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Author
Zeng, Li

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Essays on International Finance and Trade

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Li Zeng

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Committee in charge:

Professor Maurice Obstfeld
Professor Andrew K. Rose
Professor J. Brad DeLong
Professor Michael Jansson

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Abstract

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Professor Maurice Obstfeld, Chair

Chapter one of this dissertation tries to empirically answer the question of what explains China's foreign reserve holdings. The author finds that some VEC models motivated by modern mercantilism perform better in predicting China's foreign reserves than selected ARIMA models and those based on precautionary demand theories, both with statistical significance. This suggests that previous econometric studies, whose emphasis has been overwhelmingly placed on precautionary motives, might have missed some more important factors in explaining China's reserve holdings. Further structural analyses of these models show that the buildup of reserve stocks by China has a negative (depreciating) effect on its real exchange rate and a positive impact on its export growth. These findings seem to corroborate the conjecture that the recent increase in China's reserve holdings is a part of its export-led development strategy, and they also suggest that such a policy is effective for China.

The second chapter studies the impact of trade liberalization on imports, exports, and overall trade balance for a large sample of developing countries, using two recently constructed measures of trade liberalization dates. The authors find strong and consistent evidence that trade liberalization leads to higher imports and exports. However, in contrast to Santos-Paulino and Thirwall (2004), who find a robustly negative impact of trade liberalization on the overall trade balance, the authors find only mixed evidence of such a negative impact. In particular, the authors find little evidence of a statistically significant negative impact using their first measure of liberalization dates, which extends Li (2004). Using a second measure of liberalization dates compiled by Wacziarg and Welch (2003), the authors find some evidence that liberalization worsens the trade balance, but the evidence is not robust across different estimation specifications, and the estimated impact is smaller than that reported by Santos-Paulino and Thirwall (2004).

The last chapter empirically tests the proposition that domestic real oil prices can help to forecast real effective exchange rates. For most countries in the sample, little

evidence was found to support such a proposition. There were two exceptions, Japan and Norway, who showed some consistent signs in favor of the oil-price exchange rate model. But the statistics favoring the model are mostly insignificant. The author also finds that the relations between oil prices and real exchange rates are not stable over time. The conjecture that the oil-price exchange rate model can produce better forecasts on exchange rates when oil prices are more volatile is not consistent with the pattern seen in the subsample results.

CHAPTER ONE

What Explains China's Foreign Reserves
-- An Empirical Study from a Time Series Perspective

Li Zeng[1]

This paper tries to empirically answer the question of what explains China's foreign reserve holdings. We find that some VEC models motivated by modern mercantilism perform better in predicting China's foreign reserves than selected ARIMA models and those based on precautionary demand theories, both with statistical significance. This suggests that previous econometric studies, whose emphasis has been overwhelmingly placed on precautionary motives, might have missed some more important factors in explaining China's reserve holdings. Further structural analyses of these models show that the buildup of reserve stocks by China has a negative (depreciating) effect on its real exchange rate and a positive impact on its export growth. These findings seem to corroborate the conjecture that the recent increase in China's reserve holdings is a part of its export-led development strategy, and they also suggest that such a policy is effective for China. [JEL F31, F37]

1. Introduction

This paper will empirically study China's foreign reserve holdings, which have increased dramatically over the past two decades[2]. The first panel of Figure 1 shows that by the end of 2007 the stock of China's foreign reserves had reached over 1.5 trillion US dollars, nearly 100 times the 16 billion US dollars held in 1982. This amount also far exceeds the second largest reserve stock in the world, 0.97 trillion US dollars held by Japan. Such increase is quite remarkable even after the rapid growth of the Chinese economy is taken into consideration. The second panel of Figure 1 shows that the reserve to GDP ratio of China has risen from 5.5% to 47.6% over the period from 1982 to 2007, an almost nine-fold increase. This not only contrasts sharply to what happened to the advanced economies, whose average reserve to GDP ratio dropped from 6.4% in 1982 to 5.3% in 2006, but also dwarfs the increase in reserve holdings by the other emerging market economies, whose average reserve to GDP ratio has increased from 5.6% to 21% over the same period[3]. The

[1] I'm grateful to Professor Maurice Obstfeld for his great guidance. I also want to thank Professor Andrew Rose, Professor Pierre-Olivier Gourinchas, Professor Yuriy Gorodnichenko, Professor Yingyi Qian, Professor Brad DeLong, and Gewei Wang for helpful discussions. Email: lzeng@econ.berkeley.edu.

[2] Foreign reserves discussed in this paper are defined as the sum of gold, SDRs, foreign exchange reserves and reserve position with the IMF.

[3] Advanced economies are countries with IFS codes less than 199, with the exception of Turkey and South Africa. Emerging market economies are those in the Morgan Stanley emerging market index, as of April 25, 2008.

last panel of Figure 1 measures reserve holdings as months of national imports covered by them. In 1994, the reserve holdings of China, as well as the average of the other emerging market economies, covered approximately 6 months of imports. However, since then, the increase in China's reserve holdings has outpaced others and by the end of 2006 its months of imports covered by reserves is 16, much higher the average of 9 months for the other emerging market economies.

The enormous size of China's current foreign reserve stock has brought its cost efficiency into question. Bai, Hsieh and Qian (2006) find that the aggregate real rate of return to capital in China has averaged above 20% since 1978. This implies a potentially very high opportunity cost for China to hold foreign reserves, which are believed to be mostly in the form of low-yielding US Treasury and other securities. Added to the concerns of opportunity costs are the valuation effects emphasized by Gourinchas and Rey (2007). Unlike the case for the United States where depreciation of the US dollar has stabilizing effects through external adjustments, the valuation effects of dollar depreciation can lead, or probably have led, to huge losses to the value of the Chinese reserve stock and bring big balance sheet risks to the People's Bank of China[4]. There are also other costs and risks associated with holding a huge amount of foreign reserves. For instance, the continuing rise in the level of reserve stocks can push up the costs of sterilizing interventions and make it increasingly difficult to neutralize the inflationary monetary impact of reserve accumulation. The surge in the Chinese inflation rate in 2007 and early 2008 is an example of this possibility[5].

From a broader perspective, the large stock of China's foreign reserves is often perceived as a symptom of the current global imbalances. Such a view is well summarized by a quote from the Federal Reserve Board Chairman, Ben Bernanke, during a recent speech on the current financial turmoil:

The Sources of the Financial Turmoil: A Longer-Term Perspective

" ... a substantial increase in the net supply of saving in emerging market economies contributed to both the U.S. housing boom and the broader credit boom. The sources of this increase in net saving included rapid growth in high-saving East Asian countries and, outside of China, reduced investment rates in that region; large buildups in foreign exchange reserves in a number of emerging markets; and the enormous increases in the revenues received by exporters of oil and other commodities. The pressure of these net savings flows led to lower long-term real interest rates around the world, stimulated asset prices (including house prices), and pushed current accounts toward deficit in the

[4] A back-of-the-envelope calculation suggests that the loss to the value of China's reserve stock, caused by the dollar depreciation from 2005 to present, can easily exceed 3% of its annual GDP: $830 \times 0.6 \times (1 - 6.8/8.1)/2400 = 3.3\%$, where 830 billion dollars was China's reserve level at the end of 2005, 0.6 is the assumed proportion of dollar assets in the reserve stock, 8.1 was the RMB/dollar rate at the end of 2005 and 6.8 is the current RMB/dollar rate, and 2400 billion dollars was China's GDP in 2005.

[5] See Rodrik (2006) for a more general discussion on the cost of holding foreign reserves.

industrial countries - notably the United States - that received these flows."

Important and interesting questions arise here: Why is China holding such a huge amount of foreign reserves, if the associated costs are very high and they have caused unease among its major trade partners? What are the underlying forces that have driven the growth in China's foreign reserves? Is China's current reserve holding behavior different from what we observe in its past? This paper will try to answer these questions empirically, testing a set of alternative explanations.

Recent work on international reserves indicates two possible directions for explaining a country's reserve holdings. One direction of research can be summarized as the precautionary demand theories for foreign reserves. These theories suggest that countries use foreign reserves as a preventative measure against domestic and external shocks, and the reserve stocks are built up, by and large, on the countries' own initiatives. Studies along these lines include: Flood and Marion (2001), Aizenman and Marion (2003, 2004), Aizenman and Lee (2005), Jeanne and Ranciere (2006) and Jeanne (2007), who emphasize the self-insurance role of reserve stocks against external shocks such as sudden stops in capital inflows; McKinnon and Schnab (2003), McKinnon (2006), and Obstfeld, Shambaugh and Taylor (2008), who stress the importance of domestic financial stabilities in determining countries' reserve levels.

The other direction is what has been called modern mercantilism, which is a part of the broader Revived Bretton Woods System theory, first developed by Dooley, Folkerts-Landau and Garber (2003)[6]. The modern mercantilism theory suggests that export promotion through an undervalued currency is a long-term development strategy adopted by China currently, and it has been used by other countries during certain stages of their development, e.g. Japan. It views the vast amount of foreign reserves held by China as merely a by-product of its export-led growth policy.

In this paper we construct econometric models to explore China's foreign reserve holdings following both directions. Unlike most recent empirical studies, however, we will examine the questions from a time series perspective[7]. As we will argue below, the vector error correction (VEC) model adopted by this paper has a few important advantages over the cross-country panel method, used by most previous studies, in investigating China's reserve holding behavior.

The central finding in this paper is that models motivated by modern mercantilism perform better in predicting China's foreign reserve holdings than those based on precautionary demand theories, as well as selected ARIMA models, with statistical significance. This

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[6] Dooley, Folkerts-Landau and Garber (2004a, 2004b, 2004c, 2005a, 2005b and 2007) is a series of working papers on the Revived Bretton Woods System theory. The term *modern mercantilism* is borrowed from Aizenman and Lee (2005).

[7] A few early studies on reserves also took time-series perspectives, including Edwards (1983, 1984), and Ford and Huang (1994).

suggests that previous econometric studies, whose emphasis has been overwhelmingly placed on precautionary motives, might have missed some more important factors in explaining China's foreign reserves. Further structural analyses of these models show that the buildup of reserve stocks by China has a negative (depreciating) effect on its real exchange rate and a positive impact on its export growth, which corroborates the conjecture that the recent increase in China's reserve holdings is a part of its export-led development strategy. When interpreted in the broader context of current global imbalances, such results seem to support the view that China should play a more active role in the adjustment process by reevaluating its currency and/or adopting a more flexible exchange rate regime.

Among considerations related to precautionary motives, imports and associated risks, such as real openness of the economy, appear to be the most important determinants for China's reserves holdings. Inflation volatility and the interest rate differential between China and the US are also helpful predictors. But our results suggest that if China wants to reduce its inflation volatility by piling up reserves, that has not been an effective policy tool. The estimated impact of the interest rate differential on China's reserve holdings is opposite to that predicted by the buffer stock model. One possible explanation for such a finding is that higher relative returns to capital in China are likely to attract more capital inflows to China, outweighing the opportunity cost consideration suggested by the buffer stock model. We find little evidence to support volatility of exports and domestic financial depth as crucial factors in explaining China's reserve holdings.

A natural question to ask is why these results obtained from the time-series VEC model are more credible than the previous ones produced by cross-country panel estimations. We answer this by carefully considering some serious limitations of using panel method to analyze China's reserve holdings and showing how the VEC model can help to overcome them.

First, cross-country panel regressions assume that the coefficients of interest are the same, or at least close to each other, across all the countries in the sample. However, on the issue of foreign reserves, there are compelling reasons to believe that China may act quite differently from other countries, especially the emerging market economies, the country group that it presumably best fits in. For instance, from a historical perspective, state economy has played a much more important role in China than in most other emerging market countries. For most of the period that we study, none of the firms and households in China were allowed to hold foreign exchanges[8]. This means that the Chinese reserve stocks that we are looking at are not only the official foreign reserves, but indeed the foreign reserves held by China. Since there is little hidden cushion in the private sectors and households that can help to protect China from external shocks, it should not be surprising if China has a more prudent attitude toward reserve accumulations.

Another problem of using panel methods to investigate foreign reserves is the danger of spurious regression, which seems to be largely ignored by the current empirical literature.

[8] With the exception of remittances from abroad.

