	0.10	0.14	0.19	0.25
THA	0.26	0.50	0.37	0.38	0.15	0.19	0.26	0.35	0.45	0.51	0.60	0.68	0.76
TWN	0.25	0.10	0.16	0.25	0.38	0.34	0.44	0.45	0.49	0.58	0.66	0.74	0.80

Table 3.4: p-values of the Wald Test for H_0: $I_y(h, \delta) = -I_y(h, -\delta)$

(Bivariate VAR with NEER and CPI, 6 lags, $\tilde{x}_t = x_t^{c,-} = min(0, \Delta x_t)$, $I_y(.)$ = Response of CPI, h = Horizon, **δ = 2 Standard Deviation Exchange Rate Shock**)

Horizon	0	1	2	3	4	5	6	7	8	9	10	11	12
BGD	0.28	0.17	**0.09**	**0.06**	0.11	0.18	0.26	0.32	0.40	0.43	0.52	0.60	0.67
CHN	0.52	**0.00**	**0.00**	**0.01**	**0.01**	**0.00**	**0.01**	**0.01**	**0.02**	**0.03**	**0.05**	**0.07**	0.10
IDN	**0.06**	0.16	**0.02**	**0.02**	**0.04**	**0.07**	0.11	0.15	0.21	0.28	0.36	0.43	0.51
IND	0.93	0.99	0.86	0.70	0.27	0.17	0.24	0.32	0.41	0.50	0.59	0.67	0.74
JPN	0.22	0.15	0.21	0.29	0.24	0.27	0.27	0.36	0.42	0.52	0.55	0.61	0.69
KOR	0.51	0.31	0.28	**0.07**	**0.09**	0.12	**0.03**	**0.05**	**0.06**	**0.09**	0.11	0.15	0.10
MYS	0.61	**0.04**	**0.08**	**0.06**	**0.08**	**0.06**	**0.06**	**0.06**	**0.09**	0.11	0.15	0.18	0.23
PAK	**0.08**	0.21	0.30	**0.00**	**0.01**	**0.01**	**0.02**	**0.03**	**0.04**	**0.07**	**0.07**	0.10	0.12
PHL	0.20	**0.03**	**0.06**	**0.00**	**0.01**	**0.01**	**0.01**	**0.02**	**0.03**	**0.04**	**0.05**	**0.07**	0.10
SGP	0.25	0.33	0.52	0.62	0.55	0.41	0.33	0.41	0.49	0.58	0.65	0.73	0.79
THA	0.35	0.64	0.67	0.61	0.14	0.17	0.24	0.32	0.41	0.49	0.58	0.66	0.74
TWN	0.40	0.13	0.13	0.22	0.22	0.31	0.41	0.50	0.60	0.69	0.77	0.82	0.87

Table 3.5: p-values of the Wald Test for $\boldsymbol{H_0: I_y(h,\delta) = -I_y(h,-\delta)}$

(Bivariate VAR with NEER and CPI, 6 lags, $\tilde{\boldsymbol{x}}_{\boldsymbol{t}} = \boldsymbol{x}_{\boldsymbol{t}}^{\boldsymbol{c,+}} = \boldsymbol{max}(\boldsymbol{0}, \Delta \boldsymbol{x}_{\boldsymbol{t}})$, $I_y(.)$ = Response of CPI, h = Horizon, δ = 1 Standard Deviation Exchange Rate Shock)

Horizon	0	1	2	3	4	5	6	7	8	9	10	11	12
BGD	0.25	0.23	0.17	0.10	0.15	0.21	0.25	0.34	0.41	0.47	0.55	0.62	0.69
CHN	0.56	**0.00**	**0.01**	**0.02**	**0.02**	**0.03**	**0.05**	**0.07**	0.11	0.14	0.16	0.22	0.27
IDN	**0.04**	0.11	**0.04**	**0.06**	**0.08**	0.14	0.14	0.17	0.23	0.30	0.38	0.42	0.49
IND	0.87	0.98	0.96	0.84	0.51	0.61	0.71	0.79	0.86	0.91	0.94	0.96	0.97
JPN	0.21	**0.09**	0.11	0.17	0.24	0.22	0.28	0.38	0.47	0.55	0.61	0.69	0.75
KOR	0.46	0.27	0.13	**0.06**	**0.03**	**0.06**	**0.01**	**0.02**	**0.03**	**0.05**	**0.08**	0.11	0.10
MYS	0.56	**0.02**	**0.04**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.02**	**0.03**
PAK	**0.08**	0.21	0.31	**0.06**	**0.09**	0.10	0.15	0.22	0.29	0.36	0.45	0.53	0.54
PHL	0.16	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.01**	**0.01**	**0.01**	**0.02**	**0.02**	**0.03**
SGP	0.28	0.32	0.51	0.62	0.48	0.49	0.52	0.63	0.72	0.79	0.85	0.90	0.94
THA	0.24	0.50	0.49	0.38	0.10	0.15	0.21	0.26	0.32	0.40	0.49	0.55	0.57
TWN	0.30	0.15	0.22	0.35	0.48	0.55	0.60	0.68	0.77	0.83	0.88	0.92	0.95

Table 3.6: p-values of the Wald Test for $\boldsymbol{H_0: I_y(h,\delta: x_t^{c,-}) = I_y(h,\delta: x_t^{c,+})}$

(Bivariate VAR with NEER and CPI, 6 lags, $\tilde{\boldsymbol{x}}_{\boldsymbol{t}} = \boldsymbol{x}_{\boldsymbol{t}}^{\boldsymbol{c,-}} = \boldsymbol{min}(\boldsymbol{0}, \Delta \boldsymbol{x}_{\boldsymbol{t}})$, $\tilde{\boldsymbol{x}}_{\boldsymbol{t}} = \boldsymbol{x}_{\boldsymbol{t}}^{\boldsymbol{c,+}} = \boldsymbol{max}(\boldsymbol{0}, \Delta \boldsymbol{x}_{\boldsymbol{t}})$ $I_y(.)$ = Response of CPI, h = Horizon, δ = 1 Standard Deviation Exchange Rate Shock)

Horizon	0	1	2	3	4	5	6	7	8	9	10	11	12
BGD	0.53	0.67	0.85	0.94	0.96	0.95	0.96	0.98	0.99	0.99	1.00	1.00	1.00
CHN	0.71	0.93	0.98	0.86	0.91	0.86	0.92	0.95	0.98	0.99	0.99	1.00	1.00
IDN	**0.08**	0.12	**0.06**	**0.02**	**0.02**	**0.04**	**0.06**	**0.09**	0.13	0.10	0.15	0.19	0.24
IND	0.56	0.45	0.56	0.70	0.68	0.79	0.87	0.90	0.87	0.91	0.95	0.97	0.98
JPN	0.94	0.92	0.96	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
KOR	**0.02**	**0.05**	**0.08**	0.14	0.22	0.32	0.42	0.52	0.61	0.68	0.70	0.77	0.80
MYS	0.66	0.89	0.85	0.89	0.91	0.95	0.97	0.98	0.99	1.00	1.00	1.00	1.00
PAK	0.34	0.46	0.48	0.57	0.64	0.71	0.83	0.89	0.93	0.96	0.98	0.99	0.99
PHL	0.51	0.54	0.37	0.38	0.44	0.50	0.37	0.48	0.57	0.66	0.70	0.74	0.80
SGP	0.82	0.88	0.94	0.97	0.90	0.81	0.80	0.87	0.91	0.94	0.96	0.98	0.99
THA	0.78	0.88	0.86	0.62	0.67	0.76	0.80	0.83	0.89	0.93	0.96	0.98	0.99
TWN	0.20	0.33	0.53	0.57	0.69	0.71	0.79	0.86	0.91	0.94	0.96	0.98	0.99

Table 3.7: p-values of the F-Test of $H_0: \gamma_i = 0$ for all $i = 0, 1, \dots, p$., in Equation (2)

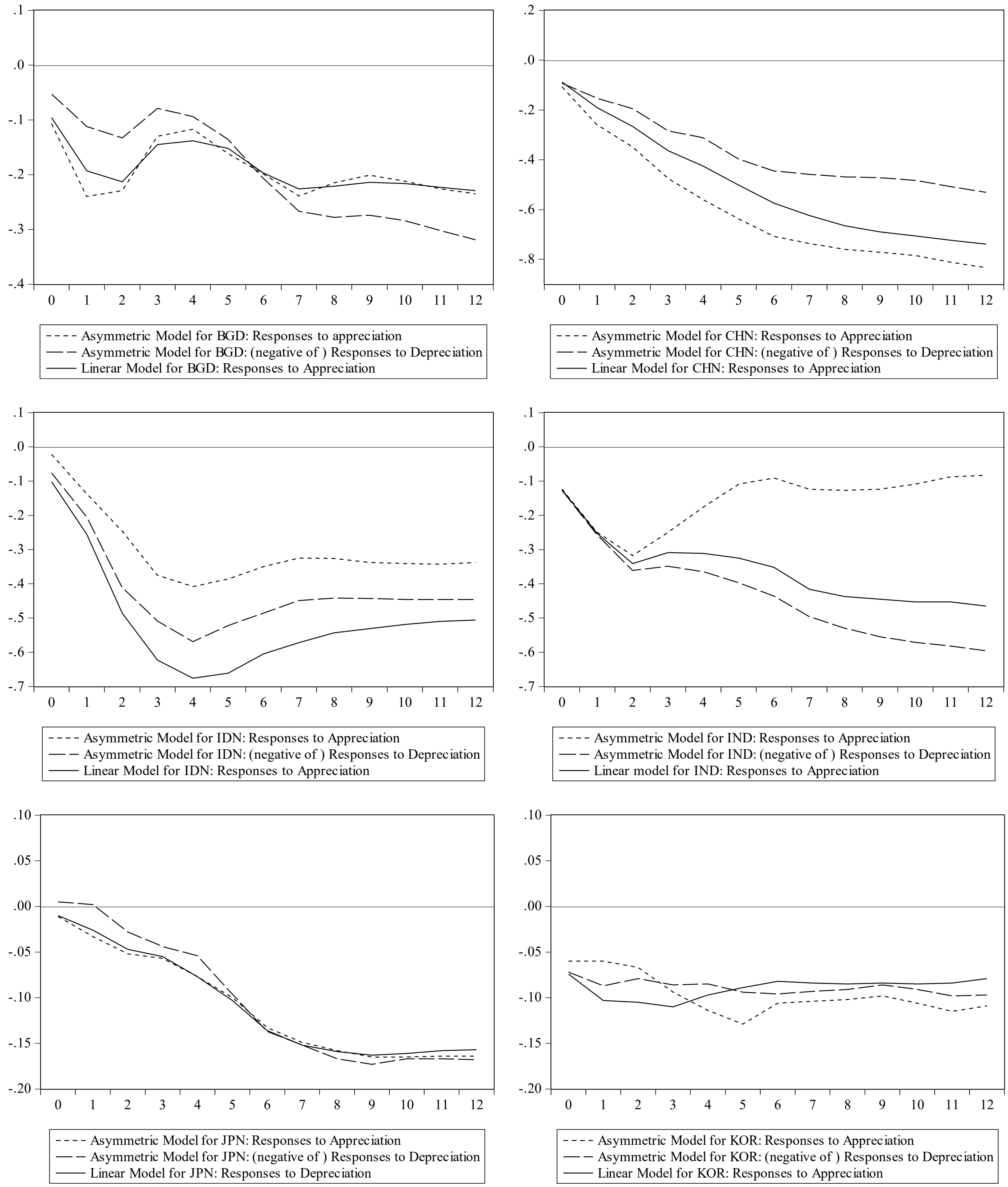
	p = 4	*p = 5*	*p = 6*	*p = 7*	*p = 8*
BGD	**0.07**	**0.04**	**0.05**	0.16	0.16
CHN	**0.02**	**0.01**	**0.00**	**0.00**	**0.00**
IDN	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**
IND	0.42	0.33	0.39	0.58	0.64
JPN	0.16	0.12	0.11	0.20	0.25
KOR	0.35	0.37	0.13	0.66	0.74
MYS	**0.04**	**0.06**	**0.06**	**0.04**	**0.04**
PAK	0.17	**0.05**	**0.00**	**0.02**	**0.07**
PHL	0.11	0.25	0.22	0.11	0.04
SGP	0.75	0.93	0.92	0.61	0.41
THA	**0.00**	**0.00**	**0.00**	**0.01**	**0.02**
TWN	0.41	0.49	0.61	0.72	0.68

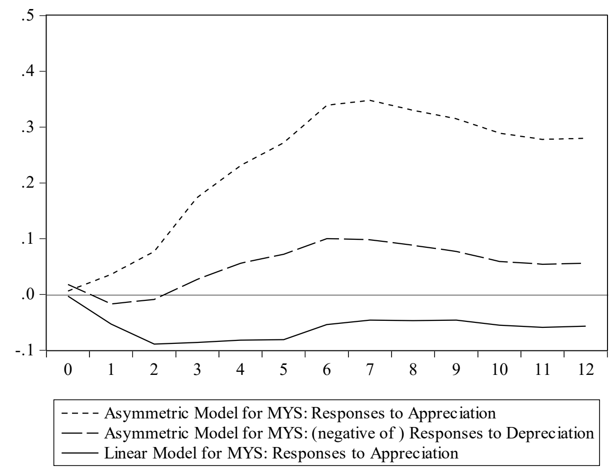
Table 3.8: p-values of the Wald Test for $H_0: I_y(h, \delta) = -I_y(h, -\delta)$

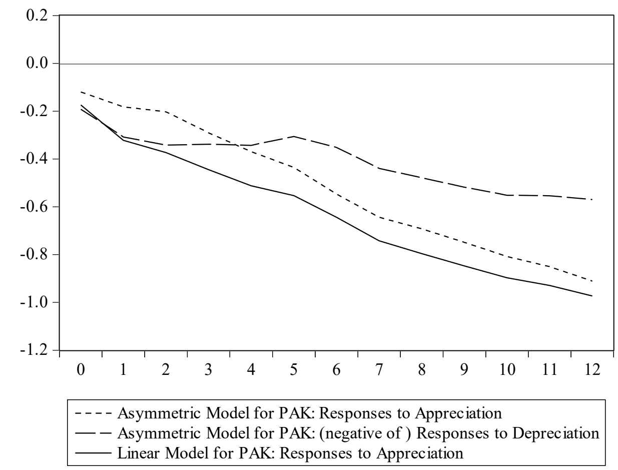
(Bivariate VAR with USD-based Nominal Exchange Rate and CPI, 6 lags, $\tilde{x}_t = x_t^{c,-} = min(0, \Delta x_t)$, $I_y(.)$ = Response of CPI, h = Horizon, δ = 1 Standard Deviation Exchange Rate Shock)

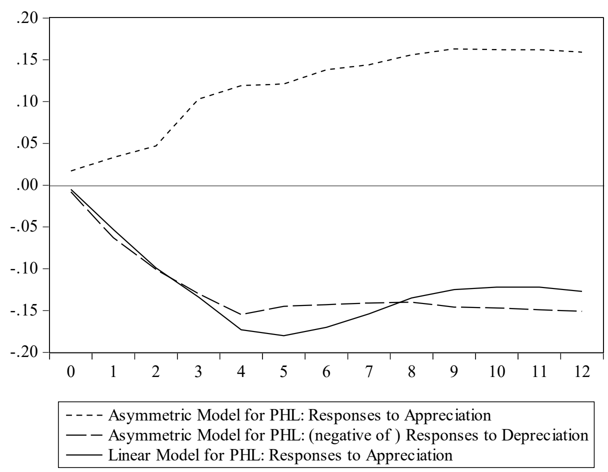
Horizon	0	1	2	3	4	5	6	7	8	9	10	11	12
BGD	0.15	0.11	0.16	0.26	0.23	0.20	0.15	0.21	0.29	0.27	0.35	0.36	0.39
CHN	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.01**	**0.01**	**0.01**	**0.01**	**0.00**	**0.00**	**0.00**	**0.00**
IDN	0.38	0.38	**0.00**	**0.01**	**0.01**	**0.01**	**0.01**	**0.01**	**0.02**	**0.03**	**0.04**	**0.06**	**0.09**
IND	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.01**	**0.01**	**0.02**	**0.02**
JPN	0.26	0.27	0.34	0.48	0.62	0.64	0.65	0.74	0.82	0.87	0.91	0.94	0.95
KOR	**0.07**	**0.04**	**0.02**	**0.03**	**0.06**	**0.03**	**0.01**	**0.01**	**0.02**	**0.03**	**0.04**	**0.04**	**0.06**
MYS	0.96	0.10	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**
PAK	**0.05**	**0.08**	0.10	0.15	0.23	0.31	0.32	0.32	0.39	0.48	0.57	0.65	0.72
PHL	**0.02**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**
SGP	0.13	**0.02**	**0.05**	**0.08**	0.13	0.14	**0.07**	0.10	0.15	0.20	0.27	0.33	0.34
THA	**0.00**	**0.01**	**0.02**	**0.04**	**0.06**	**0.09**	0.12	0.18	0.23	0.30	0.34	0.41	0.47
TWN	0.86	0.96	0.86	0.83	0.36	0.35	0.04	0.06	0.09	0.12	0.14	0.18	0.23

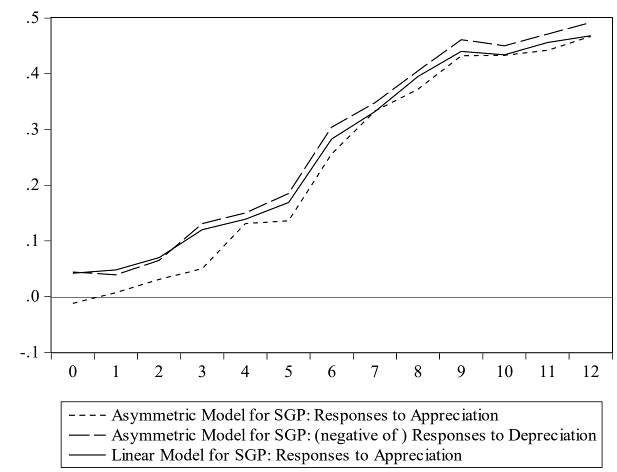
Figure 3.1: The Response of CPI to 1% Change in NEER

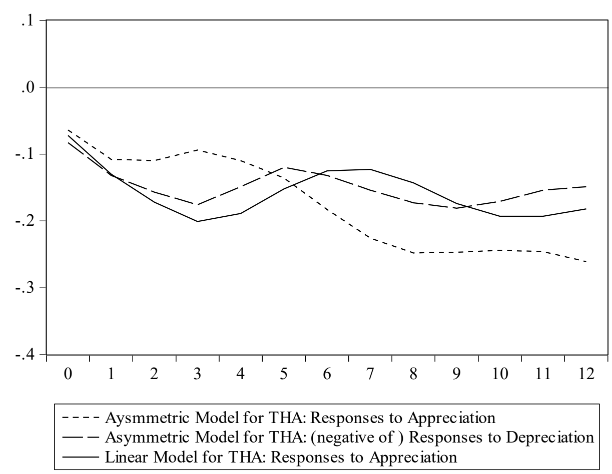



Asymmetric Model for MYS: Responses to Appreciation
Asymmetric Model for MYS: (negative of) Responses to Depreciation
Linear Model for MYS: Responses to Appreciation


Asymmetric Model for PAK: Responses to Appreciation
Asymmetric Model for PAK: (negative of) Responses to Depreciation
Linear Model for PAK: Responses to Appreciation


Asymmetric Model for PHL: Responses to Appreciation
Asymmetric Model for PHL: (negative of) Responses to Depreciation
Linear Model for PHL: Responses to Appreciation


Asymmetric Model for SGP: Responses to Appreciation
Asymmetric Model for SGP: (negative of) Responses to Depreciation
Linear Model for SGP: Responses to Appreciation


Aysmmetric Model for THA: Responses to Appreciation
Asymmetric Model for THA: (negative of) Responses to Depreciation
Linear Model for THA: Responses to Appreciation

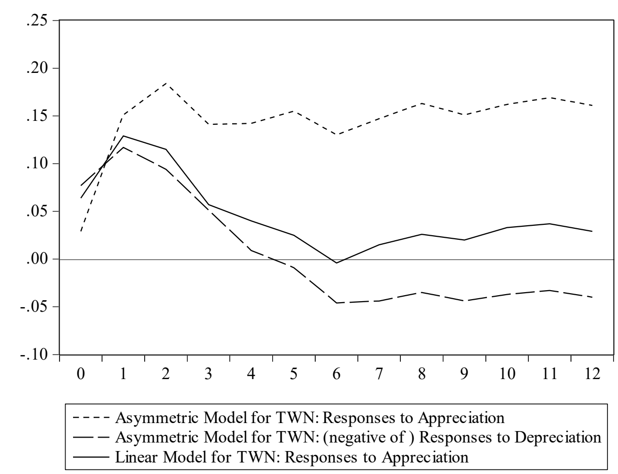
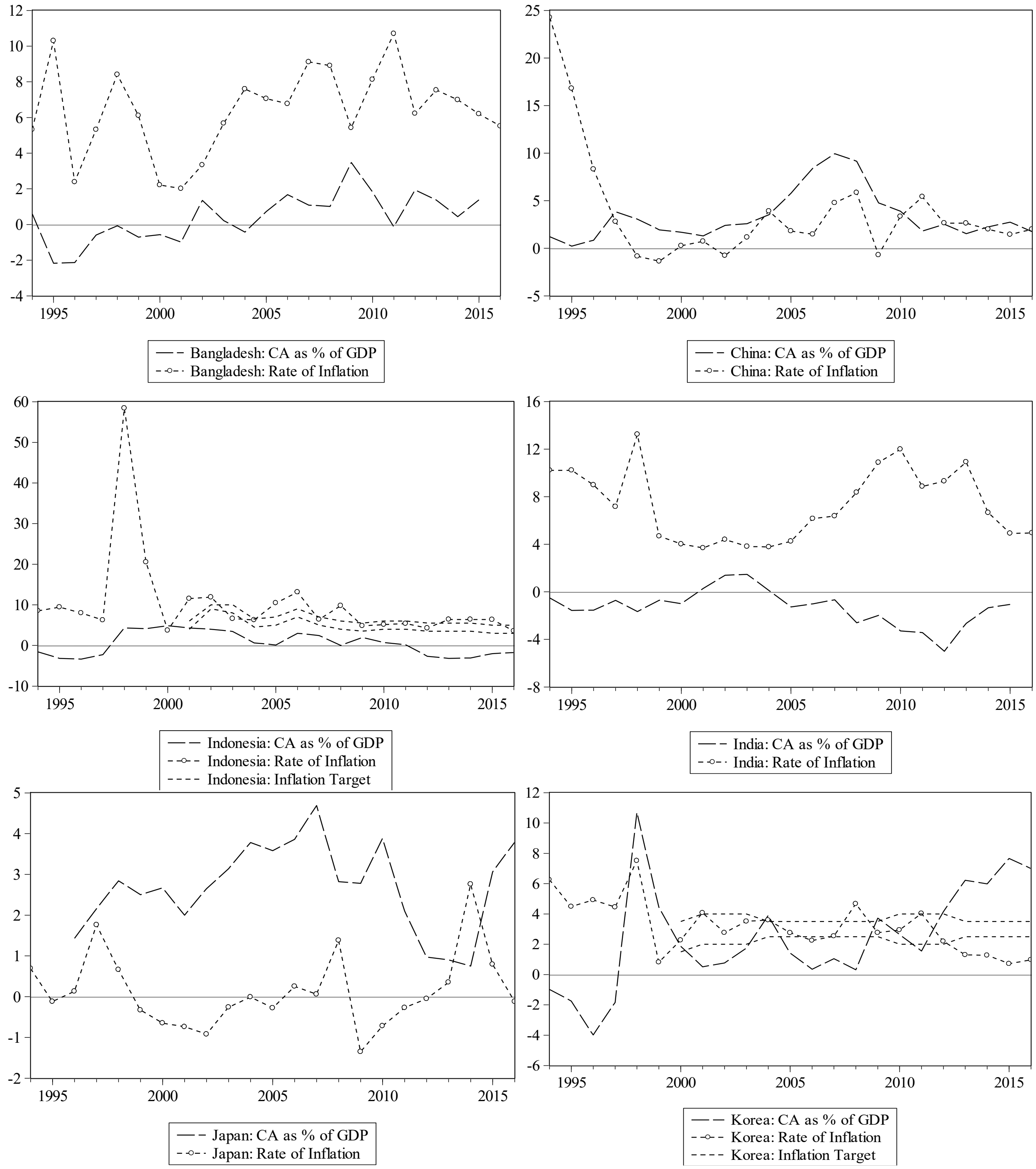

Asymmetric Model for TWN: Responses to Appreciation
Asymmetric Model for TWN: (negative of) Responses to Depreciation
Linear Model for TWN: Responses to Appreciation

Figure 3.2: Rate of Inflation and Current Account (CA) Balance as % of GDP



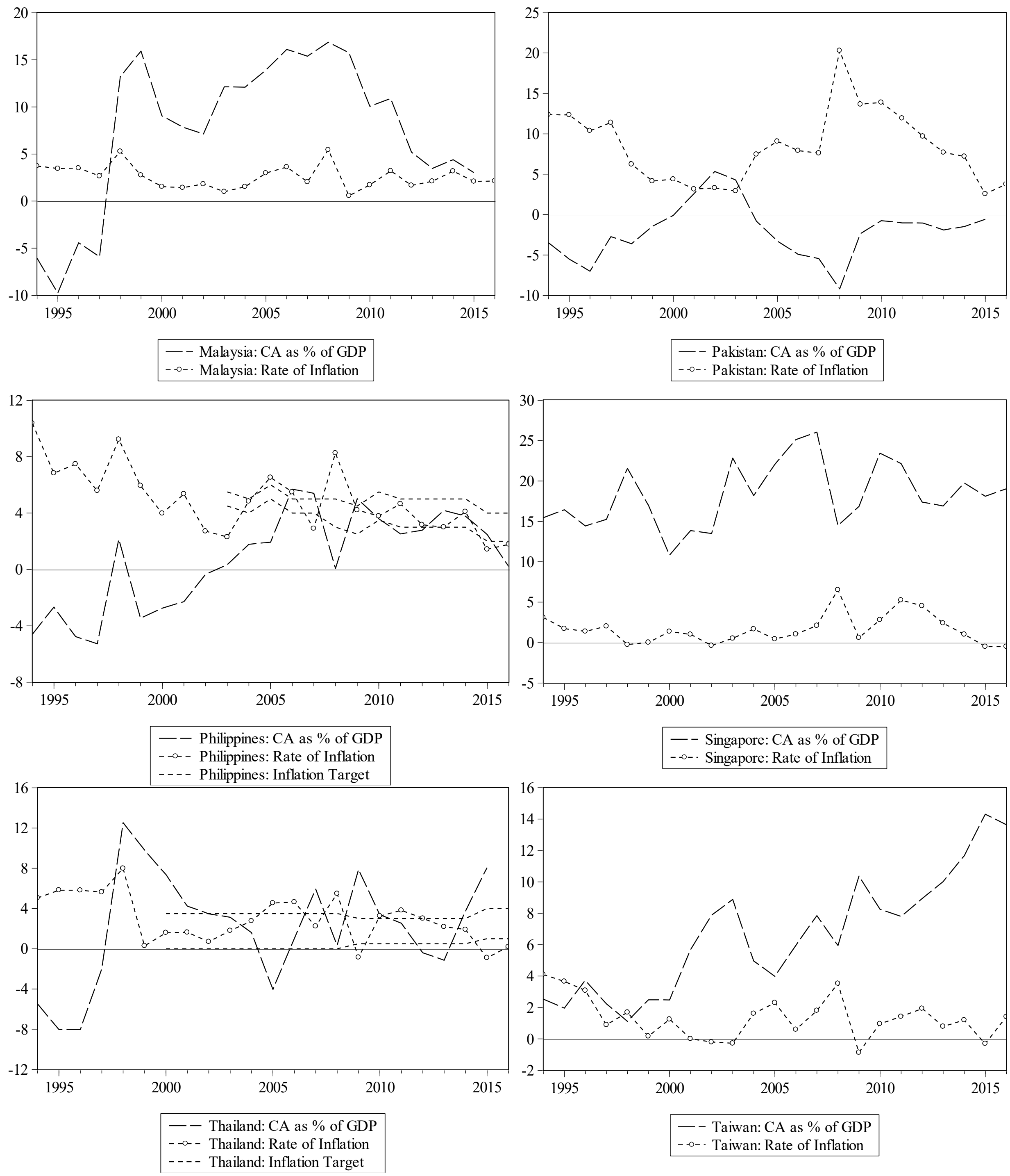
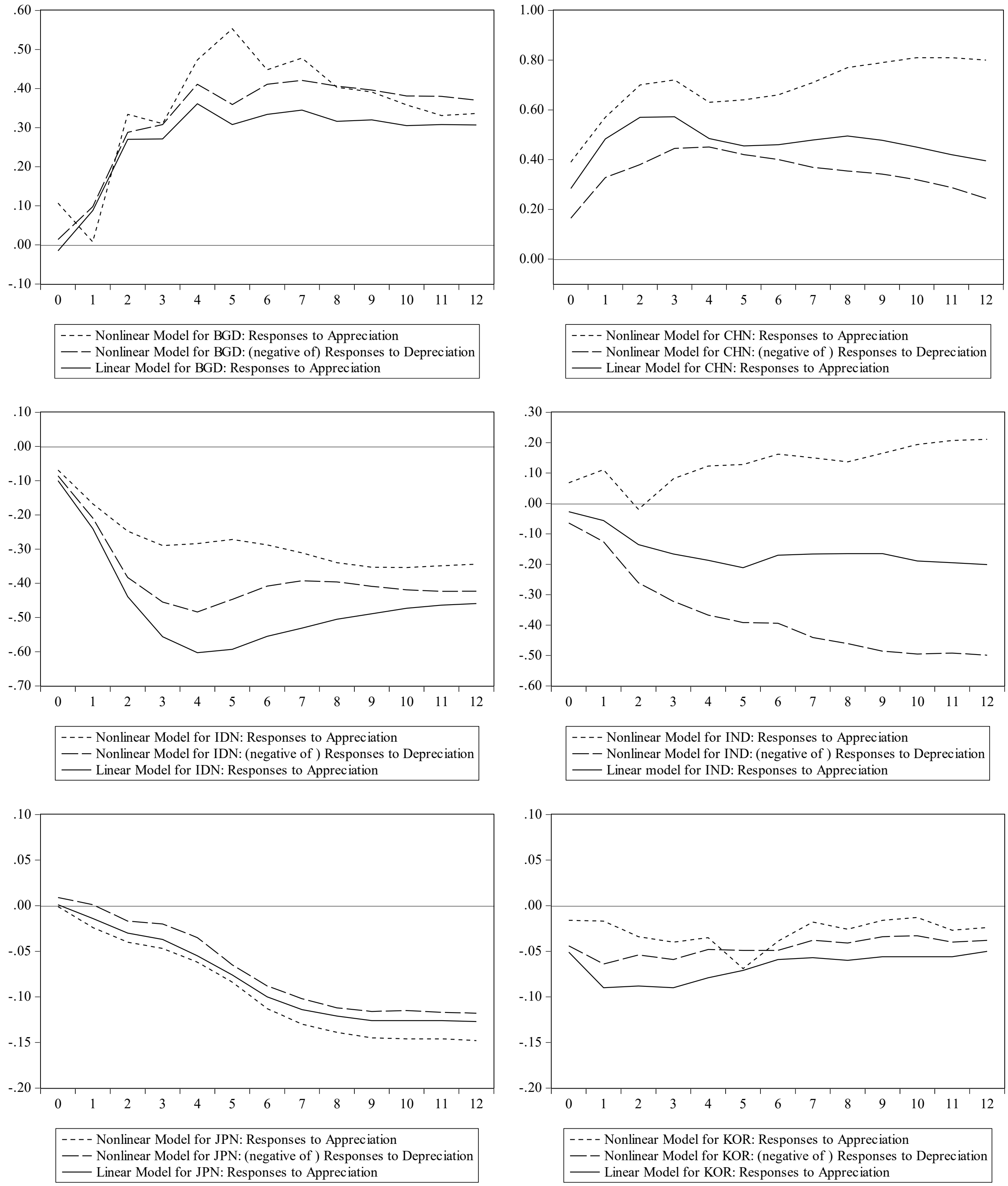
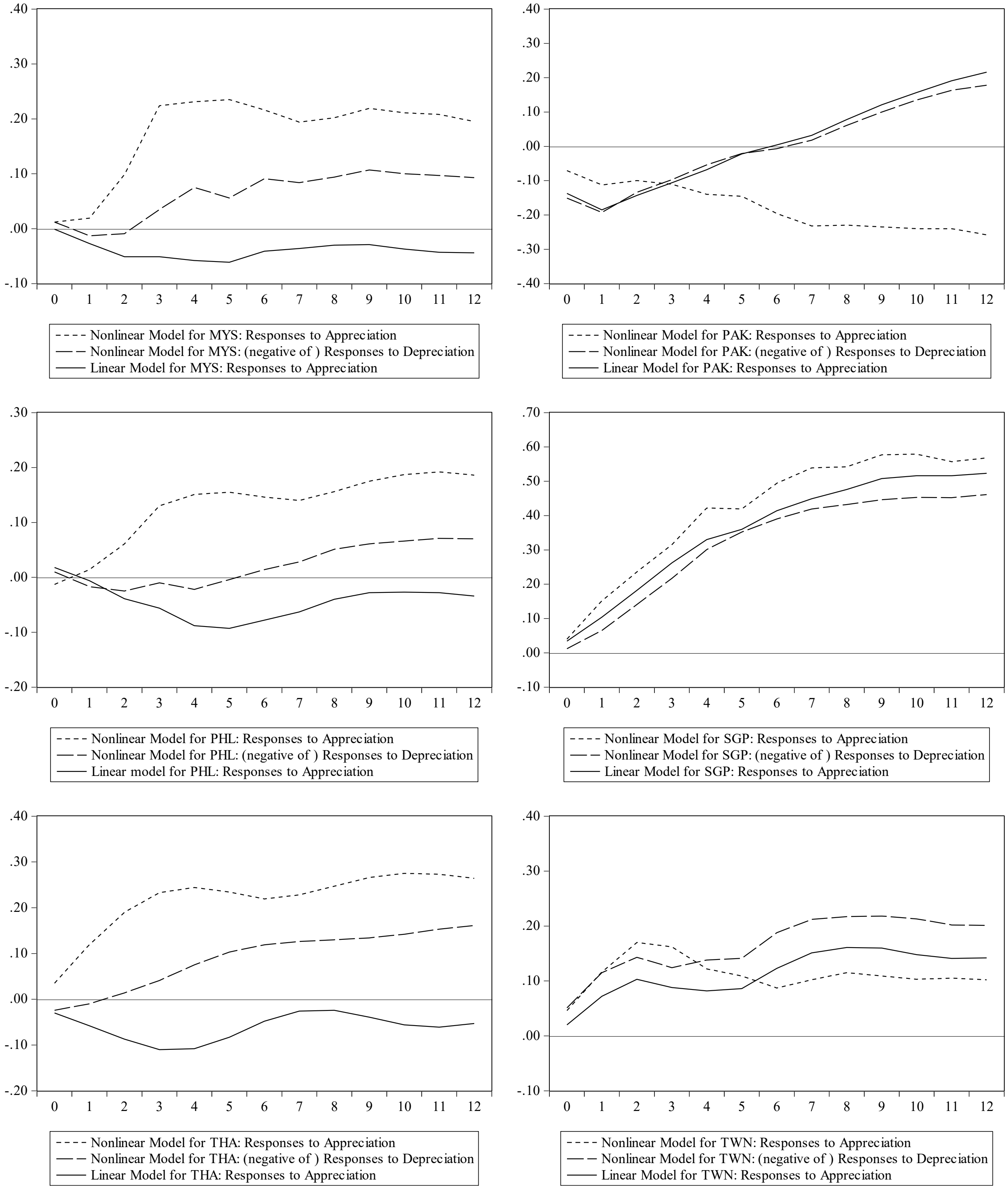