Foreign reserve levels, as well as some explanatory variables that have been experimented with by economists, often appear to be nonstationary time series. In such cases cointegration tests should be performed before one can claim that a stable relation among these variables has been identified. In practice, however, few studies employing panel methods have actually done so. Two assumptions are therefore implicitly assumed by those studies: the variables included in the panel regressions are cointegrated, and furthermore, the cointegrating relations among these variables are identical across all the sample countries. The example shown in Figure 2 and Table 1 casts doubts on both of them. Figure 2 shows the time series of reserves and imports, both as ratios to GDP, for China and Korea, who almost always show up on the same panel in studies on foreign reserves. Although the two reserve series display a similar upward trend over time, it is obvious that the paths of imports are very different. Indeed, the Johansen cointegration test results reported in the first panel of Table 1 suggest that the reserve and import series for China are cointegrated, but they are not in the case of Korea. Nonetheless, in the second panel of the table, where we report regressions of reserves on imports, we find significant coefficients not only for China, but also for Korea and for the case where both countries are included. This illustration of spurious regression not only confirms the heterogeneity concern we pointed out earlier, but also shows that without careful consideration of the nonstationarity issue, conclusions based on panel regressions can hardly be convincing.

The endogeneity problem is widely acknowledged in the empirical literature on foreign reserves. For instance, theoretically speaking a country's need for international reserves depends on its exchange rate regime. But in the reverse direction, the abundance of a country's international reserves can affect its exchange rate arrangement as well. As we saw in the 1997 Asian Financial Crisis, many countries were forced off their *de facto* pegs to the US dollar when their international reserves were nearly depleted. However, few existing studies have provided satisfactory solutions to the problem. Part of this is probably due to the difficulty in finding good instrumental variables.

Finally, many existing empirical studies are based on theoretical models predicting the optimal reserve levels for an economy. Since panel regressions do not distinguish between the long-term equilibria and short-run fluctuations *per se*, empirical studies employing such methods are therefore hampered by the fact that the researchers do not observe the optimal levels of reserve stocks, only the actual holdings instead. As a consequence, measurement errors in the variable to be explained, the optimal reserve holdings, may interact with the constructed regressors to generate a misleading correlation between reserve holdings and its potential determinants.

The VEC model can handle all the empirical difficulties mentioned above better. First, since the VEC model studies foreign reserves from a time-series perspective, cross-country heterogeneity will not cause any concerns in the estimations. Second, all the variables involved in a VEC model are treated as endogenous, allowing us to study the dynamics among them. In addition, cointegration tests are performed when the model is estimated through the Johansen maximum likelihood method, which ensures that we are not running spurious regressions. Finally, the estimation results of the VEC model separate the long-run equilibrium relations from the short-run adjustment dynamics, which, at least to some extent, solves the measurement error problem on optimal reserve levels.

The rest of the paper is organized as the following. Section 2 introduces the VEC model, discusses the regressors that will be included in estimations, and sketches the scheme of our empirical investigation. Section 3 provides more details on the estimations and discusses all the empirical results. Section 4 concludes the paper with a summary of the main findings and a brief discussion on future work.

2. Model, Regressors, and Investigation Scheme

VEC Model

The workhorse tool in this paper is the vector error correction (VEC) model. The reduced form VEC model that we will estimate is the following:

$$\Delta y_t = \alpha(\beta' y_{t-1} + \mu) + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + \gamma + \varepsilon_t, \qquad (1)$$

where $y_{(n\times 1)}$ is the vector of regressors, $\beta_{(n\times k)}$ with $k < n$ is the matrix of cointegrating vectors, $\alpha_{(n\times k)}$ and $\Gamma_{i\,(n\times n)}$ are matrices of adjustment parameters, $\mu_{(n\times 1)}$ is a constant vector that allows the cointegrating equations to have nonzero means, the constant vector $\gamma_{(n\times 1)}$ allows linear trends in the levels of the data, p is the number of lags included when the model is written in levels, and ε_t is the *i.i.d.* residual vector with assumed distribution $N(0,\Omega)$[9,10]. This equation can be interpreted as saying that the current period adjustments in *y* are affected by their own historical values, the deviations from their long-run relations in levels in the last period, and some random shocks.

Equation (1) is called the reduced form because the coefficients for Δy_t on the left hand side is restricted to be an identity matrix. The consequence is that the elements of ε_t are not necessarily orthogonal to each other (or equivalently, Ω is not a diagonal matrix), therefore we can not simply interpret the estimation results of equation (1) as causal

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[9] γ and $\alpha\mu$ are orthogonal to each other.

[10] Equation (1) nests two forms of the vector autoregression (VAR) model. If $k = n$, then $(\alpha\beta')$ is an unrestricted $n \times n$ matrix and equation (1) is in fact a VAR model in levels. The VAR model in levels fits the data better than the VEC model and therefore would be preferable if all the variables in *y* are stationary series. But since the variable of our primary interests, China's foreign reserves, is an $I(1)$ process, the VEC model is more appropriate and efficient. If $k = 0$, then equation (1) becomes a VAR model in first order difference. Such a model is appropriate if all the variables in *y* are nonstationary $I(1)$ processes and no cointegrating relation exists among them. We will, however, just ignore these situations, because they imply that no other variables in the model have a stable long-run relation in levels with the variable that we are trying to explain.

relations among the regressors. To determine causation, we will continue to estimate the structural coefficient matrix, A, in the structural form of the VEC model, which is equation (2) below:

$$A\Delta y_t = \overline{\alpha}(\beta' y_{t-1} + \mu) + \sum_{i=1}^{p-1} \overline{\Gamma}_i \Delta y_{t-i} + \overline{\gamma} + v_t , \tag{2}$$

where $\overline{\alpha} = A \times \alpha$, $\overline{\Gamma}_i = A \times \Gamma_i$, $\overline{\gamma} = A \times \gamma$, and most importantly, $v_t = A \times \varepsilon_t$ and it has a diagonal variance and covariance matrix, Σ. Since there are no contemporaneous correlations among the elements of v_t, they are regarded as structural (exogenous) shocks to the corresponding regressors in y.

Tests of cointegrating relations among regressors and estimations of equation (1) will both be executed using Johansen's maximum likelihood method. Identifications of the coefficients in matrix A will be based on the second moment conditions derived from the equation $v_t = A \times \varepsilon_t$:

$$\Sigma = A\Omega A' . \tag{3}$$

Since the number of elements in A exceeds the number of moment conditions implied by equation (3), identification restrictions will have to be imposed. In this paper we will follow the just identification strategy proposed by Bernanke (1986). Details of the restrictions imposed will be discussed in the related parts of section 3.

Regressors

The variable of our primary interests in this paper is China's foreign reserve holdings scaled by GDP (*res*). The other regressors included in the VEC models are motivated by previous theoretical and empirical work on foreign reserves. The construction details and summary statistics for all the variables are provided in the data appendix. The reasons for the inclusion of each regressor are briefly discussed below.

We start with variables related to precautionary motives of holding reserves. Import (*imp*) is the most robust regressor for reserves found by the empirical literature. Reserves are the "financing option of last resort" in covering a country's import demand, providing a natural link between these two variables. In a broader sense, imports can also be interpreted as a measure for real openness and therefore the vulnerability to external shocks of an economy. The buffer stock model introduced by Frenkel and Jovanovic (1981) suggests that the optimal level of reserve holdings by a country is affected by the opportunity cost of holding them. This is proxied by the interest rate differential between China and the US (*idif*) in the paper. Another factor that we consider is external debt (*debt*). The importance of external debts in determining a country's reserve levels has regained much attention since the 1997

Asian Financial Crisis, which can be seen from the now-famous Guidotti-Greenspan rule[11]. Due to data availability, we will only study the impact of total external debts in this paper. We also check whether volatility of exports (*vexp*) can help to explain China's reserve holdings. It can be regarded as a measure for volatility in international transactions. If a country holds reserves to smooth its international transactions, then higher volatility in exports would justify the holding of more reserves. Two variables associated with domestic financial stabilities, volatility of inflation (*vcpid*) and M2 to GDP ratio (*m2gdp*), are also examined. The inclusion of inflation volatility is based on the view, shared by McKinnon and Schnab (2003) and McKinnon (2006), that China should peg its currency to the US dollar so as to provide a nominal anchor to its domestic price level and for most of the time reserves are helpful to maintaining such a peg. M2 to GDP ratio is a predictor for reserve holdings proposed by Obstfeld, Shambaugh and Taylor (2008), who emphasize the cushion effect of reserves against potential capital flight caused by domestic financial instabilities.

The modern mercantilism theory provides an alternative thought on explaining China's foreign reserves, claiming that China has been piling up reserves to keep an undervalued currency and promote exports. We take two different approaches to assess this story. One is to first estimate the deviation of the Chinese exchange rate from some benchmark value and then include such deviation in the VEC model to see whether it helps to explain China's foreign reserves. The deviation measure (*penn*) used in this paper is based on the empirical regularity called the Balassa-Samuelson relation, which predicts an association between higher (appreciated) real exchange rate and higher level of real per capita income[12]. The other way is to directly include exchange rate (*exrt*) in the VEC model and inspect how it interacts with reserves over time. When we proceed in this second way, we will also control for some other factors that may affect the exchange rate, including the real GDP growth rate of China relative to the rest of the world (*rgg*) and China's external debt level (*debt*). One last regressor studied in the paper is the real export growth rate (*expg*). It is at the heart of the modern mercantilism theory and affects both reserves and exchange rates. We want to know whether buildup of reserve stocks has helped China to promote exports.

Investigation Scheme

The VEC model usually performs better with parsimonious specifications. It would be inefficient, and practically infeasible, to include all the regressors that we study in a single VEC model. Therefore an investigation scheme is designed, to help us identify the best predictors for China's reserve holdings while keeping the estimation results tractable.

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[11] The emphasis on the role of external debts in determining a country's reserve levels can in fact be traced back to at least a century ago. Please see Obstfeld, Shambaugh and Taylor (2008) for a quote from *Treatise on Money* (1930) by John Maynard Keynes.

[12] One may want to turn to more sophisticated "fundamentals-based" models to search for the equilibrium values of exchange rates. However, as Dunaway, Leigh and Li (2006) showed specifically for China, small changes in model specifications, explanatory variable definitions, and time periods used in estimation can lead to very substantial differences in equilibrium real exchange rate estimates.

We divide the regressors other than *res* into three groups. The first one consists of the "usual suspects" that have been tested extensively in the empirical literature on reserves, including *imp*, *idif*, *debt* and *vexp*. There are two goals that we want to realize through the study of this group of regressors. One is to check how well these variables can explain China's reserve holdings and the other is to search for some benchmark specifications, to which we can augment other regressors later on. We need such benchmark specifications because, as we will see in section 3, some regressors do not appear to be cointegrated with *res* by themselves. However, this does not necessarily imply that they cannot contribute to the explanations of reserves at all. We will do further examinations by adding them to the benchmark specifications before reaching the conclusions. The second group includes the two regressors associated with domestic financial stabilities, *vcpid* and *m2gdp*, and variables motivated by the modern mercantilism theory are in the third group. For these two groups, we will first check how well they can explain reserves by themselves. If the results are not satisfactory, we will then add them to the benchmark specifications and see whether they can bring some improvements.

Time-series models such as VEC and VAR are best known for their predictive capability. This is the key for our cross-specification comparisons. We will compare predictions by different models using the Diebold-Mariano test, taking both in-sample and out-of-sample forecasts into consideration. Furthermore, for models with good predictive power, we will check and see whether their implications are consistent with theories and previous empirical findings that motivate them. One possible way to do this is to look at the long-run relations among the regressors implied by the models, but a more informative method is to study the impulse response functions (IRF). When causal relations among the regressors are of particular interests, we will continue and estimate the structural form of the VEC model, and then implement post-estimation analysis with tools such as structural impulse response functions and forecast error variance decompositions (FEVD)[13].

All planned investigations will be carried out for both the full sample (1983Q1-2005Q4) and a subsample (1996Q1-2005Q4).

3. Estimations and Empirical Findings

Data available for estimations in this study cover 1983Q1-2007Q4. The last two years (2006 and 2007) are saved for out-of-sample predictions, therefore the full sample used to estimate the VEC models ranges from 1983Q1 to 2005Q4. The Chinese economy has been

[13] There are several reasons why we want to study the nonstructural IRF, even for those models whose structural forms are estimated. First, for prediction purposes, we should look at the nonstructural IRF. Second, since previous panel regressions had no satisfactory solutions to the endogeneity problem, it is possible that they were merely estimating the correlations between reserves and the explanatory variables. If it is indeed so, we should look at the nonstructural IRF to see whether our results match previous findings. Finally, it is interesting to compare the structural and nonstructural results. The difference between them shows how crucial it is to take the endogeneity among the variables into consideration in order to find out the true causal relations among them.

in fast transitions since the start of its economic reform and opening in the late 1970s. Particularly, in the mid 1990s there were several important policy shifts regarding to China's exchange rate regime and foreign exchange management. To check whether our results are sensitive to the time frame, we will also estimate the models using the subsample from 1996Q1 to 2005Q4.

3.1 Full Sample (1983Q1-2005Q4)

Following the investigation scheme discussed in section 2, estimations using the full sample are summarized in Tables 2.1-3. The Diebold-Mariano test results for prediction comparisons are reported in Table 3[14].

Precautionary demand (1)

The VEC models in Table 2.1 involve four regressors other than *res*. They are *imp*, *idif*, *debt* and *vexp*. An exhaustive search strategy is taken here, that is, models with all possible combinations of the four variables are estimated.

The first result to notice is that *imp* is the only variable that is cointegrated with *res* by itself, which is why the numbering of the models is inconsecutive in the first column of the table[15]. As we will see later, such a result stays true even if we take into consideration the other two variables motivated by precautionary demand, *vcpid* and *m2gdp*. It suggests that if there is only one long-term determinant for China's foreign reserve holdings among these variables, it is most likely to be *imp*.

A more surprising result is that the simple VEC specification with only *res* and *imp* (F1_01) produces the best predictions for *res* among all the models in Table 2.1. From the table we can see that this specification has the smallest root mean squared error (RMSE) and root mean squared forecast error (RMSFE) for the *res* equation[16]. We can also see this result from the first part of Table 3, where model F1_01 is used as the benchmark for the Diebold-Mariano test. The test results show that the predictions for *res* by all the other specifications in Table 2.1 are worse than those by model F1_01. Although the differences are not statistically significant, such results at least show that the three variables, *idif*, *debt* and *vexp*, are not very essential to the explanation of China's reserve holdings. As mentioned earlier, one purpose of studying this group of regressors is to look for benchmark specifications for later estimations. An ideal benchmark would be a simple model that produces a good fit for the variable to be explained. Model F1_01 meets both conditions. In addition, we will also use the model with *res*, *imp* and *idif* (F1_05) as a

[14] Predictions by different models were also compared using the Davidson-MacKinnon test. Since the results are largely consistent with those given by the Diebold-Mariano test, they are not reported.

[15] Here is an example of how we name the models. "Model F1_05" means that is the fifth model that we estimate for the first group of regressors using the full sample.

[16] For VEC models, RMSE is a measure for average in-sample prediction error. RMSFE is calculated based on one-period-ahead out-of-sample forecast errors.

benchmark, which is a clear "second best" in terms of predictions among the models in Table 2.1.

One way to find out whether the estimation results are consistent with theories and previous empirical findings is to check the long-run relations among the regressors implied by the *res* equations. For instance, we can see from Table 2.1 that the long-run coefficients for *imp* in the *res* equations are all positive. This implies that China's reserve holdings will increase when its import to GDP ratio rises, which is in line with the theory that motivates *imp* as a regressor for *res*. It is worth noticing, however, that there are a few flaws in this method of interpreting a VEC model. First, the long-run interactions among the regressors shown by the VEC system may be different from those implied by a single equation[17]. Second, the long-run relations do not reveal the adjustment process through which the system returns to equilibrium after receiving a shock. Finally, as common to most time series models, individual coefficients for a VEC model are often statistically insignificant, making the interpretation on each one of them less meaningful. For these reasons, another tool for analyzing VEC estimation results, the impulse response functions (IRF), is more widely used.

The IRF of the two selected benchmark models are plotted in Figures 3.1 and 3.2, respectively. Both figures show that when *imp* receives a positive shock, *res* will first decrease, but then soon increase over time. It will rise for about 1.6 to 1.7 percentage point in 50 periods if the shock in *imp* is of unit size. On the other hand, *imp* will also increase following a positive shock in *res*, but the response is less significant and has a much smaller size, only about 0.14 percentage point after 50 periods if the shock to *res* is of unit size. Although the result that *res* will respond positively to shocks in *imp* coincide with findings by previous studies, the adjustment processes depicted in these figures have different implications for the appropriate econometric tools that should be applied to analyzing reserve holdings. The IRF graphs show that it takes *res* very long to fully respond to the shock in *imp*, which implies that for most of the time the actual reserve levels might be different from the equilibrium ones. This articulates the measurement error problem associated with panel methods mentioned in the introduction.

Figure 3.2 shows that *res* will increase following a positive shock in *idif*. This is opposite to the prediction of the buffer stock model, which treats *idif* as a measure for the opportunity cost of holding reserves. One possible explanation for such a finding is that higher relative returns to capital in China, reflected by higher interest rate differentials between China and the US, are likely to attract more capital inflows to China, and therefore push up its foreign reserve levels[18].

[17] This complication can be caused by the multiple cointegrating relations among the regressors.

[18] We also estimate the structural forms of models F1_01 and F1_05, using the Cholesky decomposition and assuming *res* to be the "most endogenous" variable. For these models, the structural dynamics are qualitatively the same as what we see in the nonstructural results. They are not reported because later we will see similar analysis for more complicated models with better fits.

Another question raised earlier is whether China's current reserve holding behavior is different from what we observed in the past. Some clues to answering this can be found in the prediction errors. In the last part of Table 2.1 we see that for all the models, the mean prediction errors on *res* are slightly negative for the subsample period from 1983 to 1995, slightly positive for 1996-2005, and positive but of much bigger sizes for the out-of-sample period 2006-2007. This pattern seems to suggest that there is a systematic bias among the out-of-sample predictions made by the VEC models in Table 2.1 and China is holding more and more reserves over time. However, further examinations show that such statements are not quite accurate. Plotted in Figure 4 are the prediction errors on *res* by the benchmark model F1_01. Four big outliers can be seen in the graph, two in the early period, one for 1992Q3 and the other 1994Q1, two toward the end of the time line, one in 2004Q4 and the other 2007Q1[19]. When these outliers are excluded, there is no obvious shift of patterns among the prediction errors. In the last few columns of Table 2.1 we see that after excluding 2004Q4, model F1_01's mean prediction error on *res* for 1996-2005 is almost zero, and the mean for 2006-2007 without 2007Q1 actually turns slightly negative. Similar results are found for most other models as well, including those we will see later in Tables 2.2 and 2.3. This suggests that other than the two big positive shocks received in 2004Q4 and 2007Q1, China's reserve holding behavior in recent years is not very much different from before.

Precautionary demand (2)

We next study the two variables related to China's domestic financial stability, *vcpid* and *m2gdp*. Since no cointegrating relations are identified between these two variables and *res*, whether individually or combined together, we proceed by adding them to the two benchmark specifications and checking whether any improvement can be achieved. The estimation results are summarized in Table 2.2[20].

When *vcpid* is added to the benchmark specifications, the in-sample predictions for *res* by both models get better. The RMSEs of the *res* equations in models F2_04 (*res*, *imp*, *vcpid*) and F2_07 (*res*, *imp*, *idif*, *vcpid*) are 0.743 and 0.744, respectively, both smaller than the 0.768 of model F1_01 and 0.776 of model F1_05. This comes with the cost of slightly worse out-of-sample forecasts. The RMSFE for *res* of model F2_04 is 1.52, very close to the 1.517 of model F1_01, and model F2_07's 1.551 is slightly larger than the 1.535 of model F1_05. Taking both the in-sample and out-of-sample predictions into consideration, the Diebold-Mariano test results in Table 3 show that both new models, F2_04 and F2_07, predict *res* better than the more fitting benchmark F1_01 does, and the improvement in model F2_07 is significant. Indeed F2_07 is the only model in Tables 2.1 and 2.2 whose predictions on *res* beat model F1_01 significantly. These results suggest that *vcpid* is a

[19] The causes for the early outliers are relatively clear. The negative spike in 1992Q3 is most likely the result of China's large trade deficit in that period. The positive shock in 1994Q1 can be explained by the upsoaring of China's trade surplus following the RMB devaluation from 5.7 to 8.7 yuan/dollar.

[20] No cointegrating relation is found when the model includes *imp*, *idif*, *vcpid* and *m2gdp*, which would otherwise be model F2_09 in the table.

useful predictor for China's reserve holdings.

The addition of *vcpid* does not change the long-run coefficients of *imp* in the *res* equations by much, which still imply that *res* will rise following a positive shock in *imp*. The long-run coefficient of *idif* for *res* stays positive in model F2_07 and has become more significant. The long-run coefficients of *vcpid* in models F2_04 and F2_07 are both positive, though insignificant. They imply that when inflation becomes more volatile, China tends to hold more foreign reserves. These inferences based on the long-run coefficients have supports from the IRF of the two models, which are plotted in Figures 3.3 and 3.4, respectively. Both figures show that *res* will respond positively to a shock in *vcpid*, and the dynamics among *res*, *imp* and *idif* are very similar to those seen in Figures 3.1 and 3.2. If maintaining a nominal peg to the US dollar so as to stabilize the domestic price is part of the reason why China builds up its reserve stock, a conjecture that seems backed up the finding that *res* will rise following a positive shock in *vcpid*, we would wonder how effective such a policy is. From the IRF in Figures 3.3 and 3.4, it seems that the sizes of *vcpid*'s responses to shocks in *res* are extremely small.

Yet we have to caution ourselves before drawing any conclusion, because the analyses based on Figures 3.3 and 3.4 are nonstructural. To figure out the exact causal relations among the regressors, we need to obtain the structural IRF by estimating the coefficient matrix *A* in equation (2). We only report the structural analysis for model F2_07 here. With similar identification assumptions, what we learn from model F2_04 are qualitatively the same as the findings for model F2_07.

The identification restrictions we impose to estimate the structural form of model F2_07 can be described by the following matrix equation:

$$A\varepsilon = \begin{pmatrix} 1 & a_1 & a_2 & a_3 \\ a_4 & 1 & 0 & 0 \\ a_5 & 0 & 1 & 0 \\ a_6 & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \varepsilon_{res} \\ \varepsilon_{imp} \\ \varepsilon_{idif} \\ \varepsilon_{vcpid} \end{pmatrix} = \begin{pmatrix} v_{res} \\ v_{imp} \\ v_{idif} \\ v_{vcpid} \end{pmatrix} = v. \tag{4}$$

With four variables in the model, the moment conditions implied by equation (3) allow ten coefficients to be estimated freely. Six of them are parameters a_1-a_6 shown in equation (4), and the other four will be the standard deviations for the elements of vector v. The first line of equation (4) is saying that in any period *res* is affected not only by its own structural shock, but also directly by contemporaneous structural shocks to all the other variables in the system. The rest of equation (4) says that structural shocks in *res* have direct impacts on the other variables as well, but variables other than *res* do not have direct interactions among them within the period when the structural shocks occur. Notice that this is weaker than assuming *imp*, *idif* and *vcpid* are exogenous to each other, because, for instance, v_{imp} can still affect ε_{vcpid} through its impact on ε_{res}. In fact the structure laid out in equation (4) implies that no variable in the system is exogenous.

The structural IRF for model F2_07 based on the estimation of equation (4) are plotted in Figure 5.1.1[21]. The graphs in the third column are the responses of *res* to different structural shocks. They show that exogenous shocks to *idif*, *imp* and *vcpid* will all cause *res* to increase. Checking the impact of *res* on *vcpid* (the third graph in the last column), we still see that the responses of *vcpid* to shocks in *res* are very close to zero, with a perceivable size only in the period when the shock occurs. The ineffectiveness for *res* to affect *vcpid* can also be seen from the forecast error variance decompositions (FEVD) for model F2_07, which are plotted in Figure 5.1.2. The third graph in the last column shows that the proportion of the forecast error variance in *vcpid* that can be attributed to shocks in *res* is near zero. These results imply that China tends to hold more reserves when its domestic inflation gets more volatile, but if it wants to reduce the inflation volatility by piling up reserves, that has not been an effective policy tool[22].