Malaysia: CA as % of GDP
Malaysia: Rate of Inflation
Pakistan: CA as % of GDP
Pakistan: Rate of Inflation
Philippines: CA as % of GDP
Philippines: Rate of Inflation
Philippines: Inflation Target
Singapore: CA as % of GDP
Singapore: Rate of Inflation
Thailand: CA as % of GDP
Thailand: Rate of Inflation
Thailand: Inflation Target
Taiwan: CA as % of GDP
Taiwan: Rate of Inflation
1995
2000
2005
2010
2015

Figure 3.3: The Response of CPI to 1% Change in Nominal Exchange Rate (USD based)




Nonlinear Model for MYS: Responses to Appreciation
Nonlinear Model for MYS: (negative of) Responses to Depreciation
Linear Model for MYS: Responses to Appreciation
Nonlinear Model for PAK: Responses to Appreciation
Nonlinear Model for PAK: (negative of) Responses to Depreciation
Linear Model for PAK: Responses to Appreciation
Nonlinear Model for PHL: Responses to Appreciation
Nonlinear Model for PHL: (negative of) Responses to Depreciation
Linear model for PHL: Responses to Appreciation
Nonlinear Model for SGP: Responses to Appreciation
Nonlinear Model for SGP: (negative of) Responses to Depreciation
Linear Model for SGP: Responses to Appreciation
Nonlinear Model for THA: Responses to Appreciation
Nonlinear Model for THA: (negative of) Responses to Depreciation
Linear Model for THA: Responses to Appreciation
Nonlinear Model for TWN: Responses to Appreciation
Nonlinear Model for TWN: (negative of) Responses to Depreciation
Linear Model for TWN: Responses to Appreciation

Appendix 3.1: IMF De Facto Classification of Exchange Rate Regime (IMF changed classification taxonomy in 2009)

Country	As of June 30, 2003	As of July 31, 2006	As of April 30, 2008	As of April 2010	As of April 2012	As of April 30, 2014
BGD	Other conventional fixed peg arrangement	Managed floating with no pre-determined path for the exchange rate	Other conventional fixed peg arrangement	Stabilized arrangement	Other managed arrangement	Stabilized arrangement
CHN	Other conventional fixed peg arrangement	Other conventional fixed peg arrangement	Crawling peg	Stabilized arrangement	Crawl-like arrangement	Crawl-like arrangement
IDN	Managed floating with no pre-determined path for the exchange rate	Independently floating	Managed floating with no pre-determined path for the exchange rate	Floating	Floating	Floating
IND	Managed floating with no pre-determined path for the exchange rate	Managed floating with no pre-determined path for the exchange rate	Managed floating with no pre-determined path for the exchange rate	Floating	Floating	Floating
JPN	Independently floating	Independently floating	Independently floating	Floating	Floating	Floating
KOR	Independently floating	Independently floating	Independently floating	Floating	Floating	Floating
MYS	Other conventional fixed peg arrangement	Managed floating with no pre-determined path for the exchange rate	Managed floating with no pre-determined path for the exchange rate	Other managed arrangement	Other managed arrangement	Other managed arrangement
PAK	Managed floating with no pre-determined path for the exchange rate	Other conventional fixed peg arrangement	Managed floating with no pre-determined path for the exchange rate	Floating	Floating	Other managed arrangement
PHL	Independently floating	Independently floating	Independently floating	Floating	Floating	Floating
SGP	Managed floating with no pre-determined path for the exchange rate	Managed floating with no pre-determined path for the exchange rate	Managed floating with no pre-determined path for the exchange rate	Other managed arrangement	Other managed arrangement	Stabilized Arrangement
THA	Managed floating with	Managed floating with	Managed floating with	Floating	Floating	Floating

	no pre-determined path for the exchange rate	no pre-determined path for the exchange rate	no pre-determined path for the exchange rate			
TWN	NA	NA	NA	NA	NA	NA

Source: Annual report on exchange rate arrangements and exchange restrictions, IMF, different issues. The details of the classification are available at:
<https://www.imf.org/external/pubs/ft/wp/2009/wp09211.pdf>

Appendix 3.2: Source of Data

Country	Consumer Price Index	Nominal Effective Exchange rate	Nominal Exchange Rate vis-à-vis USD
Bangladesh	IMF-IFS	Darvas, Zsolt (2012)[a]	IMF-IFS
China	IMF-IFS	IMF-IFS	IMF-IFS
Indonesia	IMF-IFS	BIS	IMF-IFS
India	IMF-IFS	BIS	IMF-IFS
Japan	IMF-IFS	IMF-IFS	IMF-IFS
Korea, Republic	IMF-IFS	BIS	IMF-IFS
Malaysia	IMF-IFS	IMF-IFS	IMF-IFS
Pakistan	IMF-IFS	IMF-IFS	IMF-IFS
Philippines	IMF-IFS	IMF-IFS	IMF-IFS
Singapore	IMF-IFS	IMF-IFS	IMF-IFS
Thailand	IMF-IFS	BIS	IMF-IFS
Taiwan	SB, Taiwan	BIS	CBC, Taiwan

[a] Data are updated regularly

IMF-IFS: International Financial Statistics (IFS) from International Monetary Fund (IMF)
< http://data.imf.org/?sk=5dabaff2-c5ad-4d27-a175-1253419c02d1&sId=1390030341854>
BIS: Bank for International Settlements
< http://www.bis.org/statistics/eer.htm>
CBC, Taiwan: Central Bank of the Republic of China (Taiwan)
< http://www.pxweb.cbc.gov.tw/dialog/statfile1L.asp?lang=1&strList=L>
SB, Taiwan: Statistical Bureau, Republic of China (Taiwan)
< https://eng.stat.gov.tw/ct.asp?xItem=12092&ctNode=1558&mp=5>
Darvas, Zsolt (2012) ‘Real effective exchange rates for 178 countries: a new database’, Working Paper 2012/06, Bruegel, 15 March 2012.
< http://bruegel.org/publications/datasets/real-effective-exchange-rates-for-178-countries-a-new-database/>

Chapter 4

Purchasing Power Parity in East, Southeast and South Asian Countries

4.1. Introduction

In this chapter, we examine whether the purchasing power parity (PPP) hypothesis holds for 12 Asian countries, located in East, Southeast and Southern part of Asia. The PPP hypothesis states that, in equilibrium, the nominal exchange rate between two currencies will vary proportionally with the relative price level in the two countries, so that the purchasing power of a unit of one currency would be the same in both countries. This version of PPP is often called absolute PPP. A weaker version of PPP, known as relative PPP, implies that the rate of change of one currency relative to another matches the difference in their inflation rates. Relative PPP is said to hold when the real exchange rate is constant over time; and movements from constancy represent deviations from PPP. The empirical literature, as in this chapter, examines the relative version of PPP.

While very few contemporary economists expect the PPP to hold continuously in the real world, 'most instinctively believe in some variant of PPP as an anchor for long-run real exchange rates' (Rogoff, 1996). PPP is an essential component of the economic literature both from theoretical and policy perspectives. If PPP does not hold in the long run, a large strand of open-economy macroeconomic theory may be flawed since most models are developed on the assumption that PPP holds in long run. Also, the estimates of PPP exchange rates are important for determining the degree of misalignment of the nominal exchange rate and the appropriate policy response, and for international comparisons of national income levels.

The contemporary literature on PPP mainly focuses on two issues: (1) whether PPP holds and (2) the reasons behind the observed deviations from PPP. The former re-examines the hypothesis whenever relevant new econometric techniques are developed, in order to ensure that earlier findings on the failure of PPP to hold were not due to the deficiencies of the econometric technique that had been employed previously. The second strand offers various explanations for the 'failure' of PPP to hold empirically. From the theoretical perspective, in the long run, the relative demand and supply of goods determine the equilibrium real exchange rate. Thus, any factors that affects demand or supply would affect the equilibrium rate. The factors that affect demand include, among others, shifts in the fiscal and monetary policy, domestic and foreign tastes, capital flows, etc. On the other hand, exogenous changes in the terms of trade and relative productivity can affect the supply side and cause deviations from the equilibrium rate. The Harrod-Balassa-Samuelson (HBS) hypothesis claims that the relative productivity differential among countries is an important factor behind the failure of PPP[12]. This chapter is focused on examining whether PPP holds. However, we can examine the relevance of the HBS hypothesis if real exchange rate has a deterministic trend, i.e. if real exchange rate is trend stationary.

The econometric techniques to examine whether PPP holds can be classified under two broad groups: (1) testing stationarity of the real exchange rate and (2) testing cointegration among nominal exchange rate, domestic price level and foreign price level. The real exchange rate can be calculated as: $q_t = e_t + p_t^* - p_t$, where q_t is the logarithm of the real exchange rate, e_t is the logarithm of nominal exchange rate (the amount of domestic currency required to buy one unit of foreign currency), p_t^* is the logarithm of foreign price index and p_t is the logarithm of domestic price index. Stationarity of q_t will provide support for the PPP hypothesis. Also, PPP will hold if e_t, p_t^* and p_t are cointegrated in the following regression: $e_t = \alpha + \beta p_t + \gamma p_t^* + \varepsilon_t$, where ε_t is the error term. If the series are cointegrated and the coefficient values are $\beta = 1$ and $\gamma = -1$, we call this version of PPP as strong PPP. In a bivariate setup, $e_t = \alpha + \beta rp_t + \varepsilon_t$, where rp_t is relative price. If $\beta = -1$, strong PPP holds. Under weak PPP the variables are cointegrated but the cointegrating vector can differ from unity. In such a case, an equilibrium relationship may exist but the differences in the construction of price indices, transaction costs and government policies can lead to a non-unitary relationship. In this chapter, we examine the strong version of relative PPP by testing whether the real exchange rate is stationary or not. Thus, we are imposing the above coefficient restriction, rather than testing it.

[12] According to HBS hypothesis, low wages in a low-productivity (but labour-intensive) country will cause prices to be low in its non-tradable sector, while, in a more productive economy, high wages will drive prices up in this sector. Assuming that PPP holds for traded goods, because of higher prices in non-tradable sector, overall price will be higher in more productive country. This will cause an appreciation in more productive country's real exchange rate and can lead to a breakdown in PPP. In developing countries, where productivity tends to be lower, this effect might provide a particularly useful explanation.

While the theory on PPP indicates that it should hold for open economies, it has been difficult to validate the hypothesis empirically. Sarno and Taylor (2002) provide a review of early studies on PPP while Bahmani-Oskooee and Hegerty (2009) provide a review of studies on less developed and transition economies. With little support for PPP using the unit root or cointegration method, even among the developed countries, this hypothesis has been tested extending these two methods in a panel setup, in the presence of structural breaks and in the nonlinear framework. Although analysis with panel setup, structural breaks and regime-dependent models are able to provide more support for PPP, the empirical evidence can still be described as mixed.

Due to the low power of the unit root tests[13], especially for a short sample period, PPP hypothesis is often tested in a panel setup. However, Taylor and Sarno (1998) argue that the rejection of unit root hypothesis in the null can occur even if just one of the series is stationary in the panel. Using Monte Carlo experiments, calibrated on the USD based real exchange rates among the G-5 countries with a single stationary process (with slope coefficient of 0.95) and three unit root processes, they found that the null hypothesis of nonstationarity is rejected in 65 percent of the simulations at the 5% level of significance. Therefore, a rejection of the joint null hypothesis is not very informative, so that it cannot be concluded that PPP holds for all of them. Nonlinear models are also widely used in the contemporary literature and they are more suited for regime-dependent stationarity. In this chapter, we want to examine whether PPP holds in our selected countries as a whole, not in some regimes, so that we do not use regime-dependent models. Also, our sample time period is relatively longer compared to earlier studies, and thus, we expect the unit root tests to have more power. In these circumstances, we believe that unit root tests allowing for multiple breaks will best serve our purpose.

Perron (1989) argues that, in the presence of a structural break, the standard augmented Dickey-Fuller (ADF) test is biased towards the non-rejection of the null hypothesis. A large change in the constant and/or slope of the trend function (the deterministic component of the series that include constant and time trend) will reduce the power of the test dramatically while small changes, especially in level, will reduce power slightly. Hence, it is not necessary to account for all structural changes and focus on the large shifts. For this reason, we allow up to two structural breaks to

[13] The probability of rejecting the null hypothesis of a unit root when the null is false. A major criticism of the unit root test is its low power as it is unable to distinguish regression coefficient between 0.95 and 1. Hence, the test rejects PPP even though the real exchange rate is slowly mean reverting.

accommodate two major crisis, 1997 Asian crisis and sub-prime crisis, faced by this region. However, it is not necessary that breaks will coincide with these two crises, for any particular country, as breaks can be caused by shocks purely internal to that country. To accommodate this, structural breaks will be determined endogenously in our analysis.

Structural breaks can affect the trend function of the series differently: (1) change to new trend function occurs instantaneously, known as the Additive Outlier (AO) model and (2) change to new trend function is gradual, known as Innovational Outlier (IO) model. This distinction between AO and IO models is important as the statistical procedures to test unit roots are different under these two models. Moreover, in this chapter, we focus only on the structural breaks in the constant of the trend function while the slope of the trend function remains same. In this case, there is a jump in the trend function without a change in slope. These types of breaks are known as structural shifts. Also, changes in the slope of the trend function may not be appropriate since the real exchange rate is not a trending series for most of our selected countries and breaks in slopes can change the sign of the slope of the trend function for some period. Although trend stationarity in the series will support the HBS hypothesis, such changes in slope will not be helpful for examining the hypothesis.

The objective of this chapter is to examine whether PPP holds for 12 Asian countries during the period 1974 to 2016.[14] Similar to chapter 3, the countries included in this study are: Bangladesh, China, India, Indonesia, Japan, Republic of Korea, Malaysia, Pakistan, the Philippines, Singapore, Taiwan and Thailand. As mentioned in chapter 3, almost half of the world population live in these 12 countries. Also, these countries are at different stages of economic development and some of these countries performed extremely well, in terms of economic growth, in the past two decades or so. Hence, it is essential to examine the validity of the PPP hypothesis for these countries as it will shed light on the relevance of contemporary open economy macroeconomic models (which assumes that PPP hold in the long-run) for these countries. The diverse nature of these countries will also allow us to explore whether the validity of PPP hypothesis are dependent to economic development and macroeconomic environment. For that purpose, we use Consumer Price Index (CPI) based real exchange rate vis-à-vis USD to examine

[14] The Bretton Woods System dissolved between 1968 and 1973. By 1973, major currencies began to float against each other. However, because of the oil price shock in 1973, exchange rate was volatile for most countries. To avoid the volatility at the beginning of sample period, we start our analysis from 1974.

the PPP hypothesis. Using USD-based real exchange rate allows us to examine the hypothesis for a longer period as data on the real effective exchange rate, a trade weighted real exchange rate, is available for half of the countries from 1994 or a later period. For most of our selected countries, USA is their major trading partner, USD is used as the numeraire currency in trade, they hold foreign exchange reserve in USD, a large portion of their assets and liabilities are denominated in USD and the local currency was pegged against the USD for a considerable period of time. Hence, it is reasonable to examine the PPP hypothesis vis-à-vis USD. Also, we examine the CPI-based real exchange rate as the Wholesale Price Index (WPI) or Producer Price Index (PPI) data are not available for some of our selected set of countries.

We use quarterly data from 1974:1 to 2016:4 for all countries except Bangladesh (starts at 1993:3) and China (starts at 1986:1). The data for all countries, except Taiwan, are downloaded from the International Financial Statistics (IFS) of International Monetary Fund (IMF) website. For Taiwan, the data on the nominal exchange rate and domestic price index are taken from the website of the Central Bank of the Republic of China and the website of National Statistics of the Republic of China, respectively. The nominal exchange rate is defined as the price of the US Dollar (USD) in the domestic currency, i.e., the amount of domestic currency required to buy one unit of USD. Hence, an increase in the nominal exchange rate represents a depreciation of the domestic currency. Consistent with this definition of the nominal exchange rate, the real exchange rate is the price of US commodities in terms of domestic commodities, so that an increase in real exchange rate represents a depreciation in real terms that makes foreign commodities expensive.

This chapter uses unit root methods to examine the stationarity of the real exchange rate, both without and with structural breaks. The contributions of this chapter are: (1) compared to the earlier studies, we examine stationarity of real exchange rate for a longer period of time with more efficient unit root tests. This will provide a new benchmark for future studies; (2) all the previous empirical studies on our selected Asian countries, allowing structural breaks, use tests that only examine trend stationarity of the real exchange rate, while we examine level stationarity of the series allowing for multiple breaks; (3) all the previous empirical studies on our selected country assume gradual adjustment of the trend function due to structural break(s) and use tests based on IO model. In this chapter, we make no such assumption and use tests based on both the AO and IO frameworks. Examination of both models is necessary since structural shocks can affect the real exchange rate differently across countries. Further, the adjustment process can also be different; and (4) we also examine trend stationarity using both AO and IO framework. This will allow us to compare our result with earlier studies and provide a new benchmark for future empirical research.

The summary of our results is provided in Table 4.1. Compared with earlier studies, we have found significantly greater evidence of PPP since the unit root hypothesis can be rejected for 5 (Japan, Korea, Philippines, Singapore and Taiwan) out of our 12 selected countries, using the unit root tests without accommodating structural break(s). In addition, with structural shifts, the unit root null can be rejected for an additional country, Thailand. Among the 6 countries for which there is evidence of PPP, we find little support for the HBS hypothesis in any of these countries.