In contrast to *vcpid*, the models with *m*2*gdp* added, including F2_05 (*res*, *imp*, *m*2*gdp*), F2_06 (*res*, *imp*, *vcpid*, *m*2*gdp*) and F 08 (*res*, *imp*, *idif*, *m*2*gdp*), all have substantial deteriorations in the overall fits of the *res* equation and therefore higher RMSEs, relative to the benchmark specifications. The test results in Table 3 show that their predictions on *res* are significantly worse than those of model F1_01. Such findings lend little support to the hypothesis that increased financial depth is the main reason why China is holding increasingly more reserves. Two of the models, F2_05 and F2_06, have better out-of-sample forecasts than the benchmark model F1_01. This could be a sign of the decreased sterilization ability of the People's Bank of China, which leads to a stronger correlation between China's reserve holdings and its money stocks.

Modern mercantilism

The models estimated so far are based upon economic theories associated with various precautionary motives. But do these theories really give us any edge over "dumb" models such as ARIMA? The comparison results between our theory-based models and a selected ARIMA model are reported in the last part of Table 3. They show that among the models estimated so far, the two benchmark specifications we chose and the two models with *vcpid* added to them do perform better in predicting *res* than the ARIMA(4,1,2) model does. However, the gains are not statistically significant. In this part we will estimate models motivated by the modern mercantilism theory. As a preview of the results, some of the models beat not only the ARIMA model, but also the benchmark specifications, both with statistical significance.

We attempt two different approaches to construct the VEC models following the thought of modern mercantilism. One is to first estimate the deviation of China's exchange rate from

[21] Estimation results of equation (4) and structural forms of other models are provided in Table A4.

[22] Compared to these structural results, the nonstructural dynamics for model F2_07 do not seem to be very misleading. However, as we will see later, not taking account of the endogeneity structure among the regressors will sometimes lead us to different or even wrong conclusions.

some benchmark value and then use the deviation as a regressor to explain China's reserves[23]. The other is to directly include exchange rate as a regressor in the VEC model. Results using these approaches are reported respectively in the two panels of Table 2.3. In the first panel of the table, although the deviation measure *penn* and its combination with *expg* both appear to be cointegrated with *res* (models F3_01 and F3_02), the fits of the *res* equations in these two models are much lower than the benchmark specifications. The Diebold-Mariano test results in Table 3 suggest that their predictions on *res* are significantly worse than those of model F1_01. The test between models F3_03 (*res*, *imp*, *penn*) and F1_01 favors model F3_03, but the very high p-value (0.947) implies that the difference between their predictions is just trivial. The prediction performances of models F3_04 (*res*, *imp*, *penn*, *expg*) and F3_05 (*res*, *imp*, *idif*, *penn*) are both worse than model F1_01.

The only specification that looks promising in the first panel of Table 2.3 is model F3_06 (*res*, *imp*, *idif*, *penn*, *expg*). Although its RMSE and RMSFE for *res* are both larger than those of the benchmark models, the overall fit of its *res* equation is so high that the Diebold-Mariano test suggests that its predictions on *res* are significantly better than those of model F1_01. Since this model is motivated by modern mercantilism, we have to check whether its implications are consistent with the theory, before we can claim a success. Unfortunately the model fails the check. A key piece of modern mercantilism is the negative (depreciating) effect of reserve piling on the exchange rate. But the IRF graph of *penn* to *res* in Figure 3.5 (the last picture in the second to last column) suggests the opposite. We see that *penn* will rise following a positive shock in *res*, which implies a either less undervalued or more overvalued Chinese currency.

Are these positive responses of *penn* to *res* merely a correlation or they actually reflect the structural relation between them? We estimate the structural form of model F3_06 with restrictions in equation (5)[24]:

$$A\varepsilon = \begin{pmatrix} 1 & a_1 & a_2 & a_3 & a_4 \\ a_5 & 1 & 0 & a_6 & a_7 \\ a_8 & 0 & 1 & 0 & 0 \\ a_9 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & a_{10} & 1 \end{pmatrix} \begin{pmatrix} \varepsilon_{res} \\ \varepsilon_{imp} \\ \varepsilon_{idif} \\ \varepsilon_{penn} \\ \varepsilon_{\exp g} \end{pmatrix} = \begin{pmatrix} v_{res} \\ v_{imp} \\ v_{idif} \\ v_{penn} \\ v_{\exp g} \end{pmatrix} = v\,, \tag{5}$$

where *res* is again assumed to be directly affected by all the structural shocks, and it has direct influences on all the other variables as well, except *expg*. The only variable that affects *expg* directly is *penn*. Also assumed in equation (5) is that *imp* will be directly affected by the deviation of exchange rate from its benchmark and the growth rate of real

[23] Please refer to section 2 and the data appendix for details on the deviation measure used here.

[24] We tried alternative sets of restrictions with slight modifications to equation (5). The conclusion remained the same.

exports[25]. The structural IRF in Figure 5.2.1 show that although a positive shock in *res* causes *penn* to drop (the Chinese currency depreciates) immediately, such a negative effect lasts only very briefly, for about 4 periods, and then turns positive with even larger magnitude. More importantly, the stimulus that a *res* shock gives to *expg* is also temporary (the last graph in the first column), whose size is much smaller than the subsequent decrease of *expg*. For most of the time the effect of *res* on *expg* stays negative. These can hardly be interpreted as strong evidence supporting the modern mercantilism theory.

However, the fail of models F3_01-06 is not the end of the road. Rather, a careful look at the deviation variable *penn* suggests that the problem probably lies in the imperfect benchmark measure for the exchange rate. In the data appendix we can see that the correlation between *penn* and *exrt* is almost one, and the reason is that the benchmark exchange rates based on the Balassa-Samuelson relation are not nearly as volatile as the actual exchange rate series[26]. Since *penn* is essentially the same series as *exrt*, it leads us to our second approach of constructing the VEC model, that is, to directly include *exrt* as a regressor.

Thus far we have centered the VEC models around *res*, because it is this variable that we want to explain. But if the immediate target of China's reserve policy is the exchange rate, as suggested by the modern mercantilism theory, logically we should also control other factors that influence the exchange rate in the VEC model. This is the thinking that underlies models F3_07-10, which are reported in the second panel of Table 2.3.

The main control variable added to these models is *rgg*, China's real GDP growth rate relative to the rest of the world. We use it as a proxy for relative productivity, which has been proposed by the literature as a determinant for real exchange rate. If reserves affect exchange rate as a demand factor for foreign exchanges, then external debt can be thought of as a component on the supply side. It is controlled in two of the models and turns out be a helpful regressor.

We now check the details of these models and first verify that both *expg* and *debt* are helpful to the explanation of China's reserve holdings. Export growth rate, *expg*, is left out by models F3_07 (*res*, *rgg*, *exrt*) and F3_09 (*res*, *debt*, *exrt*). When it is added to model F3_07, the improvement seen in model F3_08 (*res*, *rgg*, *exrt*, *expg*) is very significant. The R^2 of the *res* equation increases from 0.29 to 0.52, and both the RMSE and RMSFE become smaller. The improvement of model F3_10 (*res*, *debt*, *exrt*, *expg*) over F3_09 is not as dramatic, but still obvious. The R^2 of the *res* equation increases from 0.53 to 0.58, although the RMSE gets higher, from 0.76 to 0.78, due to increased number of regressors. Model F3_10's RMSFE (1.373) is much smaller than that of model F3_09 (1.473), which implies better out-of-sample forecasts by model F3_10. The contribution of *debt* can be seen in a similar way, if we compare model F3_09 to model F3_07 and model F3_10 to

[25] We assume that *imp* is directly affected by *expg* because processing trade accounts for a large proportion in China's exports.

[26] It could be related to "the exchange rate disconnect puzzle", as termed by Obstfeld and Rogoff (2000).

model F3_08.

How well do these models perform relative to the other models we discussed earlier? Models F3_08-10 turn out to have the best predictive power among all the models. From Tables 2.1-3 we see that they have the highest overall fits for the *res* equation, much higher than most other models, although their RMSEs are not particularly small due to increased number of regressors. They also have the best out-of-sample forecasts, reflected by their small RMSFEs. The Diebold-Mariano test results in the last part of Table 3 show that they are the only models whose predictions beat the ARIMA(4,1,2) model significantly. They outperform the benchmark specification F1_01 as well, and the difference is significant for model F3_10. The middle part of Table 3 compares model F3_10 to all the other models, which shows that it has the best predictions for *res* and the gains over most models are significant[27].

One last task is to verify whether the implications of these models are consistent with the story of modern mercantilism. As we will see, the three models all seem supportive to the theory that motivates them.

We start with model F3_08. At a first look, its IRF graphed in Figure 3.6 do not appear to be exactly what is envisioned by the modern mercantilism theory. In particular, they show that following a positive shock to *res*, *exrt* will appreciate (the third graph in the second column), and *expg* will first have a temporary rise, but then the impact will fluctuate around zero until it dies out over time (the third graph in the first column).

For this model, however, the structural IRF in Figure 5.3.1 and FEVD in Figure 5.3.2 suggest a different story. Their underlying structural form is estimated by imposing the following restrictions[28]:

$$A\varepsilon = \begin{pmatrix} 1 & a_1 & a_2 & 0 \\ a_3 & 1 & a_4 & a_5 \\ 0 & a_6 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \varepsilon_{res} \\ \varepsilon_{exrt} \\ \varepsilon_{\exp g} \\ \varepsilon_{rgg} \end{pmatrix} = \begin{pmatrix} v_{res} \\ v_{exrt} \\ v_{\exp g} \\ v_{rgg} \end{pmatrix} = v\,. \qquad (6)$$

Equation (6) assumes *rgg* to be exogenous to contemporaneous structural shocks to the other variables, *exrt* to be directly affected by structural shocks to all the variables, *res* by all the variables but *rgg*, and *expg* only by itself and *exrt*. In Figure 5.3.1 we see that a positive shock in *res* leads *exrt* to depreciate (the third graph in the second column), and

27 Test results not reported in tables 3 show that models F3_08 and F3_09 also have better predictions for *res* than all the early models. But their dominances are not as significant as model F3_10.

28 We also tried other sets of restrictions with slight modifications to equation (6). The results are qualitatively the same.

this is not a simple correlation between *res* and *exrt* because the second graph in the third column shows that a positive shock in *exrt* will cause *res* to increase. In addition, the figure shows that *res* has a sizable positive impact on *expg* right after the shock, which becomes smaller over time but stays positive throughout. Furthermore, Figure 5.3.2 indicates that shocks to *res* account for a big proportion in the forecast error variances of both *exrt* and *expg*. These results all match the story told by the modern mercantilism theory, and the importance of taking account of the endogeneity structure is manifested by the difference between the structural and nonstructural dynamics.

Figures 3.7 and 3.8 are the IRF of models F3_09 and F 10, respectively. Since the dynamics in Figure 3.7 are similar to those seen in Figure 3.8, and because model F3_09 misses the variable *expg*, we will focus our discussions on model F3_10. Figure 3.8 shows that following a positive shock in *res*, *exrt* will first appreciate (the second to last graph in the third column). Such a positive effect lasts for about 15 periods and then turns negative, with an increased size over time until becoming constant after about 35 periods of the initial shock. In the mean time, except for a very short period, the influence of *res* on *expg* stays positive and remains sizable even after 50 periods of the shock. These findings seem to support the modern mercantilism theory, because we see both of its key components, the negative effect of *res* on *exrt* and its positive effect on *expg*. But their supports are much weakened by the result that the exchange rate will first appreciate for a substantial period of time following a positive shock in reserves.

Nonetheless, further examinations show that such appreciation is not the structural response of the exchange rate to an exogenous shock in reserves. The estimation of the structural form of model F3_10 is based on the following identification assumptions:

$$A\varepsilon = \begin{pmatrix} 1 & a_1 & a_2 & a_3 & 0 \\ a_4 & 1 & a_5 & a_6 & a_7 \\ 0 & a_8 & 1 & 0 & 0 \\ a_9 & a_{10} & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \varepsilon_{res} \\ \varepsilon_{exrt} \\ \varepsilon_{\exp g} \\ \varepsilon_{debt} \\ \varepsilon_{rgg} \end{pmatrix} = \begin{pmatrix} v_{res} \\ v_{exrt} \\ v_{\exp g} \\ v_{debt} \\ v_{rgg} \end{pmatrix} = v\,. \qquad (7)$$

Equation (7) is similar to equation (6). With *debt* in the system, it assumes that this newly added variable will directly interact with *res* and *exrt*, but not *expg* and *rgg*. The structural IRF are graphed in Figure 5.4.1, which shows a clearer negative effect of *res* on *exrt* than Figure 3.8 does. In the second last graph of the third column, we see that a positive structural shock in *res* will cause *exrt* to depreciate immediately, and the impact will decrease over time but always be negative. The influence of *res* on *expg* stays mostly positive and has a notable size for the about 10 periods (the second to last graph in the second column). The FEVD graphs in Figure 5.4.2 show that *res* is the most important variable in explaining the short-term forecast errors in *exrt*, but such importance fades away as the forecast horizon increases. For *expg*, *res* accounts for about 10% of the variance in its short-term forecast errors, and as the forecast horizon increases, the proportion gradually reduces to around 5%. These results all look confirmative to the mercantilism explanation for China's reserve holdings and they also suggest that such a policy is effective for China.

3.2 Subsample (1996Q1-2005Q4)

Factors such as exchange rate regime and financial openness are often considered in the empirical literature on reserves. We control these aspects and test the robustness of our previous findings by studying a subsample period, 1996Q1-2005Q4, during which the exchange rate arrangements and foreign reserve management policies for China remained relatively stable[29]. We follow the same investigation scheme as before. The estimations are summarized in Tables 4.1-3, and the Diebold-Mariano test information between predictions is reported in Table 5. As we will see, the results are largely consistent with those for the full sample.

Precautionary demand (1)

In Table 4.1 the model that produces the best predictions for *res* is again the simplest specification with only *res* and *imp* (S1_01). Although some other models have better overall fits for *res* and therefore smaller RMSEs, the test results in Table 5 show that if both the in-sample and out-of-sample forecasts are taken into consideration, predictions by model S1_01 are still the most accurate ones. The only regressor other than *imp* that is cointegrated with *res* by itself is *idif* (S1_02). But due to a larger number of regressors coupled with a poorer fit for the *res* equation, model S1_02 has a much larger RMSE than model S1_01. Another difference between Table 4.1 and Table 2.1 is that model S1_05 (*res*, *imp*, *idif*) is no longer the "second best" in Table 4.1. But for comparison purposes, we will keep it as a benchmark specification for later estimations.

The IRF of model S1_01 are plotted in Figure 6.1. The interactions between *res* and *imp* look qualitatively very similar to those seen in Figure 3.1, that is, *res* will rise following a positive shock in *imp*, and vice versa. Quantitatively, the long-run responses of *res* to shocks in *imp* seem to have larger sizes than the previous results for the full sample, which probably reflects a more cautious standpoint taken by the Chinese policymakers when the economy becomes more open.

As before, model S1_01's prediction errors on *res* graphed in Figure 7.1 suggest 2004Q4 and 2007Q1 as the two largest outliers. It should be noted, however, that unlike for the full

[29] Because of the 1997 Asian Financial Crisis, many previous panel studies take that year as the cutting points for their subsamples. However, China is a special case in at least two senses. First, thanks to an almost closed domestic financial market, the impact of the 1997 Asian Financial Crisis on China was much smaller than those on many other countries. Second, there were no significant changes in China's foreign reserve and exchange rate policies around the time of the financial crisis. In our sample period, other than some new measures adopted in late 2005, China's most significant reform steps regarding to foreign reserves and exchange rates were all taken between 1994 and early 1996. 1994 and 1995 are excluded from our subsample to avoid some transitional fluctuations, especially those caused by the RMB devaluation in 1994.

sample, such a pattern in the prediction errors is not robust across all the subsample models. For instance, model S3_10, which will be discussed later, shows in Figure 7.2 that China also received a big positive shock in 2007Q2. One possible reason why the results become weaker could be the fewer degrees of freedom in the subsample estimations.

Comparing model S1_01 to model F1_01, we see that the full sample estimation has a smaller RMSFE for the *res* equation, which implies better out-of-sample forecasts. The same is also true for most other specifications, suggesting that the early years in the sample period, 1983-1995, probably contain useful information for the predictions of China's reserve holdings.

Precautionary demand (2)

The Diebold-Mariano test results in Table 5 indicate that four models in Table 4.2 have better predictions on *res* than model S1_01. But since no improvements brought about by these models are actually significant, we will keep our focus on model S2_07 and compare its results to model F2_07, which was found earlier to have one of the best predictive performances on *res* for the full sample[30].

The IRF of model S2_07 are plotted in Figure 6.2. The dynamics between *res* and the other regressors are similar to what we saw in Figure 3.4 for model F2_07. The structural form of model S2_07 is estimated with the same assumptions as those for model F2_07, and the structural IRF and FEVD are in Figures 8.1.1 and 8.1.2, respectively. A difference between Figure 8.1.1 and Figure 5.2.1 is that Figure 8.1.1 shows a more evident negative effect of *res* on *vcpid* (the third graph in the last column). Nonetheless, Figure 8.1.2 still suggests that shocks in *res* play almost no role in accounting for the forecast error variance of *vcpid*. Together these results do not bring much change to the conclusions we reached before, that is, inflation volatility is a helpful predictor for China's reserve holdings, but piling up reserves does not appear to be an effective policy tool for China to reduce its inflation volatility.

Modern mercantilism

Previously in the full sample study, in spite of model F3_06's good predictions on *res*, we refrained from claiming it a success for the reason that the dynamics shown by the model were not consistent with its underlying theory. The judgment is easier to make in the subsample case, because model S3_06 does not have good predictions on *res* in the first place. In Table 5 we see that its predictions are worse than those by the benchmark model S1_01, and significantly worse than those of model S3_10. As for the other models in the first panel of Table 4.3, there is no substantial difference between them and their full-sample counterparts in Table 2.3, except that no cointegrating relation is identified for specification S3_03, whereas the regressors were found cointegrated in model F3_03.

[30] The models that beat model S1_01 are S2_05, S2_06, S2_07 and S2_09. Model S2_07 has the best predictions on *res* among them, albeit by insignificant margins.

On the other hand, there are a couple of interesting changes in the second panel of Table 4.3. The first one is that unlike model F3_07, whose predictions on *res* are much worse than those of model F1_01, model S3_07 beats its benchmark, model S1_01. This implies that, at least relative to *imp*, *exrt* has a closer relation with *res* in the subsample period than in the full sample. The other change is that the relative fit of model S3_09 is not as good as model F3_09. In the full sample, the addition of either *expg* or *debt* brings a big improvement to model F3_07, and the fits of models F3_08 and F3_09 are close to each other. In the subsample, adding *expg* still improves model S3_07, but adding *debt* alone actually makes it worse. This suggests that *expg* is probably playing a more crucial role in explaining *res* during the subsample period.

In the full sample, models F3_08-10 have the best predictions on *res* among all the models. The same is still true for models S3_08 and S3_10. In the first part of Table 5, we see that model S3_10 is the only one whose predictions on *res* are significantly better than model S1_01. Although the improvement of model S3_08 over model S1_01 is insignificant, it has the lowest p-value for the Diebold-Mariano test among all the models other than S3_10 which beat model S1_01. The second part of Table 5 shows that model S3_10 has the best predictions on *res* among all the models[31]. In the last part of Table 5, we see that model S3_08 is the only one which significantly outperforms the ARIMA(1,1,1) model in predicting *res*. Among the other models beating ARIMA(1,1,1), model S3_10 has the second lowest p-value, only next to model S3_07.

The structural analyses on models S3_08 and S3_10 also support the modern mercantilism theory. Before checking the structural results, we still first take a look at the nonstructural IRF of the two models, graphed in Figures 6.3 and 6.4, respectively. The effects of a positive shock to *res* are similar in these two figures, that is, it has only temporary negative effects on *exrt* and temporary positive effects on *expg*. Again, taking account of the endogeneity structure among the regressors brings us different stories. The structural forms of models S3_08 and S3_10 are estimated with the same assumptions as before, and their structural IRF are Figure 8.2.1 and Figure 8.3.1, respectively. Both figures suggest that the negative effect of *res* on *exrt* and its positive impact on *expg* are permanent. Besides, the FEVD of model S3_08, Figure 8.2.2, suggests that *res* accounts for about 20% of the forecast error variance in *exrt*, and its share in *expg*'s forecast error variance mostly falls between 7% and 20%. The same proportions shown by Figure 8.3.2, the FEVD of model S3_10, are even bigger. In summary, the evidence supporting the modern mercantilism explanation of China's reserve holdings that we found earlier is robust in the subsample period of 1996Q1-2005Q4, and the difference between the structural and nonstructural results demonstrates once again that the endogeneity among the regressors must be taken into consideration in order to identify the true causal relations among them.

4. Conclusion and Future Work

[31] Tests between model S3_08 and the other models, except model S3_10, have similar results.

This paper tries to empirically explain the stunning increase in China's foreign reserve holdings over the past two decades. Unlike most other studies in the current literature, it investigates the question from a time series perspective, because panel regressions, the commonly adopted methods, face several critical limitations in analyzing reserve holdings.

The difficulties encountered by panel regressions include heterogeneity across the sample countries, nonstationarity of the data series, endogeneity among the regressors and measurement errors associated with the optimal reserve levels. We show in the paper that all these problems can lead to serious bias in the estimation results of panel regressions, and the time-series VEC model, on the other hand, can handle them better. Another consideration favoring the VEC model is that its dynamic analysis allows us to evaluate the effectiveness of China's reserve policy in achieving different goals.

The VEC models estimated in the paper are motivated by two strands of early work on reserves, the precautionary demand theories and the modern mercantilism theory. For both the full sample (1983Q1-2005Q4) and the subsample (1996Q1-2005Q4), we find that models based on modern mercantilism perform better in predicting China's foreign reserves than selected ARIMA models and those constructed on precautionary grounds, both with statistical significance. This suggests that previous econometric studies, whose emphasis has been overwhelmingly placed on precautionary motives, might have missed some more important factors in explaining China's reserve holdings. Further structural analyses on these models show that the buildup of reserve stocks by China has a negative (depreciating) effect on its real exchange rate and a positive impact on its export growth, which corroborates the conjecture that the recent increase in China's reserve holdings is a part of its export-led development strategy. When interpreted in the broader context of current global imbalances, such results seem to support the view that China should play a more active role in the adjustment process by reevaluating its currency and/or adopting a more flexible exchange rate regime.

Among considerations related to precautionary motives, imports and associated risks, such as real openness of the economy, appear to be the most important determinants for China's reserves holdings. Volatility of inflation and the interest rate differential between China and the US are also helpful predictors. But our results suggest that if China wants to reduce its inflation volatility by piling up reserves, it has not been an effective policy tool. The estimated impact of interest rate differential on China's reserve holdings is opposite to that predicted by the Frenkel and Jovanovic buffer stock model, probably because higher returns to capital in China, reflected by higher interest rate differentials, attract more capital inflows to China and therefore push its reserve levels up. We find little evidence to support volatility of exports and domestic financial depth as crucial factors in explaining China's reserve holdings.

Due to a relatively short sample period, the observations available for our estimations are not particularly abundant, which is especially evident when we study the subsample. One possible way to alleviate such a scarcity of observations is to use information with higher frequency, such as monthly data. Using such data will also bring an extra benefit for the structural analysis, because the identification assumptions imposed for structural estimations often hold better with a shorter period span.

We have argued that panel regressions are not the most appropriate tools for analyzing foreign reserve holdings. But it by no means implies that cross-country information is not important for the understanding of reserve holdings. Rather, our work suggests that a better way to exploit such information is probably to first apply the method of this paper to individual countries, correctly identify the most relevant determinants for each one of them, and then do the cross-country comparison.