The organization of this chapter is as follows. Section 4.2 gives a brief review of the methods used in the empirical literature, along with a review of some recent studies that use these methods. Section 4.3 presents our results using different unit root tests without incorporating structural breaks. Section 4.4 provides a description of the unit root tests allowing for structural breaks. The section also outlines the difference in the testing procedures of the AO and IO models. Using these two frameworks and allowing up to two structural shifts, we present our results in this section. Section 4.5 is the concluding part of this chapter.

Table 4.1: Summary of Results

Test	Level Stationary	Trend Stationary
ADF	Japan[c], Korea[b], Philippines[c] Singapore[c]	Korea[c]
DF-GLS	Korea[a], Philippines[c]	Korea[b], Philippines[c]
MZt	Korea[a], Philippines[b], Taiwan[c]	Korea[a], Philippines[a], Singapore[c], Taiwan[c]
AO-1 Break: Perron	..	..
AO-2 Breaks: Perron	Japan[c], Taiwan[b], Thailand[b]	Taiwan[b], Thailand[b]
IO-1 Break: Perron (level), ZA (trend)	Pakistan[c]	Japan[b]
IO-2 Breaks: Perron (level), LP (trend)	India[c], Japan[c], Taiwan[b], Thailand[c]	Taiwan[b], Thailand[c]
IO-1 Break: LS	NA	Korea[c]
IO-2 Breaks: LS	NA	Korea[b], China[c]

[a], [b] and [c] represents 1%, 5% and 10% level of significance, respectively.

4.2. Literature Review

In this section, we briefly discuss the methods used in the empirical literature to test the PPP hypothesis and some empirical studies on our selected Asian countries. The empirical literature on PPP is quite large, and the sophistication of the testing procedures employed has evolved with advances in econometric techniques. Nevertheless, it is useful to separate the contemporary models into two broad categories: linear models and nonlinear models. Both types can be estimated using unit root tests as well as using cointegration techniques. For example, the Markov switching model allows testing for the unit root of a variable in the ADF framework, as well as the examination of cointegration among the variables using the residual based cointegration tests. In this chapter, unless otherwise stated, the empirical studies examine the PPP hypothesis with: (1) CPI based real exchange rate for the selected country vis-à-vis USA and (2) monthly data to increase the number of observations since most of these studies deal with a relatively short period of time.

4.2.1. Linear Models

From the mid-1980s onwards, a standard approach has been to employ variants of the augmented Dickey-Fuller (ADF) test to investigate unit root in the real exchange rate. The standard ADF unit root test adds augmenting (truncation) lags to control for serial correlation. Phillips and Perron (PP) (1988) estimate the ADF equation without the augmenting lags and modify the test statistic to correct for serial correlation. This modified test statistic has the same asymptotic distribution as the ADF t-statistic.

Elliott, Rothenberg and Stock (1996) propose a two-step unit root testing procedure, known as DF-GLS. In the first step, the data are de-meaned or de-trended (by both constant and trend) using the generalized least square (GLS) method. The ADF based test is then performed on the de-meaned or de-trended series. The DF-GLS test statistic follows the Dickey-Fuller distribution for the de-meaned case only. For the de-trended case, Elliott et al. (1996) provide simulated critical values. Using simulation, they also show that the DF-GLS test has higher power than the ADF test.

Schwert (1989) finds that, if the series has an ARMA representation with a large and negative moving average errors, both the ADF and PP tests are severely size distorted and rejects the unit root null too often when it is true. Perron and Ng (1996) modify the PP tests to correct for this size distortion[15]. These tests are known as the M-tests. Ng and Perron (2001) use GLS de-trending procedure to create an efficient version of the M-tests proposed in Perron and Ng (1996). These tests are efficient as they do not exhibit severe size distortions and have higher power compared to the PP test.

The ADF, PP, DF-GLS and M-tests assume a unit root under the null hypothesis. There are also some tests where the null hypothesis is stationarity of the series. The most commonly used such test is the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) (1992) test. This is a one-sided right-tailed test, which follows the same procedure of the PP test to correct for serial correlation. The critical values for this LaGrange Multiplier (LM) based test statistic are calculated via simulation.

Cointegration analysis, originally developed by Engle and Granger (1987), specifies that any two $I(1)$ series are cointegrated if a linear combination of the two is $I(0)$. In our context, if both the nominal exchange rate e_t and the relative price rp_t are $I(1)$ and the linear combination of these two, which is the error term ε_t, is $I(0)$, then e_t and rp_t is cointegrated and PPP will hold. This method, also known as the residual-based cointegration analysis, can be employed for more than two variables. The Johansen (1988) and Johansen and Juselius (1990) method, on the other hand, use a vector autoregressive (VAR) framework for the cointegration analysis. This method has two notable advantages over the residual-based method: (1) there is no need to specify the dependent variable in the estimation process, and (2) it is possible to identify all the cointegrating relations among the variables. Despite its limitations, the residual-based method is used more frequently because of its flexibility to accommodate structural breaks, panel or nonlinear models.

Empirical studies employing unit root tests during the floating exchange rate regime often do not reject the unit root hypothesis of the real exchange rate for our selected Asian countries or developing countries (see, Bahmani-Oskooee and Hegerty (2009) for details); so that they tend to reject PPP. Doğanlar and Özmen (2000), using the ADF and PP tests, do not reject the unit root hypothesis for any of the 18 developing countries (including India, Indonesia, South Korea, Pakistan, Philippines and Sri Lanka) for the period 1986 to 1997. For a longer time period of 1973

[15] The probability of rejecting the null hypothesis of a unit root when the null is true.

to 1998, using the ADF and PP tests, Achy (2003) investigates PPP for 38 middle-income countries (including Indonesia, Korea, Malaysia, Philippines and Thailand) and also fails to reject the unit root hypothesis against the alternative hypothesis of level and trend stationarity for any of the 5 countries selected for analysis in this chapter. Hooi and Smyth (2007) applies the ADF test for 15 Asian countries (including Bangladesh, China, India, Indonesia, Korea, Malaysia, Pakistan, Philippines, Singapore and Thailand) for the period 1990 to 2005; and are able to reject the unit root null against the alternative hypothesis of trend stationarity for India, Pakistan, Philippines and Thailand. Bahmani-Oskooee, Kutan and Zhou (2009) examine PPP for 113 countries, including all the countries of this chapter except Taiwan, using real effective exchange rates from 1980 to 2005. Based on the KPSS test, they find the series to be level stationary for Singapore and trend stationary for Bangladesh, Korea, Malaysia, Singapore and Thailand.

Since the unit root test examines the strong version of PPP, the weaker version can be analysed with cointegration analysis. Theoretically, this weaker version should provide more support for PPP. Lee (1999) examines PPP for 13 Asia-Pacific countries (including Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan and Thailand) with both CPI and WPI based real exchange rates from 1957 to 1994. The unit root hypothesis, based on the PP test, is not rejected for any of these 8 Asian countries with both CPI and WPI based real exchange rate. With a dynamic version of the bivariate error correction model, with nominal exchange rate and CPI ratio, PPP holds for Korea and Philippines at the 5 percent level of significance. With the WPI ratio, PPP holds for two additional countries, Malaysia and Singapore. On the other hand, several studies have found that cointegration tests do not provide more support for PPP hypothesis than the unit root tests. For example, Doğanlar (1999) employs both the residual-based and Johansen's cointegration techniques to examine PPP for 5 Asian countries (India, Indonesia, Pakistan, Philippines and Turkey) for the period 1980 to 1995. He finds that PPP holds only for Turkey. Janjua and Ahmad (2006) examine PPP for four South Asian countries (Bangladesh, India, Pakistan and Sri Lanka) with CPI and WPI based real exchange rate for the period 1984 to 2002. Based on ADF and PP tests, PPP hypothesis is rejected for all countries, with only weak evidence of PPP for Pakistan using the residual based cointegration test. Doğanlar, Bal and Özmen (2009) employ Johansen's cointegration test on the data of 10 emerging countries (including India, Indonesia, South Korea, Pakistan and Philippines) for the period 1995 to 2005. They report that PPP holds in none of these 5 Asian countries. Kim and Jei (2013) examine PPP for Japan and

Korea for the period 1974 to 2011. Using Johansen's cointegration method, they also fail to find evidence of PPP. There is thus a very limited support for PPP even with cointegration analysis. A possible reason for this finding can be the shortness of the time period included in these studies.

Now looking at the findings for industrialized countries, with unit root tests and the cointegration method, the support for PPP among the major industrialized countries against one another is also mixed (see, Sarno and Taylor (2002) for details). The results tend to largely depend on the numeraire currency, estimation method and, especially, the time period under investigation. For example, Taylor (1988), using bivariate residual based cointegration test, finds no evidence of PPP for five industrialized countries (Canada, France, Japan, UK and West Germany) for the period 1973 to 1985. On the other hand, Cheung and Lai (1993), using Johansen's cointegration test, find significant evidence supporting PPP for all five industrialized countries (Canada, France, Switzerland, UK and West Germany) under study for the period 1974 to 1989.

Given the lack of support for PPP employing the unit root and cointegration tests, several studies examine whether the real exchange rate is stationary but has a long memory process, known as fractional integration[16]. The fractionally integrated process can be estimated with the parametric autoregressive fractionally integrated moving average (ARFIMA) method or the semi-parametric methods, as in Geweke and Porter-Hudak (GPH) (1983). For fractional cointegration, the residual-based test can be applied by examining whether the residual series exhibits long memory. To illustrate the findings using such methods, Chou and Shih (1997) use quarterly data from 1965 to 1992 for 4 East Asian countries (Hong Kong, Korea, Singapore and Taiwan). With Johansen's method (with a time trend), PPP holds for Hong Kong and Singapore. Using fractional cointegration (with a trend), PPP seems to hold for only Korea. Gil-Alana and Jiang (2013), using both parametric and semiparametric method for the period 1994 to 2010, find no evidence of PPP for China.

A number of studies have found that the tests typically employed during the 1980s and 1990s, using data from floating exchange rate regimes, to examine the stationarity of the real

[16] In its general form, the real exchange rate process may be expressed as: $\mathrm{B(L)}(1-\mathrm{L})^{\mathrm{d}}\mathrm{q_t} = \mathrm{E(L)}\varepsilon_{\mathrm{t}}$, where $\mathrm{B(L)}$ and $\mathrm{E(L)}$ are both lag polynomials with roots lying outside the unit circle, and ε_{t} is a white noise process. Fractionally integrated processes are stationary but are more persistent than pure autoregressive moving average (ARMA) processes. If $\mathrm{d}=0$, then the real exchange rate is stationarity, while $\mathrm{d} \in (0,0.5)$ implies weak stationarity, $\mathrm{d} \in (0.5,1)$ implies mean-reverting nonstationary process and $\mathrm{d} \geq 1$ implies that the real exchange rate follows a random walk.

exchange rate may have very low power to reject a null hypothesis. The Monte Carlo experiment of Shiller and Perron (1985) demonstrates that the power of the test depends on the span of the data rather than the number of observations, so that merely increasing the frequency of the observation from annual to monthly will not solve the power problem. Inspired by these findings, several studies examine long span data to examine the PPP hypothesis. Lothian and Taylor (1996) use two centuries of the annual real exchange rate data (1791 to 1990) for USD-Pound Sterling and Franc-Pound Sterling; and find strong evidence for mean reversion in real exchange rates (with standard DF and PP unit root test). For Japan, Ito (1997) uses DF and ADF unit root test to examine PPP vis-à-vis USA with annual data from 1879 to 1995. Using the CPI based real exchange rate, the unit root tests reject the validity of PPP. Using the WPI based real exchange rate, the unit root null is rejected against trend stationarity; and thus, provides support for the HBS hypothesis.

Due to the unavailability of long span data for most countries and to circumvent the power problem, some studies have reverted to panel analysis to examine PPP. Hakkio (1984) and Abuaf and Jorion (1990) are two early studies that test the PPP hypothesis using the panel analysis and find support for PPP. A series of studies followed this approach and found evidence supporting PPP. Two widely used panel unit root test in the empirical literature are Levin, Lin and Chu (LLC) (2002) and Im, Pesaran and Shin (IPS) (2003) test. The null hypotheses of these two tests are different since a common unit root process is the null hypothesis in the LLC test while the country-specific unit root is the null hypothesis in the IPS test. Among other studies is that by Baharumshah, Aggarwal and Tze-Haw (2007) who employ both LLC and IPS tests for 6 Asian countries (Indonesia, Malaysia, Philippines, South Korea, Singapore and Thailand) for the period 1973 to 2006. Using the USD and the Yen as the numeraire currency and CPI based real exchange rates, the unit root hypothesis is rejected under both numeraire currencies.

Both the LLC and IPS tests assume cross-sectional independence (e.g., Japan's real exchange rate does not depend on Korea's real exchange rate) among the series. A simulation-based study of Banerjee, Marcellino and Osbat (2004) shows that cross-sectional dependence can lead to strong size distortions in these tests. To correct for this problem, Pesaran (2007) proposes a method to allow for cross-section dependence within the IPS framework. Drine and Rault (2008) employ panel unit root and panel cointegration test on 80 developed and developing countries, by categorizing them into different groups. Based on panel unit root tests with cross-sectional dependence, PPP holds for OECD countries only. Based on the panel cointegration test, PPP holds for Middle East and North African countries. However, in African, Asian (including China, Hong Kong, India, Indonesia, Korea, Malaysia, Philippines, Singapore and Thailand), Latin American and Central and Eastern European countries, PPP does not hold. Hassan, Hoque and Koku (2014) employ panel unit root test for 5 South Asian countries (Bangladesh, India, Nepal, Pakistan and Sri Lanka) allowing for cross-sectional dependence between the series. Covering the period 1957 to 2014 (data for Bangladesh and Nepal start from 1993 and 1963, respectively), the panel unit root test results support the validity of PPP in these countries.

The evidence supporting PPP from the panel studies is significantly higher compared to individual country based analysis. However, as mentioned in the introduction of this chapter, based on Taylor and Sarno (1998), the rejection of the unit root does not necessarily mean that all series in the panel are stationary. Also, the performance of the panel unit root tests depends on the countries included in the panel. To illustrate, Alba and Papell (2007) find stronger evidence of PPP if countries are homogenous in terms of low inflation, closer to the numeraire country, more open to trade, moderate nominal exchange rate volatility or similar economic growth rate to the numeraire country.

To overcome the power problem of the unit root tests, this chapter employs unit root tests in the presence of structural breaks and are discussed in detail in section 4.4. A point to be noted here is that these breaks can also be included in the cointegration tests and in panel setup.

4.2.2. Nonlinear Models

As discussed above, firm support for PPP has remained elusive with linear models. As a result, nonlinear methods have been growing in popularity because they can capture asymmetries, transaction costs and other distortions better than the linear model. Nonlinear models can be broadly classified into two groups: threshold models (where the regime switch depends on an observable variable) and Markov (or simple) switching models (where the regime switch depends on an unobservable state). A brief description of nonlinear models is provided in Appendix 4.1.

The threshold regression model allows for two or more branches governed by the values of the threshold variable in the estimation process. Rather than using one discrete threshold, tests of PPP can also incorporate a band in which a variable behaves as if it has a unit root, but can be mean reverting outside the band. Obstfeld and Taylor (1997) have investigated the nonlinear nature of the adjustment process in terms of a threshold autoregression (TAR) model that allows for a transaction costs' band within which no adjustment takes place, while, outside of the band, the process switches abruptly to become stationary autoregressive. This method is particularly relevant in countries in which government responds to changes in the exchange rate. In general, government policy is likely to be more active once exchange rates move farther away from their targets and also depends on the direction of the deviations from PPP. If the deviation from PPP represents an increase in the price for foreign goods, governments often try to accommodate this price increase in order to promote exports. Therefore, the exchange rate movements might not be symmetric in adjusting toward the equilibrium rate.

Enders and Granger (1998) employ an ADF-type TAR model to examine the null hypothesis of the unit root against the alternative of the stationary TAR process. They also provide critical values for the joint test of the unit root in both regimes. Caner and Hansen (2001) also use a two-regime TAR model and develop asymptotic theory for inference in the context of nonlinearity and nonstationarity. They propose a Wald-type statistic which allows examination of the unit root in both regimes and of the unit root in any one regime. Chen, Chang, Zhang and Lee (2011) employ the threshold unit root test of Caner and Hansen (2001) for China and Taiwan for the period 1986 to 2009. Their result supports the nonlinear model and finds that the unit root hypothesis can be rejected for China and Taiwan at the 10 percent and 5 percent levels of significance, respectively.

Cointegration analysis can be used with the residual-based framework by applying the threshold unit root test on the residual series, as in Enders and Siklos (2001). Using this framework, Chang, Su, Zhu and Liu (2010) examine PPP for 4 BRIC countries (Brazil, Russia, India and China) for the period 1992 to 2006. The linear residual based cointegration test rejects cointegration relations for all four countries while threshold cointegration test provide support for PPP for all countries other than China. However, using this method on a slightly longer time period (1985 to 2008), Lu and Chang (2011) find a cointegrating relationship for China, using both the USD and the Yen as the numeraire currency.

In threshold models, the regime switches when the threshold variable crosses a certain threshold, so that there is a sharp cut-off between the branches. An alternative is the Smooth

Transition Regression (STR) model where the regime switches gradually in a smooth fashion. In the Smooth Transition Autoregression (STAR) models, adjustment takes place in every period but the speed of adjustment varies with the extent of the deviation from parity. Two simple transition functions, suggested by Granger and Teräsvirta (1993) and Teräsvirta (1994), are the exponential STAR (ESTAR) and the logistic STAR (LSTAR). For an LSTAR model, switches between two regimes are asymmetric whereas the adjustment is symmetric for the ESTAR model.

Taylor, Peel, and Sarno (2001) find evidence of nonlinear mean reversion of real exchange rates for four industrialized countries (UK, USA, Germany and Japan) during the post-Bretton Woods period. Their estimated ESTAR models imply that the real exchange rate behaves like a unit root process when they are closer to the long run equilibrium level and become mean reverting when they are further away from the equilibrium. Baharumshah, Liew and Chowdhury (2010) use quarterly data from 1965 to 2004 and CPI based real exchange rate for 6 Asian countries (Indonesia, Malaysia, Philippines, Singapore, South Korea and Thailand) to examine the PPP hypothesis. Estimating ADF-type STAR model, they find evidence of nonlinear mean reversion for all countries. They also find the adjustment to be asymmetric for all countries except Malaysia.

Michael, Nobay, and Peel (1997) apply the STAR model using monthly data for the interwar period (1921-25) for the France-USA, France-UK and UK-USA, as well as long span annual data for UK-USA (1791 to 1992) and France-UK (1802 to 1992). Using residual based cointegration, they find no evidence of PPP. However, applying the ESTAR method to the residual series, they find that the residual series exhibit nonlinear mean reversion, and thus support PPP. Following the same methodology, Chen and Wu (2000) and Hong and Oh (2009) also find nonlinear mean reversion in the residual series for Japan and Taiwan from 1974 to 1997, and for Korea with data from 1980 to 2007, respectively.

Kapetanios, Shin and Snell (KSS) (2003) develop a unit root test against the alternative hypothesis of stationary ESTAR. Using this test and quarterly data from 1968 to 2001, Liew, Baharumshah and Chong (2004) reject the unit root hypothesis in 8 USD-based (India, Indonesia, Japan, Korea, Malaysia, Pakistan, Singapore and Thailand) and 6 Japanese yen-based real exchange rates among the 11 Asian countries. On the contrary, the standard ADF test fails to reject the null for all countries. Bahmani-Oskooee, Kutan and Zhou (2008) examine the PPP hypothesis with the real effective exchange rate for 88 developing countries. Using monthly data from 1980 to 2003 and employing the KSS test, they find mean reversion for Indonesia, Korea, Malaysia, Pakistan, Singapore and Thailand. However, the process is not mean reverting for Bangladesh, China, India and the Philippines.

For TAR and STAR models, the regimes are solely determined by the magnitude of an observable weakly exogenous variable. On the contrary, in Markov switching models the regimes are determined by an unobserved state or regime variable that follows a discrete state Markov process. To illustrate some results using this method, Kanas (2009) examines real exchange rate stationarity for 43 developing countries with monthly data ranging from 1977 to 2004 (some countries have shorter data span). 36 countries show evidence of regime-dependent stationarity, i.e., in one regime the real exchange rate is stationary and in another regime it is not. He studies 8 countries (India, Indonesia, Korea, Malaysia, Pakistan, Philippines, Singapore and Thailand) that are discussed in the present chapter. Out of these 8 countries, only Malaysia and Singapore proved not to have regime-dependent stationarity at the 5% level of significance.

To sum up, nonlinear models test either regime-dependent stationarity of the real exchange rate or mean reversion of the series, i.e., examines global stationarity and assume that the series is locally nonstationary. In light of these, we believe that unit root tests in the presence of structural breaks are more suited for testing the PPP hypothesis.

4.3. Unit Root Tests (with No Structural Break)

We start with three t-statistic based test, the ADF, DF-GLS and M-test[17], without accommodating structural breaks. A brief description of these tests is provided in the previous section. Results from the unit root tests depend on the number of augmenting lags included in the test. For these tests, we select the number of lags based on three different criteria, namely the modified Akaike information criteria (M-AIC), the T test procedure (also known as general-to-specific) and a fixed number of lags. We expect these three lag selection schemes to provide a better picture of the unit root behavior of the series. Said and Dickey (1984) suggested using at least $T^{1/3}$ number of augmenting lags, where T is the number of observations. Based on this suggestion, we chose a fixed lag of 5. The T test procedure, as suggested by Ng and Perron (1995),

[17] Among the class of M-test, we only perform the t-statistic based test, namely MZ_t. As suggested in Ng and Perron (2001), we use GLS-detrended autoregressive estimate of the spectral density at frequency zero.

starts by pre-defining a maximum number of lags. It keeps dropping the last lag until the last included lag has a significance level less than the pre-specified level of significance. Generally, following Schwert (1989), the maximum number of lags is chosen as $integer\{12(T/100)^{1/4}\}$, which gives a maximum lag of 13 for our selected countries. We fix the pre-defined level of significance at 10 percent in this chapter.

It has been widely observed that the Akaike information criteria (AIC) and the Schwarz Bayesian information criteria (BIC) tend to select a very small number of augmenting lags. If the ARMA process has a large negative moving average error, a high order augmented autoregression is required for the test to have good size. Ng and Perron (2001) propose a class of modified information criteria (MIC) with a penalty factor that takes account of the bias in the sum of the autoregressive coefficients. This bias decreases as the augmenting lag increases; and thus, reduces the penalty factor. In Monte Carlo experiments, the MIC yields significant size improvements to the DF-GLS test and M-AIC performs relatively better than others. One point to note here is that it is problematic to have too many lags or too few lags. If the lag length is too small, the serial correlation in the error is not removed, so that the test statistic would be biased. On the contrary, the power of the test suffers if the lag length is too large. However, Monte Carlo experiments suggest that it is better to have too many lags than having too few lags.

The results of the unit root tests are presented in Table 4.2. The table shows that the T test generally selects more lags than the M-AIC. This table indicates the unit root hypothesis can be rejected only for Korea based on the three unit root tests. For Korea, under the three unit root tests with three different lag selection criteria for each, the chosen lag varies between 2 to 5. For all other countries, M-AIC and T test chose considerably different lags and the results of the unit root tests are sensitive to the number of augmenting lags. The unit root hypothesis is rejected for more countries when the number of augmenting lags is chosen based on the T test criterion, compared to M-AIC and fixed lags.

At the 5 percent level of significance, the unit root null can be rejected against the alternative of level stationarity for Korea under the three tests, and for Philippines under the MZ_t test[18]. At the 10 percent level of significance, the ADF test rejects the null for Japan, Philippines

[18] A modified t-statistics based test based on Perron and Ng (1996).

and Singapore; the DF-GLS test rejects the null for Philippines; and the MZ_t test rejects the null for Taiwan. We have to be cautious for the case of Japan since only 3 lags are chosen based on the M-AIC and the other two tests fail to reject the unit root null. The result for Japan may have size distortions due to presence of a large negative moving average error. However, the ADF tests are less size distorted compared to the PP test (for details, see Perron and Ng (1996)).

At the 10 percent level of significance, the ADF test rejects the unit root null against the alternative of trend stationarity for Korea only. The MZ_t test rejects unit root null against trend stationarity for Korea and Philippines at the 1 percent level of significance; and for Singapore and Taiwan at the 10 percent level of significance. PPP seems to hold for Korea and Philippines at the 5 percent and 10 percent levels of significance, respectively, under the DF-GLS test. For Korea, Philippines, Singapore and Taiwan, the real exchange rate is both level stationary and trend stationary under these three tests, which implies that the time trend is negligible. We also performed the KPSS test but do not report the results here. We examine with both the Bartlett and quadratic kernels and apply the Newey-West method to select the bandwidth parameter. For both kernels, at the 10 percent level of significance, the real exchange rate is stationary for Philippines and Singapore; and trend stationary for Korea.

Compared to the earlier studies, we found more evidence of PPP, in 5 out of 12 of our sample of countries. The difference in our findings from those of studies mentioned in the previous section can be due to the fact that we cover a longer period and use more efficient unit root tests. Unlike previous studies, we find more evidence of level stationarity. Doğanlar and Özmen (2000) and Achy (2003), using the ADF and PP tests, could not reject the unit root hypothesis against level stationarity and/or trend stationarity for any of the countries common to their study and this chapter. Hooi and Smyth (2007) find evidence of trend stationarity for India, Pakistan, Philippines and Thailand for the period 1990 to 2005. Bahmani-Oskooee, Kutan and Zhou (2009), based on the KPSS test, find that the real effective exchange rate is level stationary for Singapore and trend stationary for Bangladesh, Korea, Malaysia, Singapore and Thailand. In our case, the unit root hypothesis is not rejected against the alternative hypothesis of level or trend stationarity in any of the South Asian countries although the results for East- and Southeast Asian countries between earlier studies and ours are somewhat comparable.

4.4. Unit Root Test with Structural Break(s)

Perron (1989) argues that, in the presence of a structural break, the standard ADF test is biased towards the non-rejection of the null hypothesis. He provides a procedure that uses a modified Dickey-Fuller (DF) framework and includes dummy variables to account for one known structural break in the intercept or slope of the trend function, or in both of them. This procedure allows for a break under both the null and alternative hypotheses. Moreover, the test is invariant to the magnitude of the break, though the critical values of the test statistic depend on the location of the break. He also presents two versions of the model that differ in the treatment of the break dynamics: the IO model and the AO model. The IO model assumes that the breaks in the trend function occur gradually and follow the same dynamic path of the innovation, while the AO model assumes that the break occurs immediately with a sudden change in the trend function. The IO model for trending data with break in both intercept and trend is:

$$\Delta q_t = \alpha + \beta t + \theta DI_t(TB) + \eta DT_t(TB) + \xi D_t(TB) + (\rho - 1)q_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta q_{t-j} + \varepsilon_t$$

where, intercept break dummy $DI_t = \begin{cases} 1 & for\ t \geq TB \\ 0 & Otherwise \end{cases}$

trend break dummy $DT_t = \begin{cases} t - TB + 1 & for\ t \geq TB \\ 0 & Otherwise \end{cases}$

and one-time break dummy for intercept $D_t = \begin{cases} 1 & for\ t = TB \\ 0 & Otherwise \end{cases}$

For non-trending data with intercept break $\beta = \eta = 0$; for trending data with intercept break $\eta = 0$; and for trending data with break in trend $\theta = \xi = 0$.

For the AO model, in a two-step procedure, data is de-trended using OLS with the appropriate intercept and trend variables in the first step:

$$y_t = \alpha + \beta t + \theta DI_t(TB) + \eta DT_t(TB) + \tilde{y}_t$$

where $\tilde{y}_t$ is the residual from the first step. In the second step:

$$\Delta \tilde{y}_t = \sum_{i=0}^{p} \xi_i D_{t-i}(TB) + (\rho - 1)\tilde{y}_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta \tilde{y}_{t-j} + \varepsilon_t$$

In this specification $(p+1)$ one-time break dummy for intercept is needed to make the test statistics invariant to nuisance parameters. Both the IO and AO frameworks can accommodate multiple breaks.

Zivot and Andrews (ZA) (1992) develop a unit root test procedure where the structural break is determined endogenously by minimizing the Dickey-Fuller test statistic. This technique selects the break date that provides the most evidence against the unit root null and in favor of the alternative hypothesis of stationarity with a structural break. It adopts the IO type specification and does not include the one-time shift in the level under the alternative hypothesis. The critical values of the ZA test are different from Perron (1989) and are invariant to the location of the structural break. Perron and Vogelsang (1992) and Perron (1997) perform the unit root test for both the IO and AO specifications. As in ZA, these papers assume no break in the unit root null but include the one-time shift in the level under the alternative hypothesis. Lumsdaine and Papell (LP) (1997) extend the ZA test to accommodate two structural breaks.

Lee and Strazicich (LS) (2003) criticize the ZA and LP tests by arguing that rejection of the unit root null implies a rejection of the unit root without break rather than a rejection of the unit root. They, then, develop a two break Lagrange Multiplier based unit root test that has breaks in null and alternative hypotheses. Rejection of the null hypothesis unambiguously implies trend stationarity. This unit root test is invariant to the location of the break only if the break is in the intercept. If the break is in both the intercept and trend, the LM test is not invariant to the location of the break. Moreover, this test does not diverge in any of these two cases and does not show any systematic pattern of over-rejecting the null. Lee and Strazicich (2004) provide the critical values for the test under one structural break.

The ZA, LP and LS tests use the IO framework and examine the null hypothesis of unit root with a drift against the alternative of stationarity with a deterministic time trend. However, for PPP to hold, the real exchange rate has to be level stationary; stationarity around a deterministic trend would only provide support for the HBS hypothesis. Thus, it is equally, if not more, important to examine the null hypothesis of unit root without a drift against the alternative of stationarity with a drift. Also, because of the breaks, there can be an instantaneous change in the trend function, and this requires AO-type models. However, the empirical literature on the Asian countries do not address these two issues. This makes our examination of both level stationarity and trend stationarity under both AO and IO frameworks a major contribution to this literature. These models can be estimated using Perron's work, in different papers, on unit root tests in the presence of structural breaks. From now on, we refer to these two models as AO-Perron and IO-Perron. This chapter only examines structural shifts, i.e., break(s) in the constant of the trend function. The rationale for doing so is discussed in the introductory section.

All the tests used in this section choose the break endogenously at the point that gives the minimum test statistic, i.e. the maximum evidence against the null hypothesis. The breaks thus selected may not coincide with the known exogenous breaks that we expect for some countries. However, our results show that, in general, at least one selected break coincides with the Asian financial crisis or the most recent global economic crisis for the East and Southeast Asian countries. For the South Asian countries, the endogenously selected breaks are more diverse and rarely coincides with these two major crises. As the results of the unit root tests are affected by the number of augmenting lags, we choose lags based on three lag selection criteria, namely AIC[19], T test and a fixed number of lags. The details and rationale for these lag selection criteria are presented in the previous section. To avoid spurious breaks at the beginning and at the end of the sample period, we trim 10 percent of the data at both ends and select the breaks in the middle 80 percent region.

We start with the null hypothesis of the unit root against the alternative of level stationarity with break(s) in the constant for both the AO and IO models. The null hypothesis does not have a constant or trend, so that there is no question of allowing breaks under the null. The results of the AO-Perron and IO-Perron tests are given in Table 4.3 and Table 4.4, respectively. As expected, the T test tends to select more lags compared with the AIC. Allowing for 1 structural break in the AO model, the null cannot be rejected for any of the countries. However, when 2 breaks in the constant are allowed, the null hypothesis can be rejected for Japan, Taiwan and Thailand. The null hypothesis is rejected for Japan at the 10 percent level of significance, with 11 truncation lags based on the T test criterion. For Taiwan and Thailand, the null is rejected at the 5 percent level of significance. The relevant issue here is that the null is rejected when only 1 truncation lag is chosen by AIC for both countries. We are more conservative about the case of Thailand as the 1 truncation

[19] Selecting lag based on M-AIC is computationally demanding. Since M-AIC chose small number of lags for most countries in unit root analysis without structural break, for tests in the presence of breaks we simply use AIC.

lag may not be sufficient to remove serial correlation. For Taiwan, the unit root null is already rejected by MZ_t test. The IO model with 1 structural break rejects the unit root hypothesis only for Pakistan (at the 10 percent level of significance) with a fixed lag of 5. Allowing 2 breaks in the constant improves rejection of the null hypothesis significantly. The null is rejected for India, Japan and Thailand at the 10 percent level of significance; and for Taiwan at the 5 percent level of significance. For these four countries, more truncation lags are selected by the lag selection criteria under this specification.

The AO-Perron test with 1 break, similar to the test of level stationarity, does not reject the unit root hypothesis against the alternative of trend stationarity for any of the countries. With 2 structural breaks, the null hypothesis can be rejected for Taiwan and Thailand at the 5 percent level of significance when AIC selects only 1 augmenting lag (Table 4.5). To examine trend stationarity using the IO framework, we use ZA, LP and LS tests instead of the IO-Perron, so that we can compare our results with earlier studies. The ZA test, which allows 1 break, rejects the unit root hypothesis only for Japan; and the LP test, allowing 2 breaks, rejects the null for Taiwan and Thailand (Table 4.6). For these three countries, the null is rejected when several augmenting lags are selected by the lag selection criteria. ZA and LP tests are often criticized for over-rejecting the null as they do not allow breaks under the null hypothesis. Allowing breaks in the null, the LS test rejects the null hypothesis for Korea when 1 break is allowed; and for China and Korea when 2 breaks are allowed. In our analysis, ZA and LP tests do not reject the null for any countries for which the null is not rejected by the AO-Perron and IO-Perron tests of level stationarity. It is also possible to allow breaks under the null hypothesis in the AO models, as in Carrion-i-Silvestre, Kim and Perron (2009). However, we do not employ this test in this chapter as trend stationarity tests under AO framework, without allowing breaks under the null, do not reject null for any additional countries; so that not having breaks under the null is not a major issue here.

Based on the unit root analysis in the presence of structural shifts, the real exchange rate is stationary for Japan, Taiwan and Thailand. The real exchange rate in these three countries, along with Korea, are also trend stationary which implies that the trend is negligible. However, unit root tests without allowing structural breaks also find the real exchange rate to be stationary for Korea and Taiwan. Unit root tests in presence of breaks provide more support of stationarity for Japan (only the ADF test supported level stationarity for it) and provide new evidence of level stationarity for Thailand. We are not certain that the PPP holds for China, India and Pakistan although there are some evidence for it. The reasons are: (1) the null is rejected at the 10 percent level of significance, (2) among all the tests performed in this chapter, the null is rejected under only one specification, and (3) within that specification, the null is rejected under only one lag selection criteria. We experimented with different truncation lags but the unit root hypothesis is not rejected for these three countries.

Based on our results, PPP seems to hold more for relatively high income countries in our sample of countries. For example, in terms of per capita GDP, the top four high income countries are Singapore, Taiwan, Japan and Korea; and PPP holds in all of these countries. Other than Malaysia, PPP also seems to hold for countries with low and stable inflation (see Figure 3.2) and for countries with high volume of trade with the USA. This is consistent with the results of Alba and Papell (2007) as USD is the numeraire currency in our analysis, and thus, we are comparing PPP hypothesis vis-à-vis USA. Although China and India are major trading partners of USA, their inflation is more volatile and they intervene more in the foreign exchange market (for details, see a de facto classification of exchange rate regime by IMF, provided in Appendix 3.1).