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Table - 1 Illustration of Spurious Regression

Panel 1. Johansen Cointegration Test

Variables: *res* , *imp* Sample period: 1983Q1-2005Q4

Country	Obs.	Lags[!]	Rank = 0			Rank = 1		
			Trace statistic	10% cv	1% cv	Trace statistic	10% cv	1% cv
China	91	1	27.53	13.43	20.04	2.67*	2.71	6.65
Korea	91	1	9.92*	13.43	20.04	1.01	2.71	6.65

[!] Selection of lag order is based on Schwarz-Bayesian information criterion

* Rank of cointegration suggested by test results

Panel 2. OLS and Panel Regressions

Country	Dep. Var.	Obs.	Coef. of *imp*	t-stat	Adj. R2
China	*res*	92	1.31	14.13	0.69
Korea	*res*	92	0.89	3.25	0.10
China & Korea	*res*	184	0.34	4.19	0.08

Table 2.1 VEC Models - Precautionary Demand (1), 1983Q1-2005Q4

Model	Regressors	# obs.	# lag	# ci.	R^{2*}	RMSE	RMSFE	Implied long-run relation, res=				Mean Prediction Error						
								imp[#] (5.43)	idif (6.04)	debt (2.08)	vexp (0.48)	83-95	96-05	06-07	04Q4	96-05 w/o 04Q4	07Q1	06-07 w/o 07Q1
F1_01[%]	imp	90	2	1	0.39	0.768	1.517	5.62 (1.48)				-0.05	0.07	0.38	2.67	0.00	4.07	-0.14
F1_05	imp idif	90	2	2	0.39	0.776	1.535	4.49 (2.79)	0.39 (0.49)			-0.05	0.06	0.45	2.66	0.00	4.14	-0.08
F1_06	imp debt	89	3	1	0.35	0.817	1.809	1.39 (0.35)		0.43 (0.82)		-0.14	0.17	0.88	3.36	0.09	4.82	0.31
F1_07	imp vexp	90	2	1	0.36	0.793	1.730	2.4 (202.1)			-3.05 (2046)	-0.16	0.20	0.90	2.81	0.13	4.64	0.37
F1_08	idif debt	89	3	1	0.32	0.833	1.685		4.23 (5.23)	1.38 (11.38)		-0.13	0.16	0.54	3.73	0.07	4.42	-0.02
F1_09	idif vexp	90	2	1	0.25	0.858	1.640		2.14 (0.9)		1.43 (8.99)	-0.15	0.19	0.61	3.58	0.11	4.47	0.06
F1_10	debt vexp	87	5	1	0.40	0.830	1.737			19.76 (14.88)	-93.38 (79.59)	-0.11	0.13	0.65	3.05	0.06	4.39	0.12
F1_11	imp idif debt	89	3	1	0.38	0.810	1.720	1.69 (1.7)	-0.57 (0.71)	0.21 (3.34)		-0.14	0.17	0.75	3.32	0.09	4.58	0.21
F1_12	imp idif vexp	90	2	2	0.36	0.799	1.731	2.4 (6.16)	0.11 (1.59)		-2.48 (27.68)	-0.15	0.19	0.90	2.79	0.12	4.64	0.37
F1_13	imp debt vexp	90	2	1	0.29	0.837	1.678	1.14 (1.24)		-0.72 (1.1)	3.36 (7.62)	-0.13	0.16	0.71	3.28	0.08	4.56	0.16
F1_14	idif debt vexp	89	3	1	0.33	0.840	1.641		2.5 (1.46)	7.29 (8.86)	-27.53 (34.18)	-0.13	0.15	0.40	3.60	0.07	4.25	-0.15
F1_15	imp idif debt vexp	89	3	1	0.39	0.814	1.658	2.02 (2.09)	-0.62 (0.93)	-2.69 (7.88)	11.93 (21.59)	-0.14	0.17	0.63	3.15	0.09	4.41	0.08

* Statistics and coefficients reported in this table are those of the *res* equations.

[#] In parentheses are standard deviations of regressors and estimated coefficients.

[%] Inconsecutive numbering of models is due to specifications for which no cointegrating relations are identified. Models missing include: F1_02 (*res* , *idif*), F1_03 (*res* , *debt*) and F1_04 (*res* , *vexp*).

Table 2.2 VEC Models - Precautionary Demand (2), 1983Q1-2005Q4

Model	Regressors	# obs.	# lag	# ci.	R^{2*}	RMSE	RMSFE	Implied long-run relation, res=				Mean Prediction Error						
								imp[#] (5.43)	idif (6.04)	vcpid (0.34)	m2gdp (0.38)	83-95	96-05	06-07	04Q4	96-05 w/o 04Q4	07Q1	06-07 w/o 07Q1
F2_04[%]	imp vcpid	90	2	1	0.44	0.743	1.520	5.72 (2.83)		11.05 (45.39)		-0.09	0.11	0.35	2.31	0.06	4.11	-0.18
F2_05	imp m2gdp	91	1	1	0.31	0.805	1.501	-9.28 (10.78)			60.5 (80.09)	-0.09	0.12	0.14	3.08	0.04	3.75	-0.38
F2_06	imp vcpid m2gdp	91	1	1	0.32	0.799	1.498	-9.51 (21.38)		-26.2 (42.88)	56.31 (248.8)	-0.10	0.13	0.12	3.06	0.06	3.73	-0.39
F2_07	imp idif vcpid	90	2	2	0.45	0.744	1.551	4.04 (4.62)	0.91 (0.54)	7.34 (46.94)		-0.07	0.08	0.45	2.28	0.03	4.22	-0.09
F2_08	imp idif m2gdp	90	2	1	0.34	0.806	1.590	0.68 (6.3)	4.55 (3.23)		-20.31 (167.7)	-0.15	0.18	0.70	2.85	0.12	4.38	0.17

* Statistics and coefficients reported in this table are those of the *res* equations.

[#] In parentheses are standard deviations of regressors and estimated coefficients.

[%] Models missing include: F2_01 (*res* , *vcpid*), F2_02 (*res* , *m2gdp*), F2_03 (*res* , *vcpid* , *m2gdp*) and F2_09 (*res, imp, idif, vcpid, m2gdp*).

Table 2.3 VEC Models - Modern Mercantilism, 1983Q1-2005Q4

Model	Regressors	# obs.	# lag	# ci.	R^{2*}	RMSE	RMSFE	Implied long-run relation, res=				Mean Prediction Error						
								imp[#] (5.43)	idif (6.04)	penn (30.0)	expg (13.1)	83-95	96-05	06-07	04Q4	96-05 w/o 04Q4	07Q1	06-07 w/o 07Q1
F3_01	penn	90	2	1	0.29	0.830	1.543			0.31 (0.66)		-0.03	0.04	0.37	3.42	-0.05	4.15	-0.17
F3_02	penn expg	91	1	2	0.29	0.820	1.500			0.16 (0.29)	-0.48 (0.62)	-0.03	0.03	0.32	3.26	-0.05	3.87	-0.19
F3_03	imp penn	90	2	1	0.40	0.766	1.544	4.08 (25.08)		0.26 (3.49)		-0.09	0.11	0.48	2.55	0.05	4.16	-0.05
F3_04	imp penn expg	88	4	2	0.41	0.820	1.655	3.47 (15.97)		0.33 (3.21)	-0.61 (2.17)	-0.09	0.11	0.46	2.83	0.04	4.27	-0.08
F3_05	imp idif penn	90	2	2	0.40	0.773	1.600	3.28 (1.46)	0.27 (0.76)	0.21 (0.1)		-0.09	0.12	0.63	2.55	0.06	4.31	0.10
F3_06	imp idif penn expg	88	4	3	0.48	0.791	1.563	3.57 (2.33)	-0.21 (1.4)	0.27 (0.15)	-0.19 (0.61)	-0.08	0.10	0.44	2.45	0.04	3.97	-0.07

Model	Regressors	# obs.	# lag	# ci.	R^2	RMSE	RMSFE	Implied long-run relation, res=				Mean Prediction Error						
								rgg (3.52)	debt (2.08)	exrt (0.31)	expg (13.1)	83-95	96-05	06-07	04Q4	96-05 w/o 04Q4	07Q1	06-07 w/o 07Q1
F3_07	rgg exrt	91	1	1	0.29	0.815	1.478	-0.67 (0.13)		14.48 (17.6)		0.01	-0.01	0.19	3.29	-0.10	3.71	-0.31
F3_08	rgg exrt expg	86	6	1	0.52	0.789	1.446	-0.84 (1.15)		113.3 (206.2)	5.67 (4.51)	-0.06	0.07	0.59	2.85	0.00	3.54	0.16
F3_09	rgg debt exrt	87	5	2	0.53	0.760	1.473	1.42 (2)	0.32 (2.64)	-19.63 (8.09)		-0.09	0.11	0.05	2.49	0.05	3.41	-0.43
F3_10	rgg debt exrt expg	86	6	3	0.58	0.780	1.373	1.51 (19.63)	2.45 (30.19)	-72.38 (59.82)	-1.49 (6.38)	-0.10	0.12	0.31	2.32	0.06	3.29	-0.12

* Statistics and coefficients reported in this table are those of the *res* equations.

[#] In parentheses are standard deviations of regressors and estimated coefficients.

Table 3 Comparison of Predictions on Reserves (Full Sample) *

Benchmark 1: F1_01				Benchmark 2: F3_10				Benchmark 3: ARIMA(4,1,2)		
Model	Predictions relative to benchmark	P-value#		Model	Predictions relative to benchmark	P-value		Model	Predictions relative to benchmark	P-value
								F1_01	better	0.522
F1_05	worse	0.635		F1_05	worse	0.083		F1_05	better	0.55
F1_06	worse	0.199		F1_06	worse	0.035		F1_06	worse	0.125
F1_07	worse	0.14		F1_07	worse	0.03		F1_07	worse	0.392
F1_08	worse	0.164		F1_08	worse	0.017		F1_08	worse	0.138
F1_09	worse	0.09		F1_09	worse	0.013		F1_09	worse	0.075
F1_10	worse	0.399		F1_10	worse	0.026		F1_10	worse	0.847
F1_11	worse	0.322		F1_11	worse	0.031		F1_11	worse	0.56
F1_12	worse	0.158		F1_12	worse	0.037		F1_12	worse	0.482
F1_13	worse	0.128		F1_13	worse	0.019		F1_13	worse	0.136
F1_14	worse	0.143		F1_14	worse	0.018		F1_14	worse	0.33
F1_15	worse	0.379		F1_15	worse	0.036		F1_15	better	0.881
F2_04	better	0.147		F2_04	worse	0.093		F2_04	better	0.251
F2_05	worse	0.015		F2_05	worse	0.026		F2_05	worse	0.801
F2_06	worse	0.028		F2_06	worse	0.03		F2_06	worse	0.899
F2_07	better	0.082		F2_07	worse	0.14		F2_07	better	0.226
F2_08	worse	0.099		F2_08	worse	0.038		F2_08	worse	0.849
F3_01	worse	0.088		F3_01	worse	0.023		F3_01	worse	0.476
F3_02	worse	0.046		F3_02	worse	0.027		F3_02	worse	0.702
F3_03	better	0.947		F3_03	worse	0.032		F3_03	better	0.477
F3_04	worse	0.579		F3_04	worse	0.045		F3_04	better	0.566
F3_05	worse	0.724		F3_05	worse	0.032		F3_05	better	0.549
F3_06	better	0.075		F3_06	worse	0.15		F3_06	better	0.146
F3_07	worse	0.096		F3_07	worse	0.031		F3_07	worse	0.757
F3_08	better	0.193		F3_08	worse	0.074		F3_08	better	0.049
F3_09	better	0.16		F3_09	worse	0.042		F3_09	better	0.065
F3_10	better	0.074						F3_10	better	0.033

* Results reported in this table are based on the Diebold-Mariano test. The Davidson-MacKinnon test results, not reported, are largely consistent.

The null is that the difference in predictions is not significant. A large p-value implies that we cannot reject the null.