As mentioned earlier, previous studies on our selected Asian countries examined only trend stationarity using the IO framework and generally found additional evidence for stationarity in the presence of structural breaks. Nusair (2003) follows Perron (1989) and allow a break in slope of the trend function for 6 Asian countries (Indonesia, Korea, Malaysia, Philippines, Singapore and Thailand) with quarterly data from 1973 to 1999. Based on the ADF, PP and KPSS tests, the real exchange rate is trend stationary for Indonesia, Korea and Thailand. The test with 1 known break rejects the unit root null for an additional country, Malaysia. Hooi and Smyth (2007) examine stationarity of the real exchange rate for 15 countries (including Bangladesh, China, India, Indonesia, Korea, Malaysia, Pakistan, Philippines, Singapore and Thailand). Using the ADF test, they are able to reject unit root null for India, Pakistan, Philippines and Thailand. In addition to that, using LS with 1 structural shift, they find support for PPP in India, Indonesia, Malaysia, Korea, Pakistan and Thailand. Allowing breaks in both constant and slope of the trend function, PPP seems to hold for Bangladesh, India, Indonesia, Korea, Pakistan, the Philippines and Thailand. Soon, Baharumshah and Ahn (2015) examine the PPP hypothesis for 13 Asian countries (including China, India, Indonesia, Malaysia, Pakistan, Philippines, Singapore, South Korea, Taiwan and

Thailand) for the period 1986 to 2010. Using the Narayan and Popp (2010)[20] test with 2 structural shifts, the unit root null is rejected for Taiwan while allowing breaks in both the constant and the slope of the trend function, the null is rejected for Pakistan.

The empirical literature also examine the PPP hypothesis by allowing breaks in the cointegration method. Gregory and Hansen (1996) extend the residual based cointegration test of Engle and Granger (1987) to allow one endogenously determined break in the cointegrating vector. Johansen et al. (2000), on the other hand, extend the Johansen (1988) procedure to allow for (up to) two pre-determined break. The critical values for this test depend on the number of non-stationary relations, the location of break points and the trend specification. Nusair (2008) employs this procedure with two known breaks for 9 Asian countries (including India, Indonesia, Korea, Malaysia, Philippines, Singapore, Sri Lanka and Thailand). Using quarterly data from 1973 to 2005, the null of no cointegration is rejected for all countries. However, when breaks are not allowed, PPP seems to hold for India, Korea, Pakistan and Singapore.

Im, Lee & Tieslau (2005) propose a LM statistic based panel unit-root test with multiple structural shifts. Their simulation shows that this test is robust to the presence of structural shifts and is more powerful than the IPS test. Hooi and Smyth (2007) employ this procedure on a panel of 15 Asian countries for the period 1990 to 2005 and find evidence of PPP, but do not find evidence of PPP when structural breaks are not included. Carrion-i-Silvestre, del Barrio-Castro and Lopez-Bazo (2005) develop a test within the IPS framework, where the test statistic for each country is derived using the KPSS test with multiple breaks at different unknown dates. Following this procedure Chang, Li, Lu and Lee (2011) do not find evidence of PPP for 10 East-Asian countries for the period 1987 to 2005. Westerlund (2006) proposes a LM statistic to examine panel cointegration allowing for endogenously determined multiple breaks in level and trend. Narayan (2010) employs this approach with 1 endogenously determined break in level for 6 Asian countries (India, Malaysia, Pakistan, Philippines, Sri Lanka and Thailand) with annual data from 1967 to 2002. Using a bivariate model of the nominal exchange rate and relative prices, he finds strong

[20] This procedure is based on IO framework and allows breaks under the null hypothesis. The test examines trend stationarity and allows breaks in both constant and slope of the trend function. It endogenously selects the breaks, based on a sequential procedure. It has same specification of Perron (1989) IO model with lagged dummies of constant and trend break.

evidence of cointegration in the selected countries. In contrast, without breaks, the panel cointegration test of Pedroni (19999) provide only weak evidence of cointegration. Thus, panel studies in the presence of structural break(s) provide relatively more support for the PPP hypothesis although the results of Taylor and Sarno (1998) are still a concern for such studies.

4.5. Conclusions

This chapter examines the PPP hypothesis for 12 East, Southeast and South Asian countries that are at different stages of economic development. We use quarterly data for the period 1974 to 2016 and the CPI-based bilateral real exchange rates vis-à-vis the USD to examine the unit root hypothesis. We prefer to use the USD-based exchange rates as most of our selected countries invoice trade in USD and hold substantial amounts of this currency as foreign exchange reserves.

The chapter makes a contribution to the existing literature by examining the level stationarity of the real exchange rate in the presence of multiple structural shifts for both AO and IO model, which the previous literature had failed to do for our selected countries. It make another contribution by providing a thoroughly investigated unit root test (without accommodating structural breaks) and examination of trend stationarity in the presence of multiple structural breaks for a longer time span.

Compared to the earlier studies, we find more support for the PPP hypothesis, in 5 (Japan, Korea, Philippines, Singapore and Taiwan) out of 12 countries, using unit root tests without allowing structural breaks. The improved support for PPP can be due to the fact that we have used longer time periods and more efficient tests. Allowing up to two structural shifts, we find support for PPP in one additional country, Thailand. Also, PPP seems to hold more for relatively high income countries.

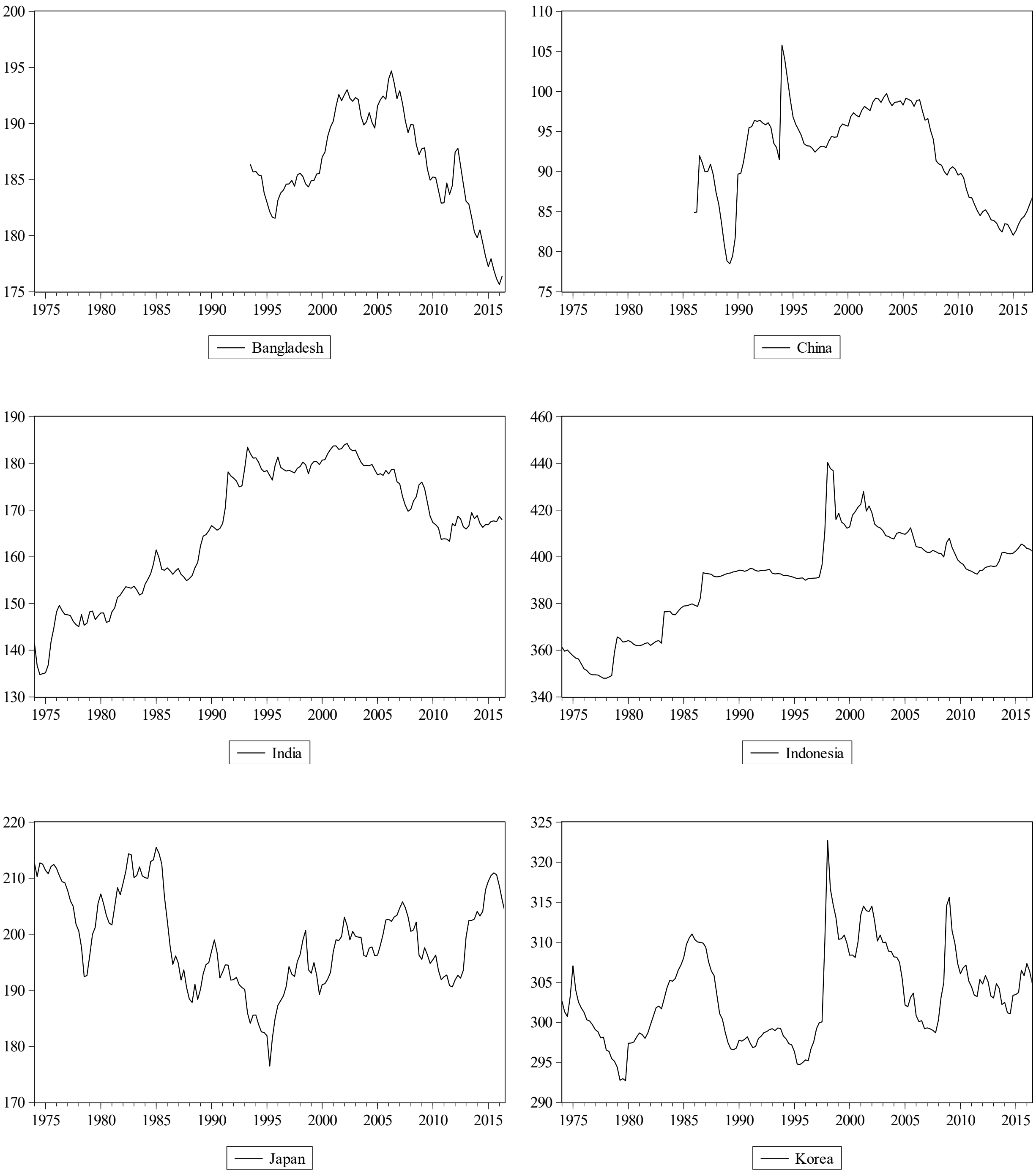
Further, our study finds little evidence of trend stationarity in these countries, so that there is no support for the HBS hypothesis. We used both the AO and IO models but did not find any major difference in terms of rejecting the null hypothesis in their results.

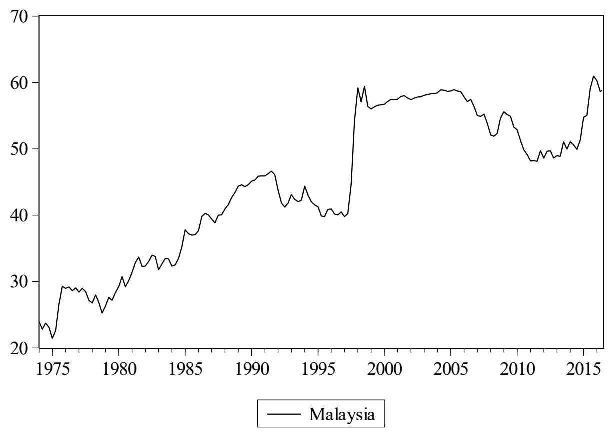
Our results differ from those of Bahmani-Oskooee et al. (2009) and Hooi and Smyth (2007), who find evidence of trend stationarity in South Asian countries. For East and Southeast Asian countries our findings are somewhat comparable to those in the earlier literature. Bahmani-Oskooee et al (2009), using the KPSS test and real effective exchange rates, find that the real effective exchange rate is trend stationary in Bangladesh, Korea, Malaysia, Singapore and Thailand. Hooi and Smyth (2007) applies the LS test on real exchange rate with data from 1990 to 2005. Based on their results, PPP seems to hold for Bangladesh, India, Indonesia, Korea, Pakistan, the Philippines and Thailand but does not hold for China, Malaysia and Singapore.

Lack of support for PPP can be due to measurement issues related to the price index. Most empirical studies use the CPI to represent the general price level of a country. A significant weight in the construction of the CPI comes from non-traded goods and services that preclude arbitrage. WPI or PPI has a higher proportion of traded goods and are able to provide more support for PPP, as found in Ito (1997) and Lee (1999). However, we cannot verify this for our selected countries as long span data on WPI or PPI are not available for most of our selected countries. Instead of using the bilateral real exchange rate, the PPP hypothesis can also be tested with the real effective exchange rate. However, data for some of our selected countries are only available after 1994.

There are usually significant transaction costs associated with the traded goods, including transportation costs, tariffs and other non-tariff barriers (such as quota) that make arbitrage of goods between countries costly. This can be a reason for lack of support for the PPP in South Asian countries and in China. 'Pricing to market' behaviour of exporting firms, as discussed in Krugman (1987), can also be a source of deviations from PPP for countries like China and India as they have large economies. In addition to these factors, Bahmani-Oskooee and Hegerty (2009) list a number of other factors for the breakdown of PPP, including natural resource endowments, military expenditures, corruption, smuggling, etc.

Figure 4.1: Real Exchange Rates (vis-à-vis the USA)

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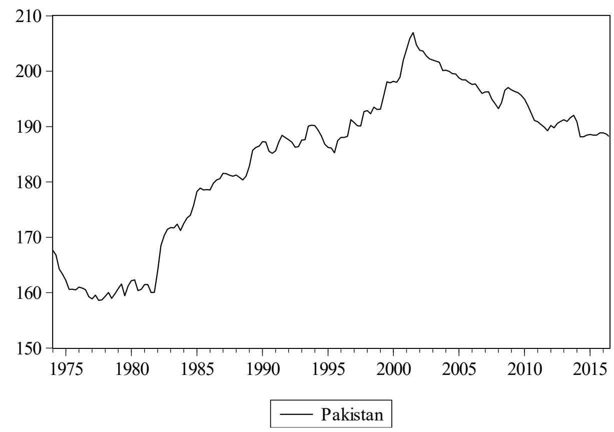
1975 1980 1985 1990 1995 2000 2005 2010 2015
Malaysia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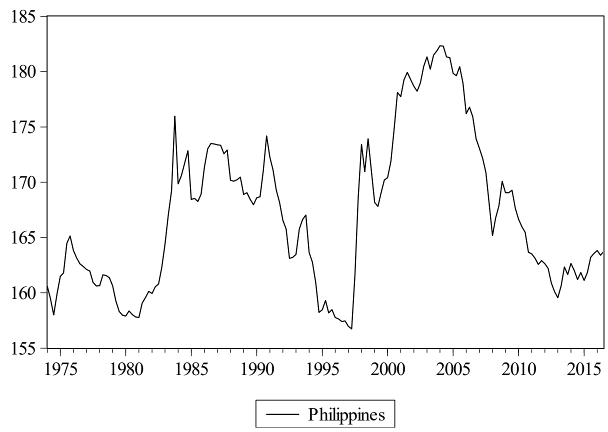
Pakistan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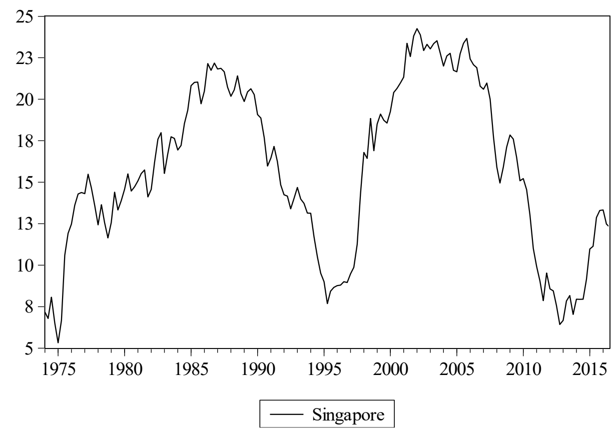
Philippines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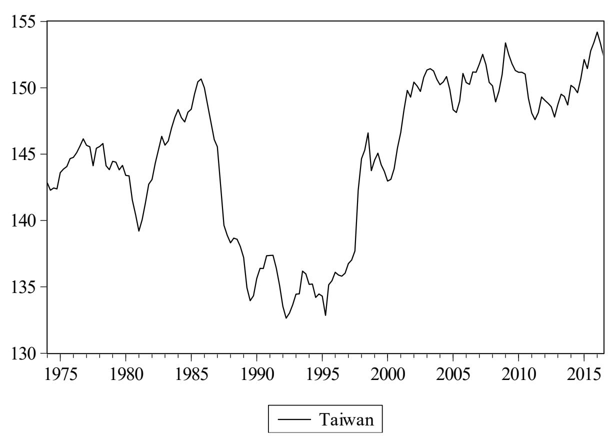
Singapore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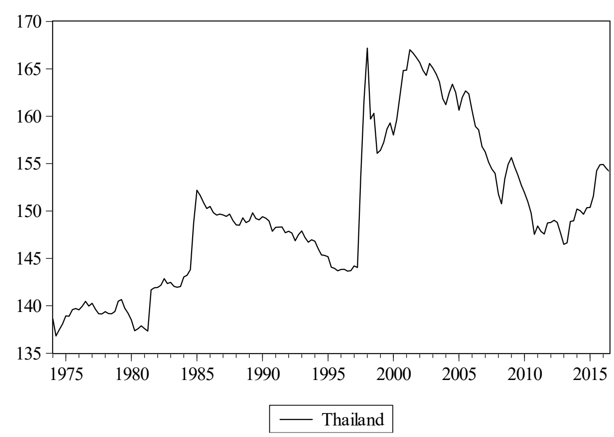
1975 1980 1985 1990 1995 2000 2005 2010 2015
Taiwan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1975 1980 1985 1990 1995 2000 2005 2010 2015
Thailand

Table 4.2: Test Statistics of Unit Root Test (no structural break)

Country	Lag Selection Criteria	ADF		DF-GLS		MZt	
		Constant	Constant and Trend	Constant	Constant and Trend	Constant	Constant and Trend
Bangladesh	Ttest	0.03 (10)	-0.24 (10)	-0.26 (10)	-0.78 (10)	-0.36 (10)	-0.95 (10)
	M-AIC	0.03 (10)	-0.24 (10)	-0.26 (10)	-0.78 (10)	-0.36 (10)	-0.95 (10)
	Fixed	-0.59 (5)	-1.00 (5)	-0.71 (5)	-1.25 (5)	-0.97 (5)	-1.41 (5)
China	Ttest	-1.85 (0)	-2.14 (0)	-1.35 (0)	-1.50 (0)	-1.32 (0)	-1.45 (0)
	M-AIC	-1.85 (0)	-2.14 (0)	-1.35 (0)	-1.50 (0)	-1.32 (0)	-1.45 (0)
	Fixed	-1.82 (5)	-2.01 (5)	-1.41 (5)	-1.61 (5)	-1.52 (5)	-1.80 (5)
India	Ttest	-1.95 (7)	-0.84 (7)	-0.15 (7)	-0.69 (7)	-0.17 (7)	-0.76 (7)
	M-AIC	-1.95 (7)	-0.84 (7)	-0.44 (1)	-1.20 (1)	-0.32 (1)	-1.05 (1)
	Fixed	-2.49 (5)	-1.01 (5)	-0.14 (5)	-0.64 (5)	-0.06 (5)	-0.55 (5)
Indonesia	Ttest	-1.64 (6)	-1.43 (6)	-0.29 (6)	-1.51 (6)	-0.27 (6)	-1.57 (6)
	M-AIC	-1.61 (3)	-1.66 (3)	-0.41 (3)	-1.72 (3)	-0.41 (3)	-1.75 (3)
	Fixed	-1.71 (5)	-1.72 (5)	-0.45 (5)	-1.78 (5)	-0.45 (5)	-1.89 (5)
Japan	Ttest	-1.98 (13)	-1.65 (13)	-0.93 (13)	-1.30 (13)	-1.05 (13)	-1.60 (13)
	M-AIC	-2.75[c] (3)	-2.59 (3)	-1.54 (3)	-2.16 (3)	-1.61 (3)	-2.36 (3)
	Fixed	-2.50 (5)	-2.29 (5)	-1.35 (5)	-1.87 (5)	-1.44 (5)	-2.16 (5)
Korea	Ttest	-2.95[b] (3)	-3.27[c] (3)	-2.95[a] (3)	-3.15[b] (3)	-3.21[a] (3)	-3.44[a] (3)
	M-AIC	-2.62[c] (4)	-2.81 (2)	-2.63[a] (4)	-2.75[c] (2)	-2.80[a] (4)	-2.81[c] (2)
	Fixed	-2.81[c] (5)	-3.13 (5)	-2.80[a] (5)	-3.01[b] (5)	-3.31[a] (5)	-3.60[a] (5)
Malaysia	Ttest	-1.51 (1)	-2.57 (1)	0.32 (1)	-2.40 (1)	0.30 (1)	-2.37 (1)
	M-AIC	-1.51 (1)	-2.57 (1)	0.32 (1)	-2.40 (1)	0.30 (1)	-2.37 (1)
	Fixed	-1.43 (5)	-2.28 (5)	0.53 (5)	-0.27 (5)	0.53 (5)	-2.23 (5)
Pakistan	Ttest	-2.43 (11)	-0.98 (11)	-0.85 (11)	-1.74 (11)	-0.68 (11)	-1.94 (11)
	M-AIC	-1.34 (1)	-0.37 (2)	-0.34 (1)	-0.84 (2)	-0.31 (1)	-0.81 (2)
	Fixed	-1.89 (5)	-0.19 (5)	-0.35 (5)	-0.88 (5)	-0.17 (5)	-0.73 (5)
Philippines	Ttest	-2.83[c] (12)	-2.93 (12)	-1.82[c] (11)	-2.85[c] (12)	-2.44[b] (11)	-4.44[a] (11)
	M-AIC	-2.15 (1)	-2.05 (1)	-1.59 (1)	-1.96 (1)	-1.59 (1)	-1.95 (1)
	Fixed	-1.83 (5)	-1.73 (5)	-1.34 (5)	-1.68 (5)	-1.42 (5)	-1.80 (5)
Singapore	Ttest	-2.51 (11)	-2.51 (11)	-1.24 (4)	-2.15 (11)	-1.28 (4)	-2.80[c] (11)
	M-AIC	-2.20 (1)	-2.21 (1)	-0.97 (1)	-1.40 (1)	-0.97 (1)	-1.35 (1)
	Fixed	-2.58[c] (5)	-2.64 (5)	-1.18 (5)	-1.72 (5)	-1.38 (5)	-1.90 (5)
Taiwan	Ttest	-1.41 (13)	-2.10 (13)	-1.33 (13)	-1.89 (13)	-1.95[c] (13)	-2.68[c] (13)
	M-AIC	-1.42 (1)	-1.83 (1)	-1.29 (1)	-1.77 (1)	-1.30 (1)	-1.78 (1)
	Fixed	-1.42 (5)	-1.91 (5)	-1.31 (5)	-1.81 (5)	-1.48 (5)	-1.98 (5)
Thailand	Ttest	-1.68 (3)	-1.78 (3)	-0.60 (3)	-1.75 (3)	-0.62 (3)	-1.80 (3)
	M-AIC	-1.68 (3)	-1.78 (3)	-0.60 (3)	-1.75 (3)	-0.62 (3)	-1.80 (3)
	Fixed	-1.67 (5)	-1.77 (5)	-0.58 (5)	-1.74 (5)	-0.63 (5)	-1.86 (5)

[a, b] and [c] represents 1%, 5% and 10% level of significance, respectively. Figures in the parentheses shows the number of lags included.

ADF critical values for the constant only case are -3.47, -2.88 and -2.58 at 1%, 5% and 10% level of significance, respectively; and for constant and trend case are -4.01, -3.44 and -3.12 at 1%, 5% and 10% level of significance, respectively.

DF-GLS critical values for the constant only case are -2.58, -1.94 and -1.61 at 1%, 5% and 10% level of significance, respectively; and for constant and trend case are -3.51, -2.97 and -2.68 at 1%, 5% and 10% level of significance, respectively. Source: Elliott, Rothenberg and Stock (1996).

Critical values for MZt at 1%, 5% and 10% level of significance are -2.58, -1.98 and -1.62, respectively for the constant only case; and 1%, 5% and 10% level of significance are -3.42, -2.91 and -2.62, respectively for the constant and trend case. Source: Ng and Perron (2001), Table 1.

Table 4.3: AO-Perron (Level Stationarity, Break in Constant)

Country	Lag Selection Criteria	1 Structural Break		2 Structural Break	
		Test Statistic	Time Break	Test Statistic	Time Break
Bangladesh	Ttest	-3.20 (8)	2009:4	-3.91 (8)	2003:4, 2009:4
	AIC	-3.20 (8)	2009:4	-4.04 (9)	2003:3; 2009:3
	Fixed	-2.71 (5)	2010:3	-3.47 (5)	1997:2; 2010:3
China	Ttest	-3.08 (0)	2007:1	-4.54 (4)	1993:4; 2009:2
	AIC	-3.08 (0)	2007:1	-4.54 (4)	1993:4; 2009:2
	Fixed	-2.80 (5)	1992:3	-4.51 (5)	1992:3; 2008:4
India	Ttest	-2.60 (7)	1985:4	-3.65 (7)	1988:4; 2007:1
	AIC	-3.48 (1)	1987:4	-4.75 (1)	1987:4; 2006:1
	Fixed	-2.87 (5)	1986:3	-4.24 (5)	1987:2; 2007:2
Indonesia	Ttest	-3.19 (3)	1982:1	-4.55 (3)	1982:1; 1997:1
	AIC	-3.19 (3)	1982:1	-5.13 (1)	1982:3; 1997:3
	Fixed	-3.11 (5)	1981:3	-4.26 (5)	1981:3; 1996:3
Japan	Ttest	-3.90 (8)	1982:4	-5.26[c] (11)	1987:2; 2000:4
	AIC	-3.87 (3)	1984:1	-5.19 (4)	1983:4; 1994:1
	Fixed	-3.59 (5)	1983:3	-4.64 (5)	1983:3; 1993:4
Korea	Ttest	-3.70 (12)	1993:1	-4.67 (3)	1996:3; 2003:2
	AIC	-3.84 (3)	1995:2	-4.83 (1)	1997:1; 2002:4
	Fixed	-3.76 (5)	1994:4	-4.31 (5)	1996:1; 2003:1
Malaysia	Ttest	-3.36 (1)	1996:3	-5.16 (1)	1984:1; 1996:4
	AIC	-3.36 (1)	1996:3	-5.16 (1)	1984:1; 1996:4
	Fixed	-3.16 (5)	1995:3	-4.75 (5)	1983:1; 1995:4
Pakistan	Ttest	-2.45 (11)	1992:3	-2.85 (11)	1980:4; 1992:3
	AIC	-4.01 (1)	1981:2	-4.19 (1)	1981:2; 1995:1
	Fixed	-3.82 (5)	1980:2	-3.74 (5)	1980:2; 2012:4
Philippines	Ttest	-3.21 (12)	1994:1	-3.93 (12)	1994:1; 2007:4
	AIC	-3.06 (13)	1995:3	-4.31 (1)	1996:4; 2006:1
	Fixed	-2.32 (5)	1995:4	-3.74 (5)	1995:4; 2007:2
Singapore	Ttest	-3.65 (4)	2008:1	-4.39 (4)	1994:1; 2007:4
	AIC	-2.90 (1)	2006:4	-3.81 (1)	1996:4; 2006:4
	Fixed	-3.37 (5)	2007:4	-4.17 (5)	1993:4; 2007:4
Taiwan	Ttest	-2.86 (1)	1997:1	-5.71[b] (1)	1986:3; 1997:1
	AIC	-2.86 (1)	1997:1	-5.71[b] (1)	1986:3; 1997:1
	Fixed	-2.87 (5)	1993:4	-4.70 (5)	1984:3; 1996:1
Thailand	Ttest	-3.08 (13)	1993:4	-3.59 (13)	1993:4; 2006:2
	AIC	-3.08 (13)	1993:4	-5.50[b] (1)	1996:4; 2005:2
	Fixed	-2.85 (5)	1995:4	-4.35 (5)	1995:4; 2005:1

[a, b] and [c] represents the rejection of null hypothesis at 1%, 5% and 10% level of significance, respectively. Number of lags included are shown in parenthesis.

Critical values at 1%, 5% and 10% level of significance are -4.95, -4.44 and -4.19, respectively for both AO and IO model with 1 structural break. Source: Perron and Vogelsang (1992)

Critical values at 1%, 5% and 10% level of significance are -5.96, -5.49 and -5.24, respectively for both AO and IO model with 2 structural break. Source: Clemente, Montanes and Reyes (1998).

Table 4.4: IO-Perron (Level Stationarity, Break in Constant)

Country	Lag Selection Criteria	1 Structural Break		2 Structural Break	
		Test Statistic	Time Break	Test Statistic	Time Break
Bangladesh	Ttest	-2.40 (10)	2011:4	-3.41 (10)	1998:3; 2011:4
	AIC	-2.40 (10)	2011:4	-3.72 (8)	1998:3; 2012:1
	Fixed	-2.77 (5)	2011:4	-3.70 (5)	1998:3; 2012:1
China	Ttest	-3.07 (0)	2007:1	-4.78 (4)	1993:4; 2007:1
	AIC	-3.07 (0)	2007:1	-4.78 (4)	1993:4; 2007:1
	Fixed	-3.02 (5)	1993:4	-4.67 (5)	1993:4; 2009:3
India	Ttest	-2.98 (7)	1987:4	-4.80 (7)	1988:3; 2006:2
	AIC	-3.50 (1)	1987:4	-4.83 (1)	1988:1; 2006:2
	Fixed	-3.27 (5)	1988:1	-5.28[c] (5)	1988:3; 2006:2
Indonesia	Ttest	-3.22 (3)	1982:4	-4.06 (3)	1982:4; 1997:1
	AIC	-3.22 (3)	1982:4	-4.06 (3)	1982:4; 1997:4
	Fixed	-3.20 (5)	1982:4	-4.11 (5)	1982:4; 1997:1
Japan	Ttest	-3.81 (3)	1984:4	-5.46[c] (11)	1990:1; 1999:3
	AIC	-3.81 (3)	1984:4	-5.02 (3)	1984:4; 1995:1
	Fixed	-3.59 (5)	1984:4	-4.73 (5)	1984:4; 2000:1
Korea	Ttest	-3.91 (3)	1997:2	-4.87 (3)	1997:2; 2003:1
	AIC	-3.91 (3)	1997:2	-4.87 (3)	1997:2; 2003:1
	Fixed	-3.78 (5)	1997:2	-4.71 (5)	1997:2; 2003:3
Malaysia	Ttest	-3.40 (1)	1997:1	-5.23 (1)	1984:1; 1997:1
	AIC	-3.40 (1)	1997:1	-5.23 (1)	1984:1; 1997:1
	Fixed	-3.39 (5)	1997:1	-5.17 (5)	1984:1; 1997:1
Pakistan	Ttest	-3.92 (11)	1981:3	-4.30 (11)	1981:3; 1995:2
	AIC	-4.32[c] (5)	1981:3	-4.47 (5)	1981:3; 2010:2
	Fixed	-4.32[c] (5)	1981:3	-4.47 (5)	1981:3; 2010:2
Philippines	Ttest	-3.37 (12)	1980:4	-4.37 (12)	1997:1; 2006:1
	AIC	-2.55 (1)	1997:1	-4.39 (1)	1997:1; 2006:1
	Fixed	-2.26 (5)	1997:1	-4.09 (5)	1997:1; 2006:1
Singapore	Ttest	-3.66 (4)	2008:4	-4.47 (4)	1997:1; 2007:1
	AIC	-2.92 (1)	2007:1	-3.95 (1)	1997:1; 2007:1
	Fixed	-3.40 (5)	2008:4	-4.19 (5)	1997:1; 2007:1
Taiwan	Ttest	-3.23 (12)	1999:4	-5.47[c] (12)	1986:4; 1997:2
	AIC	-2.91 (1)	1997:1	-5.90[b] (1)	1986:4; 1997:2
	Fixed	-2.91 (5)	1997:1	-5.73[b] (5)	1986:4; 1997:2
Thailand	Ttest	-3.47 (13)	1997:1	-5.31[c] (8)	1997:1; 2005:2
	AIC	-2.97 (3)	1997:1	-5.29[c] (3)	1997:1; 2005:2
	Fixed	-2.96 (5)	1997:1	-5.28[c] (5)	1997:1; 2005:3

[a, b] and [c] represents the rejection of null hypothesis at 1%, 5% and 10% level of significance, respectively. Number of lags included are shown in parenthesis.

Critical values at 1%, 5% and 10% level of significance are -4.95, -4.44 and -4.19, respectively for both AO and IO model with 1 structural break. Source: Perron and Vogelsang (1992)

Critical values at 1%, 5% and 10% level of significance are -5.96, -5.49 and -5.24, respectively for both AO and IO model with 2 structural break. Source: Clemente, Montanes and Reyes (1998).

Table 4.5: AO-Perron (Trend Stationarity, Break in Constant)

Country	Lag Selection Criteria	1 Structural Break		2 Structural Break	
		Test Statistic	Time Break	Test Statistic	Time Break
Bangladesh	Ttest	-3.22 (8)	2009:4	-4.49 (4)	1998:3; 2010:4
	AIC	-3.22 (8)	2009:4	-4.49 (4)	1998:3; 2010:4
	Fixed	-2.74 (5)	2010:3	-3.83 (5)	1998:2; 2010:3
China	Ttest	-3.88 (0)	1989:2	-4.56 (4)	1992:4; 2008:3
	AIC	-3.54 (4)	1993:4	-4.56 (4)	1992:4; 2008:3
	Fixed	-3.20 (5)	2007:4	-4.66 (5)	1992:3; 2008:2
India	Ttest	-2.38 (7)	2004:3	-3.73 (7)	1988:4; 2004:3
	AIC	-3.49 (1)	1987:4	-4.77 (1)	1988:2; 2005:4
	Fixed	-2.68 (5)	1987:2	-3.95 (5)	1988:4; 2005:1
Indonesia	Ttest	-3.39 (3)	2004:3	-4.60 (3)	1982:1; 1997:1
	AIC	-3.39 (3)	2004:3	-5.04 (1)	1982:3; 1997:3
	Fixed	-3.38 (5)	2004:1	-4.29 (5)	1981:3; 1996:3
Japan	Ttest	-4.46 (11)	1986:4	-5.26 (11)	1987:2; 2000:4
	AIC	-4.60 (3)	1984:1	-5.27 (3)	1984:1; 1994:2
	Fixed	-4.19 (5)	1983:3	-4.85 (5)	1984:1; 1993:4
Korea	Ttest	-4.14 (3)	1996:3	-5.39 (3)	1982:1; 1996:3
	AIC	-4.25 (1)	1997:1	-5.46 (1)	1981:3; 1997:1
	Fixed	-3.90 (5)	1996:1	-5.02 (5)	1981:3; 1996:1
Malaysia	Ttest	-3.84 (7)	2007:1	-5.21 (1)	1996:4; 2006:1
	AIC	-3.95 (1)	2006:1	-5.21 (1)	1996:4; 2006:1
	Fixed	-3.49 (5)	2005:1	-4.76 (5)	1995:4; 2006:1
Pakistan	Ttest	-2.62 (11)	2007:1	-3.09 (11)	1985:3; 2007:1
	AIC	-2.84 (1)	1981:2	-4.23 (1)	1981:2; 2008:3
	Fixed	-2.00 (5)	1980:2	-3.06 (5)	1980:2; 2008:3
Philippines	Ttest	-3.50 (12)	2008:3	-4.13 (12)	1987:3; 2007:2
	AIC	-3.01 (1)	2005:1	-5.19 (1)	1982:1; 1996:4
	Fixed	-2.52 (5)	2007:2	-3.89 (5)	1981:1; 1997:3
Singapore	Ttest	-3.49 (4)	2007:4	-4.39 (4)	1990:1; 2007:4
	AIC	-3.01 (1)	2006:4	-4.05 (1)	1989:2; 2006:4
	Fixed	-3.33 (5)	2007:3	-4.40 (5)	1989:4; 2007:3
Taiwan	Ttest	-3.47 (12)	1982:3	-4.75 (12)	1988:1; 1997:3
	AIC	-3.75 (1)	1985:2	-6.17[b] (1)	1986:3; 1997:1
	Fixed	-3.48 (5)	1984:2	-4.88 (5)	1985:3; 1996:1
Thailand	Ttest	-3.13 (3)	2004:4	-5.24 (3)	1996:2; 2005:3
	AIC	-3.74 (1)	2005:2	-6.16[b] (1)	1983:4; 1996:4
	Fixed	-2.99 (5)	2004:1	-4.80 (5)	1995:4; 2005:1

[a, b] and [c] represents the rejection of null hypothesis at 1%, 5% and 10% level of significance, respectively. Number of lags included are shown in parenthesis.

1 Structural Break: Critical values at 1%, 5% and 10% significance level are -5.61, -5.02 and -4.72, respectively. 2 Structural Break: Critical values at 1%, 5% and 10% significance level are -6.45, -5.96 and -5.69, respectively. Source: Papell and Prodan (2006).