Table 4.1 VEC Models - Precautionary Demand (1), 1996Q1-2005Q4

Model	Regressors	# obs.	# lag	# ci.	R^{2*}	RMSE	RMSFE	Implied long-run relation, res=				Mean Prediction Error						
								imp# (5.66)	idif (1.61)	debt (1.48)	vexp (0.41)	96-00	01-05	06-07	04Q4	01-05 w/o 04Q4	07Q1	06-07 w/o 07Q1
S1_01%	imp	38	2	1	0.63	0.738	1.665	2.8 (1.36)				-0.08	0.07	0.74	2.25	-0.05	4.20	0.25
S1_02	idif	36	4	1	0.60	0.838	1.650		-10.46 (3.58)			-0.25	0.20	-0.09	2.35	0.08	3.07	-0.54
S1_05	imp idif	39	1	1	0.49	0.827	1.798	2.71 (1.22)	0.73 (2.34)			-0.14	0.13	0.93	3.14	-0.03	4.52	0.42
S1_06	imp debt	39	1	1	0.50	0.822	1.790	3.19 (7.56)		3.38 (14.73)		-0.25	0.24	0.93	2.96	0.10	4.53	0.42
S1_08	idif debt	36	4	1	0.61	0.875	1.669		-5.41 (2.94)	-0.71 (3.35)		-0.21	0.17	0.16	2.22	0.06	3.26	-0.28
S1_09	idif vexp	36	4	2	0.72	0.755	1.941		-24.51 (4.74)		-157 (44.03)	-0.15	0.12	0.06	1.78	0.04	3.08	-0.37
S1_11	imp idif debt	39	1	1	0.44	0.873	2.167	2.62 (4.62)	1.66 (2.96)	3.12 (13.35)		-0.30	0.28	1.45	3.42	0.12	5.10	0.93
S1_12	imp idif vexp	36	4	1	0.79	0.690	2.404	1.38 (0.79)	2.97 (4.7)		35.38 (108.2)	-0.15	0.12	0.01	1.80	0.03	3.06	-0.42
S1_14	idif debt vexp	36	4	1	0.73	0.774	1.970		2.91 (6.14)	-2.38 (17.43)	54.41 (16.45)	-0.11	0.09	0.04	1.70	0.01	2.88	-0.37
S1_15	imp idif debt vexp	36	4	1	0.84	0.649	2.501	0.7 (2.62)	2.29 (5.53)	-1.32 (32.26)	26.24 (97.07)	-0.10	0.08	0.13	1.19	0.02	2.98	-0.28

* Statistics and coefficients reported in this table are those of the *res* equations.

\# In parentheses are standard deviations of regressors and estimated coefficients.

% Models missing include: S1_03 (*res, debt*) and S1_04 (*res, vexp*), S1_07 (*res, imp, vexp*), S1_10 (*res, debt, vexp*) and S1_13 (*res, imp, debt, vexp*).

Table 4.2 VEC Models - Precautionary Demand (2), 1996Q1-2005Q4

Model	Regressors	# obs.	# lag	# ci.	R^{2*}	RMSE	RMSFE	Implied long-run relation, res=				Mean Prediction Error						
								imp# (5.66)	idif (1.61)	vcpid (0.34)	m2gdp (0.20)	96-00	01-05	06-07	04Q4	01-05 w/o 04Q4	07Q1	06-07 w/o 07Q1
S2_03%	vcpid m2gdp	39	1	1	0.48	0.838	1.514			-317.5 (101.9)	-956.3 (411.4)	-0.19	0.18	0.02	2.94	0.04	3.70	-0.51
S2_04	imp vcpid	38	2	1	0.68	0.698	1.902	2.05 (4.59)		5.43 (26.49)		-0.13	0.12	1.14	1.82	0.03	4.61	0.65
S2_05	imp m2gdp	38	2	1	0.66	0.716	1.670	2.68 (2.17)			-16.68 (28.23)	-0.13	0.12	0.84	2.02	0.02	4.20	0.36
S2_06	imp vcpid m2gdp	38	2	1	0.72	0.662	1.803	1.62 (3.18)		6.95 (14.75)	9.03 (8.39)	-0.09	0.08	1.06	1.55	0.00	4.46	0.57
S2_07	imp idif vcpid	36	4	1	0.82	0.639	1.764	2.17 (1.33)	1.84 (5.27)	7.24 (33.52)		-0.10	0.08	0.52	1.14	0.02	3.55	0.09
S2_08	imp idif m2gdp	36	4	2	0.80	0.676	2.093	3.08 (1.39)	0.61 (2.02)		-19.99 (21.88)	-0.05	0.04	0.88	1.25	-0.02	3.50	0.51
S2_09	imp idif vcpid m2gdp	39	1	2	0.58	0.767	1.513	1.1 (2.94)	-2.27 (4.26)	21.16 (23.81)	59.04 (117.6)	-0.14	0.14	0.36	2.24	0.03	3.91	-0.14

* Statistics and coefficients reported in this table are those of the *res* equations.

In parentheses are standard deviations of regressors and estimated coefficients.

% Models missing include: S2_01 (*res* , *vcpid*) and S2_02 (*res* , *m2gdp*).

Table 4.3 VEC Models - Modern Mercantilism, 1996Q1-2005Q4

Model	Regressors	# obs.	# lag	# ci.	R^{2*}	RMSE	RMSFE	Implied long-run relation, res=				Mean Prediction Error						
								imp[#] (5.66)	idif (1.61)	penn (6.41)	expg (14.36)	96-00	01-05	06-07	04Q4	01-05 w/o 04Q4	07Q1	06-07 w/o 07Q1
S3_01[%]	penn	39	1	1	0.50	0.825	1.503			-16.38 (12.71)		-0.30	0.29	0.39	3.10	0.14	3.80	-0.10
S3_02	penn expg	38	2	1	0.43	0.931	1.574			-1.3 (0.27)	-0.5 (0.13)	-0.31	0.28	0.40	3.40	0.12	4.24	-0.15
S3_04	imp penn expg	38	2	2	0.61	0.791	1.596	7.05 (6.83)		2.87 (3.46)	-0.18 (0.33)	-0.07	0.06	0.47	2.39	-0.06	4.25	-0.07
S3_05	imp idif penn	36	4	3	0.79	0.718	2.415	1.4 (0.28)	-0.1 (1.33)	-0.43 (0.33)		-0.04	0.03	1.48	1.75	-0.06	4.43	1.06
S3_06	imp idif penn expg	36	4	4	0.87	0.639	2.476	2.58 (35.87)	1.2 (169.3)	-0.87 (34.98)	-0.12 (8.01)	0.03	-0.03	0.61	1.14	-0.09	3.12	0.25

Model	Regressors	# obs.	# lag	# ci.	R^2	RMSE	RMSFE	Implied long-run relation, res=				Mean Prediction Error						
								rgg (2.28)	debt (1.48)	exrt (0.05)	expg (14.36)	96-00	01-05	06-07	04Q4	01-05 w/o 04Q4	07Q1	06-07 w/o 07Q1
S3_07	rgg exrt	39	1	2	0.60	0.747	1.539	-12.9 (13.99)		302.4 (405.4)		-0.26	0.24	0.65	2.96	0.10	3.46	0.25
S3_08	rgg exrt expg	38	2	2	0.66	0.737	1.543	-19.98 (29.54)		-218.3 (264.6)	-1.01 (0.94)	-0.14	0.12	0.82	3.03	-0.03	3.49	0.44
S3_09	rgg debt exrt	39	1	1	0.56	0.776	1.771	114.4 (46.8)	-21.97 (31.65)	-34.82 (87.41)		-0.28	0.27	1.09	3.19	0.11	3.76	0.71
S3_10	rgg debt exrt expg	36	4	1	0.75	0.798	1.421	-1.5 (7.06)	-2.82 (5.16)	-174.6 (181.1)	-1.21 (1.75)	-0.11	0.09	0.44	2.03	-0.01	2.87	0.09

* Statistics and coefficients reported in this table are those of the *res* equations.

[#] In parentheses are standard deviations of regressors and estimated coefficients.

[%] Model missing is S3_03 (*imp, penn*).

Table 5 Comparison of Predictions on Reserves (Subsample) *

Benchmark 1: S1_01				Benchmark 2: S3_10				Benchmark 3: ARIMA(1,1,1)		
Model	Predictions relative to benchmark	P-value#		Model	Predictions relative to benchmark	P-value		Model	Predictions relative to benchmark	P-value
								S1_01	better	0.261
S1_02	worse	0.904		S1_02	worse	0.009		S1_02	better	0.557
S1_05	worse	0.047		S1_05	worse	0.031		S1_05	worse	0.403
S1_06	worse	0.041		S1_06	worse	0.026		S1_06	worse	0.41
S1_08	worse	0.906		S1_08	worse	0.013		S1_08	better	0.567
S1_09	worse	0.722		S1_09	worse	0.119		S1_09	better	0.79
S1_11	worse	0.08		S1_11	worse	0.053		S1_11	worse	0.239
S1_12	worse	0.405		S1_12	worse	0.197		S1_12	worse	0.708
S1_14	worse	0.709		S1_14	worse	0.123		S1_14	better	0.817
S1_15	worse	0.423		S1_15	worse	0.224		S1_15	worse	0.686
S2_03	worse	0.615		S2_03	worse	0.021		S2_03	better	0.576
S2_04	worse	0.524		S2_04	worse	0.144		S2_04	better	0.751
S2_05	better	0.26		S2_05	worse	0.077		S2_05	better	0.203
S2_06	better	0.894		S2_06	worse	0.198		S2_06	better	0.327
S2_07	better	0.243		S2_07	worse	0.418		S2_07	better	0.283
S2_08	worse	0.753		S2_08	worse	0.291		S2_08	better	0.877
S2_09	better	0.775		S2_09	worse	0.03		S2_09	better	0.24
S3_01	worse	0.717		S3_01	worse	0.028		S3_01	better	0.376
S3_02	worse	0.256		S3_02	worse	0.033		S3_02	worse	0.761
S3_04	better	0.7		S3_04	worse	0.035		S3_04	better	0.244
S3_05	worse	0.466		S3_05	worse	0.27		S3_05	worse	0.723
S3_06	worse	0.498		S3_06	worse	0.279		S3_06	worse	0.779
S3_07	better	0.514		S3_07	worse	0.055		S3_07	better	0.15
S3_08	better	0.233		S3_08	worse	0.198		S3_08	better	0.024
S3_09	worse	0.164		S3_09	worse	0.026		S3_09	better	0.94
S3_10	better	0.035						S3_10	better	0.156

* Results reported in this table are based on the Diebold-Mariano test.
The Davidson-MacKinnon test results, not reported, are largely consistent but often more significant.

The null is that the difference in predictions is not significant. A large p-value implies that we cannot reject the null.

Figure 1   Time Trend of China's Foreign Reserve Holdings (1982−2007)