Table 4.6: Zivot-Andrews (ZA) Test and Lumsdaine-Papell (LP) Test

(Trend Stationarity, Break in Constant)

Country	Lag Selection Criteria	ZA Test (1 Structural Break)		LP Test (2 Structural Break)	
		Test Statistic	Break Period	Test Statistic	Break Period
Bangladesh	Ttest	-2.31 (10)	2000:1	-3.45 (10)	1999:4, 2012:1
	AIC	-2.31 (10)	2000:1	-3.45 (10)	1999:4, 2012:1
	Fixed	-2.95 (5)	2012:2	-4.12 (5)	1999:4, 2012:1
China	Ttest	-3.38 (0)	2007:3	-3.62 (0)	1990:4; 2007:2
	AIC	-3.38 (0)	2007:3	-3.62 (0)	1990:4; 2007:2
	Fixed	-3.35 (5)	2007:3	-3.73 (5)	1993:4; 2007:2
India	Ttest	-3.26 (7)	1989:1	-4.45 (7)	1988:4, 2006:3
	AIC	-3.57 (1)	1988:3	-4.84 (1)	1988:4, 2006:3
	Fixed	-3.59 (5)	1989:1	-4.75 (5)	1988:4, 2006:3
Indonesia	Ttest	-3.08 (6)	2005:4	-4.31 (6)	1983:1, 1997:2
	AIC	-3.39 (3)	2005:4	-4.57 (3)	1983:1, 1997:2
	Fixed	-3.45 (5)	2005:4	-4.56 (5)	1983:1, 1997:2
Japan	Ttest	-3.93 (13)	1985:4	-4.33 (13)	1985:3, 1992:4
	AIC	-4.84[b] (3)	1985:4	-5.10 (3)	1985:3, 1992:2
	Fixed	-4.53 (5)	1985:4	-4.83 (5)	1985:3, 1992:2
Korea	Ttest	-4.21 (3)	1997:4	-5.60 (3)	1981:4, 1997:3
	AIC	-4.21 (3)	1997:4	-5.60 (3)	1981:4, 1997:3
	Fixed	-4.08 (5)	1997:4	-5.52 (5)	1981:4, 1997:3
Malaysia	Ttest	-3.95 (1)	2006:4	-5.22 (1)	1984:2, 1997:2
	AIC	-3.95 (1)	2006:4	-5.22 (1)	1984:2, 1997:2
	Fixed	-3.64 (5)	2006:4	-5.16 (5)	1984:2, 1997:2
Pakistan	Ttest	-2.13 (11)	2008:4	-2.97 (11)	2002:4, 2009:4
	AIC	-2.71 (2)	1982:1	-4.22 (2)	1981:4, 1995:3
	Fixed	-2.41 (5)	1982:1	-4.02 (5)	1981:4, 1995:3
Philippines	Ttest	-3.70 (12)	2009:4	-4.54 (12)	1990:4, 2009:3
	AIC	-3.03 (1)	1997:3	-5.64 (1)	1983:3, 1997:2
	Fixed	-2.73 (5)	2006:4	-5.45 (5)	1983:3, 1997:2
Singapore	Ttest	-3.37 (11)	1997:3	-4.33 (11)	1989:4, 2009:2
	AIC	-3.84 (4)	1997:3	-5.12 (4)	1981:4, 1997:2
	Fixed	-3.59 (5)	1997:3	-4.89 (5)	1981:4, 1997:2
Taiwan	Ttest	-4.33 (12)	1986:2	-6.10[c] (12)	1987:1, 1997:3
	AIC	-3.92 (1)	1986:2	-6.34[b] (1)	1987:1, 1997:3
	Fixed	-3.85 (5)	1986:2	-6.22[b] (5)	1987:1, 1997:3
Thailand	Ttest	-3.43 (13)	1997:3	-5.98[c] (13)	1984:2, 1997:2
	AIC	-3.30 (3)	2005:4	-5.91[c] (3)	1984:2, 1997:2
	Fixed	-3.23 (5)	2006:1	-5.91[c] (5)	1984:2, 1997:2

[a, b] and [c] represents the rejection of null hypothesis at 1%, 5% and 10% level of significance, respectively. Number of lags included are shown in parenthesis.

Critical values for ZA test at 1%, 5% and 10% level of significance are -5.34, -4.80 and -4.58, respectively. Source: Zivot and Andrews (1992)

Critical values for LP test at 1%, 5% and 10% level of significance are -6.74, -6.16 and -5.89, respectively. Source: Lumsdaine and Papell (1997)

Table 4.7: Lee and Strazicich (LS) Test (Trend Stationarity, Break in Constant)

Country	Lag Selection Method	1 Structural Break		2 Structural Breaks	
		Test Statistic	Break Period	Test Statistic	Break Period
Bangladesh	Ttest	-2.44 (8)	2005:2	-2.67 (8)	1997:4; 2005:2
	AIC	-2.44 (8)	2005:2	-2.67 (8)	1997:4; 2005:2
	Fixed	-2.06 (5)	1996:4	-2.29 (5)	1996:4; 2012:4
China	Ttest	-2.82 (5)	1993:4	-3.79^{c} (4)	1993:4; 2007:4
	AIC	-2.82 (5)	1993:4	-3.01 (5)	1993:4; 2007:4
	Fixed	-2.82 (5)	1993:4	-3.01 (5)	1993:4; 2007:4
India	Ttest	-1.27 (7)	2011:4	-1.34 (7)	1992:3; 2011:4
	AIC	-2.00 (1)	1991:2	-2.16 (1)	1991:2; 2011:4
	Fixed	-1.50 (5)	2011:4	-1.59 (5)	1988:4; 2011:4
Indonesia	Ttest	-2.42 (13)	1998:3	-2.68 (13)	1986:3; 2001:4
	AIC	-2.21 (3)	1998:4	-2.58 (3)	1983:1; 1998:4
	Fixed	-2.17 (5)	1998:4	-2.49 (5)	1983:1; 1998:4
Japan	Ttest	-3.02 (8)	1982:4	-3.31 (8)	1982:4; 1990:3
	AIC	-2.91 (3)	1987:3	-3.37 (4)	1985:3; 1995:3
	Fixed	-2.75 (5)	1985:3	-3.00 (5)	1985:3; 1990:3
Korea	Ttest	-3.44^{c} (6)	1997:3	-3.87^{b} (6)	1979:4; 1997:3
	AIC	-3.28^{c} (3)	1996:2	-3.51^{c} (3)	1979:4; 1997:3
	Fixed	-3.19 (5)	2009:1	-3.48 (5)	1979:4; 1997:3
Malaysia	Ttest	-2.75 (1)	1998:2	-2.94 (1)	1998:2; 2011:4
	AIC	-2.75 (1)	1998:2	-2.94 (1)	1998:2; 2011:4
	Fixed	-2.25 (5)	2007:4	-3.36 (5)	2007:3; 2011:4
Pakistan	Ttest	-2.36 (11)	2001:4	-2.47 (11)	1983:3; 2001:4
	AIC	-2.36 (11)	2001:4	-2.47 (11)	1983:3; 2001:4
	Fixed	-1.61 (5)	2001:4	-1.74 (5)	1979:4; 2001:4
Philippines	Ttest	-3.11 (12)	1993:1	-3.27 (12)	1993:1; 1998:3
	AIC	-2.44 (2)	2007:4	-3.27 (12)	1993:1; 1998:3
	Fixed	-1.90 (5)	1983:3	-2.12 (5)	1983:3; 1997:3
Singapore	Ttest	-2.50 (11)	1998:3	-2.62 (11)	1983:1; 1998:3
	AIC	-2.35 (4)	1998:3	-2.46 (4)	1998:3; 2007:3
	Fixed	-2.18 (5)	1998:3	-2.36 (5)	1998:3; 2007:3
Taiwan	Ttest	-2.57 (12)	2008:4	-2.76 (12)	1995:3; 2008:4
	AIC	-1.98 (1)	1995:3	-2.09 (1)	1987:2; 1997:3
	Fixed	-2.05 (5)	1997:4	-2.19 (5)	1989:1; 1997:3
Thailand	Ttest	-2.68 (13)	1998:3	-2.80 (13)	1998:1; 2001:4
	AIC	-2.80 (1)	1998:2	-2.97 (1)	1998:2; 2004:4
	Fixed	-2.02 (5)	1999:1	-2.15 (5)	1999:1; 2007:4

$^{a, b}$ and ^{c} represents 1%, 5% and 10% level of significance, respectively. Number of lags included are shown in parenthesis.

1 break: Critical values at 1%, 5% and 10% significance level are -4.239, -3.566 and -3.211, respectively. Source: Lee and Strazicich (2004).

2 breaks: Critical values at 1%, 5% and 10% significance level are -4.545, -3.842 and -3.504, respectively. Source: Lee and Strazicich (2003).

Appendix 4.1: Nonlinear Models

A threshold model generally incorporate a variable that equals zero below a certain value and one otherwise. The threshold regression model allows for two or more branches governed by the values of the threshold variable in the estimation process. A two regime model can be expressed as:

$$y_t = X_t'\beta + Z_t'\gamma_1 + \varepsilon_t \qquad \text{if } h_t < c$$

$$y_t = X_t'\beta + Z_t'\gamma_2 + \varepsilon_t \qquad \text{if } h_t \geq c$$

where the vector X contains the variables whose coefficients do not vary across regimes and vector Z is associated with regime specific variables. h is the observable transition variable and c is the threshold value. Often, the mean of the series is used as c in the threshold autoregressive (TAR) models. Using an indicator function $1(h, c)$, the above expression can be written as:

$$y_t = X_t'\beta + 1_{\{h_t<c\}}Z_t'\gamma_1 + 1_{\{h_t\geq c\}}Z_t'\gamma_2 + \varepsilon_t$$

This expression can be used to perform a ADF-type unit root test by choosing the appropriate X, Z, h and c. Often, y_{t-d} is used as a proxy for h_t, although h_t can be any other exogenous variable.

In threshold models, a regime switches when the threshold variable crosses a certain threshold, so that there is a sharp cut-off between the branches. Although the model can capture many nonlinear features usually observed in economic and financial time series, the lack of continuity in the objective function causes other problems, e.g., asymptotic distribution theory cannot be used. An alternative is the Smooth Transition Regression (STR) model where the regime switches gradually in a smooth fashion. Smooth Transition Autoregression (STAR) models are widely used in the empirical literature. Consider the following STAR model of order p :

$$y_t = \beta_0 + \beta' x_t + (\theta_0 + \theta' x_t)F(h_t, \gamma, c) + u_t$$

where y_t is a scalar, $x_t = (y_{t-1}, \dots, y_{t-p})'$, $\beta' = (\beta_1, \dots, \beta_p)$, $\theta' = (\theta_1, \dots, \theta_p)$, $\gamma > 0$ and $u_t \sim NID(0, \sigma_u^2)$. Again, y_{t-d} is often used as a proxy for h_t.

In the STAR model, adjustment takes place in every period but the speed of adjustment varies with the extent of the deviation from parity. Two simple transition functions, suggested by Granger and Teräsvirta (1993) and Teräsvirta (1994), are the exponential STAR (ESTAR) and the logistic STAR (LSTAR). The transition functions $F(y_{t-d}, \gamma, c)$ in ESTAR and LSTAR is defined as $F(y_{t-d}, \gamma, c) = 1 - \exp\{-\gamma(y_{t-d} - c)^2\}$ and $F(y_{t-d}, \gamma, c) = [1 + \exp\{-\gamma(y_{t-d} - c)\}]^{-1} - 0.5$, respectively. Hence, $0 < F(y_{t-d}, \gamma, c) < 1$ and ensures the smooth transition between regimes. γ determines the speed of transition; a small value of γ implies a slow transition.

For an LSTAR model, switches between two regimes occur smoothly and depend on the distance between the transition variable y_{t-d} and the threshold value c; and the sign of the distance. On the other hand, in the ESTAR model, the transition function is symmetrical, so that switches between two regimes occur smoothly and depend only on the distance between y_{t-d} and c, but not on the sign of the distance. Thus, in the LSTAR model, for values significantly less than c, $F(y_{t-d}, \gamma, c)$ is near zero and the coefficients are β_0 and β. On the other hand, for values significantly greater than c, $F(y_{t-d}, \gamma, c)$ is near one and the coefficients are $\beta_0 + \theta_0$ and $\beta + \theta$. For the ESTAR model, values near c have coefficients near β_0 and β; while values farther away from c have coefficients near $\beta_0 + \theta_0$ and $\beta + \theta$. The ADF type unit root test with STAR model can be expressed as:

$$\Delta q_t = \alpha + \rho q_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta q_{t-j} + \left(\alpha^* + \rho^* q_{t-1} + \sum_{j=1}^{p} \gamma_j^* \Delta q_{t-j}\right) F\left(\Delta q_{t-j}, \gamma, c\right) + \varepsilon_t$$

In this setup, it is possible to have $\rho \geq 0$. But, for mean reversion, we need $\rho^* < 0$ and $\rho + \rho^* < 0$. This means that small deviations from PPP can be characterized by a unit root but large deviations are mean reverting.

Kapetanios, Shin and Snell (KSS) (2003) develop a test to examine global stationarity for a de-meaned or de-trended series. The test assumes that $\rho = 0$ in the previous equation and also assumes ESTAR transition function. Also, assuming $-2 < \rho^* < 0$, they test $H_0: \gamma = 0$ against $H_1: \gamma > 0$. Testing the null hypothesis is not feasible since ρ^* is not identified under the null. To overcome this problem and derive the test statistic, they compute a first-order Taylor series approximation under the null and get the following auxiliary regression:

$$\Delta q_t = \sum_{j=1}^{p} \tau_j \, \Delta q_{t-j} + \delta q_{t-1}^3 + \varepsilon_t$$

The test statistic is a t-statistic to test the $H_0: \delta = 0$ against $H_1: \delta < 0$. They also provide the asymptotic critical values of the test statistic.

In the Markov regime switching setup with two regimes, the parameters governing the ADF regression are changing with an unobserved state (regime) $s_t \in \{0,1\}$ such that,

$$\Delta q_t = \alpha_0(1 - s_t) + \alpha_1 s_t + [\beta_0(1 - s_t) + \beta_1 s_t]t + [\rho_0(1 - s_t) + \rho_1 s_t]q_{t-1}$$

$$+ \sum_{j=1}^{k}\big[\gamma_{0j}(1 - s_t) + \gamma_{1j} s_t\big]\Delta q_{t-j} + \varepsilon_t$$

where $\varepsilon_t \sim IIN(0, \sigma_\varepsilon^2)$ and can be regime-specific. The state variable is specified by a 2-state Markov chain with transition probabilities:

$$p(s_t = i | s_{t-1} = j, s_{t-2} = k, \dots, I_{t-1}) = p(s_t = i | s_{t-1} = j) = p_{ji} \geq 0$$

where $i, j = 0{,}1$, I_t is the information set up to time *t* and $\sum_{i=1}^{2} p_{ji} = 1$. This model endogenously allows the coefficients to switch as the regime changes. The unit root hypothesis can be tested for both regimes, i.e., $\rho_0 = 0$ and/or $\rho_1 = 0$. Rejection of any one hypothesis would indicate regime-specific stationarity. However, the distribution of the t-statistic under the null is nonstandard, so critical values have to be constructed by simulating the model under the null.

Chapter 5

Summary and Conclusions

This thesis is an empirical investigation into the relationship between the exchange rate and the domestic price level. The summary of our results and scope for future research are briefly discussed below.

Chapter 2 examines the causality between the nominal effective exchange rates (NEER) and the consumer price index (CPI); and estimates the response of CPI to exchange rate shocks for 12 OECD countries (Australia, Canada, Finland, France, Germany, Italy, Netherlands, New Zealand, Spain, Sweden, UK and USA) for the period 1974 to 2016. We found that exchange rate fluctuations affect the domestic price level by a smaller extent in the countries under the common currency Euro than in the period prior to their adoption of the Euro. In inflation targeting countries, the exchange rate tends to predict the domestic price level better than in the countries under the common currency and the exchange rate pass-through to CPI is statistically different from zero in four out of the six countries, at least at some horizons. In this chapter, we used a VAR model with three variables because of the short sample period. However, in future research, we intend to include additional variables in the VAR model and estimate the model using the Bayesian approach. This approach relies on informative priors to shrink the VAR model towards a parsimonious one; thereby reducing the parameter uncertainty associated with an over-parameterized model.

Chapter 3 examines whether the CPI responds asymmetrically to exchange rate appreciations and depreciations in 12 Asian countries (Bangladesh, China, India, Indonesia, Japan, Republic of Korea, Malaysia, Pakistan, the Philippines, Singapore, Taiwan and Thailand). This chapter contributes to the existing empirical evidence on this issue, which is currently limited to a few developed countries and to studies based on slope-based tests. Using a recently developed response-based test, we found evidence of asymmetric responses of CPI to NEER appreciations and depreciations in 6 (China, Indonesia, Korea, Malaysia, Pakistan and Philippines) out of the 12 countries. We found evidence that the results from response-based test can be different from the slope-based test. This implies that the statistically significant coefficient(s) of the variable(s) that capture asymmetry may not ensure asymmetry in responses (or vice versa). We also found that depreciations are not necessarily passed-through to prices more than appreciations. It also would be interesting to examine the regime-dependent response of CPI to exchange rate shocks. We intend pursue this research in the future by using threshold models or Markov regime switching models.

Chapter 4 examines whether the purchasing power parity (PPP) hypothesis holds for 12 Asian countries, located in East, Southeast and South Asia. Since stationarity of the real exchange rate implies that PPP holds, we employ the unit root tests on the real exchange rate to examine the PPP hypothesis. In the presence of structural break(s), the standard unit root tests (e.g., ADF test) are biased towards non-rejection of the unit root hypothesis. Further, since country-specific and global shocks (e.g., Asian financial crisis, global recession) can cause structural breaks in the real exchange rate series, it is preferable to use the unit root test that allow for multiple structural breaks. Using these tests and comparing our findings with those from the earlier studies in the literature, we found more support for the PPP hypothesis (in 6 out of the 12 countries). The relative lack of support for the PPP hypothesis in previous studies on our selected countries could be due to their shorter time span and/or inability to accommodate structural break(s) properly. Further, our study found little evidence of trend stationarity in these countries, so that there is no support for the HBS hypothesis.

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