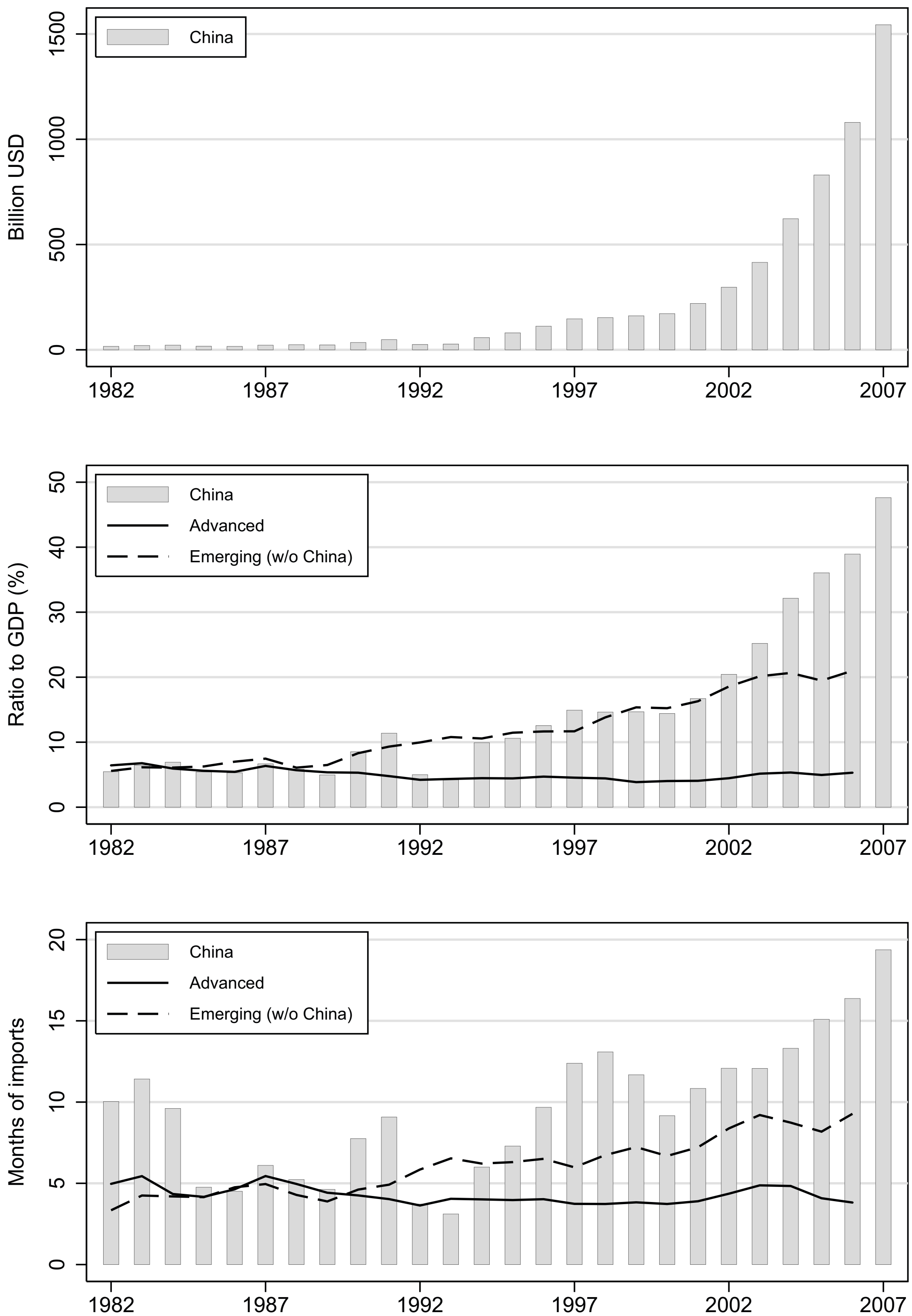
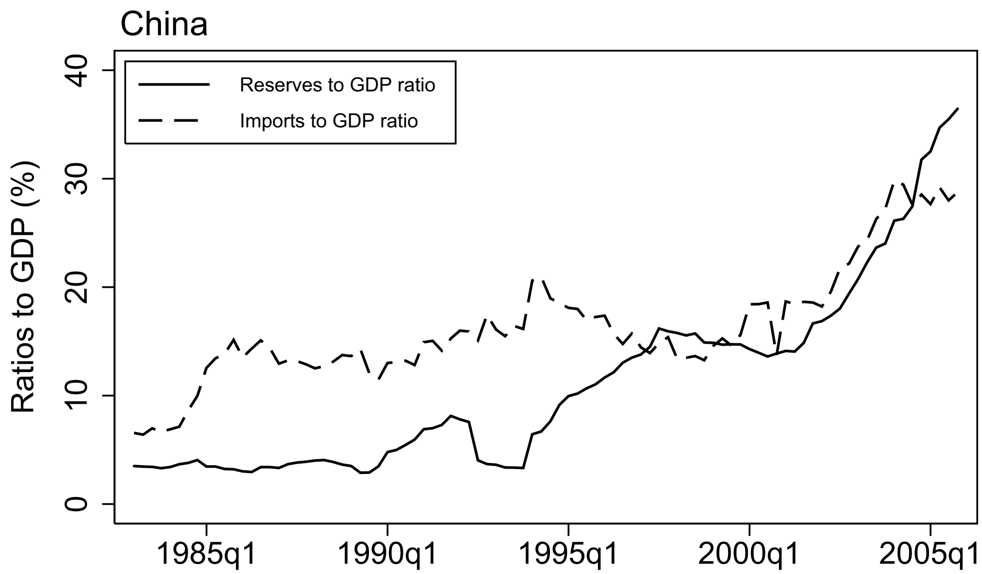


Figure 2 Reserves and Imports (China and Korea, 1983Q1−2005Q4)



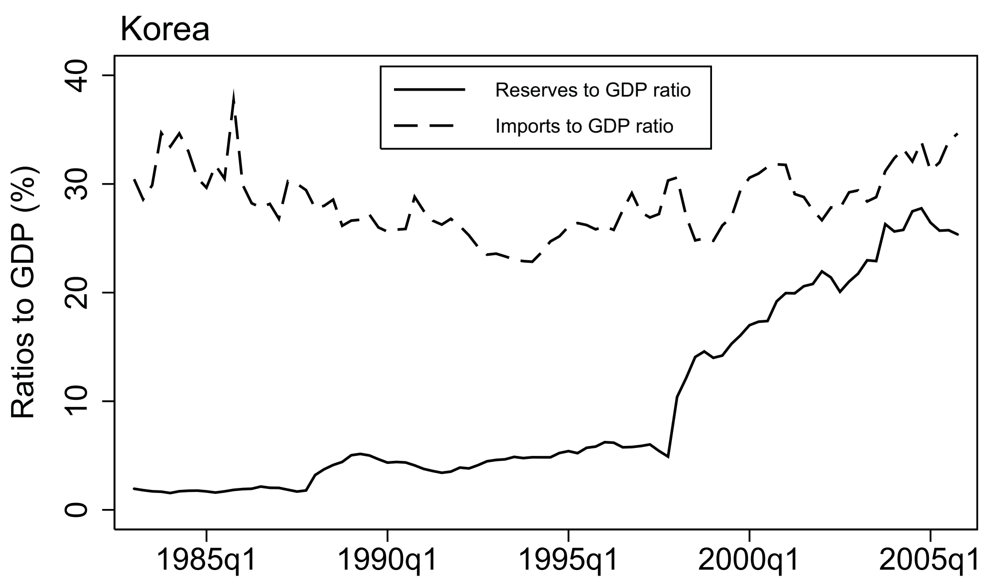
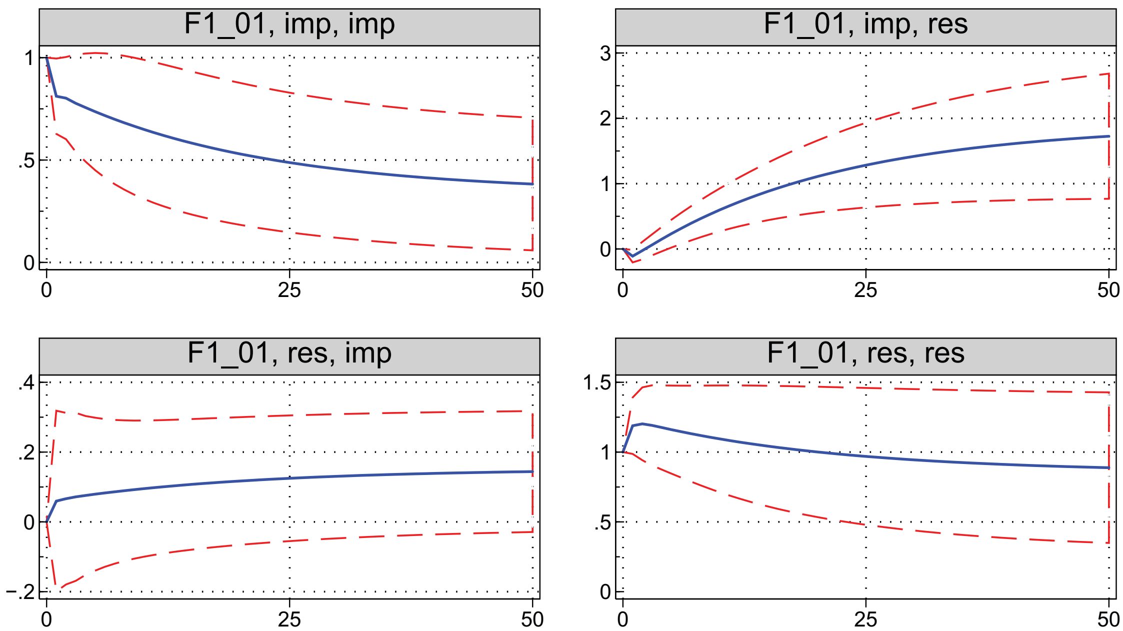
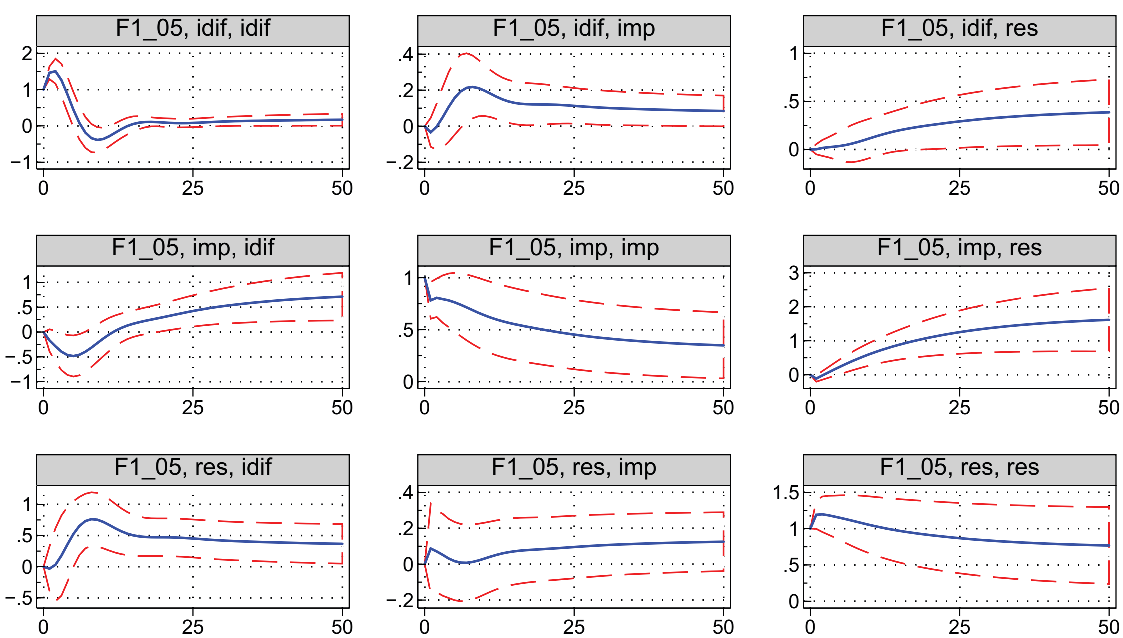


Figure 3.1 IRF of VEC Model F1_01 (Full Sample)



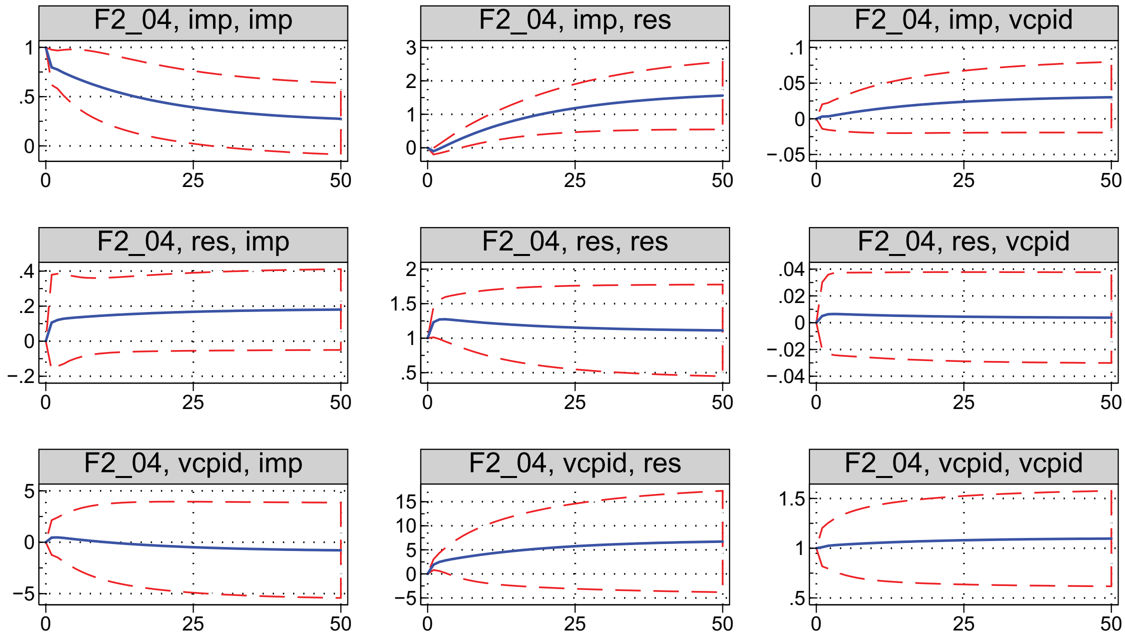
Subtitle: model, shock, response. Std(res)=8.57 Std(imp)=5.43

Figure 3.2 IRF of VEC Model F1_05 (Full Sample)



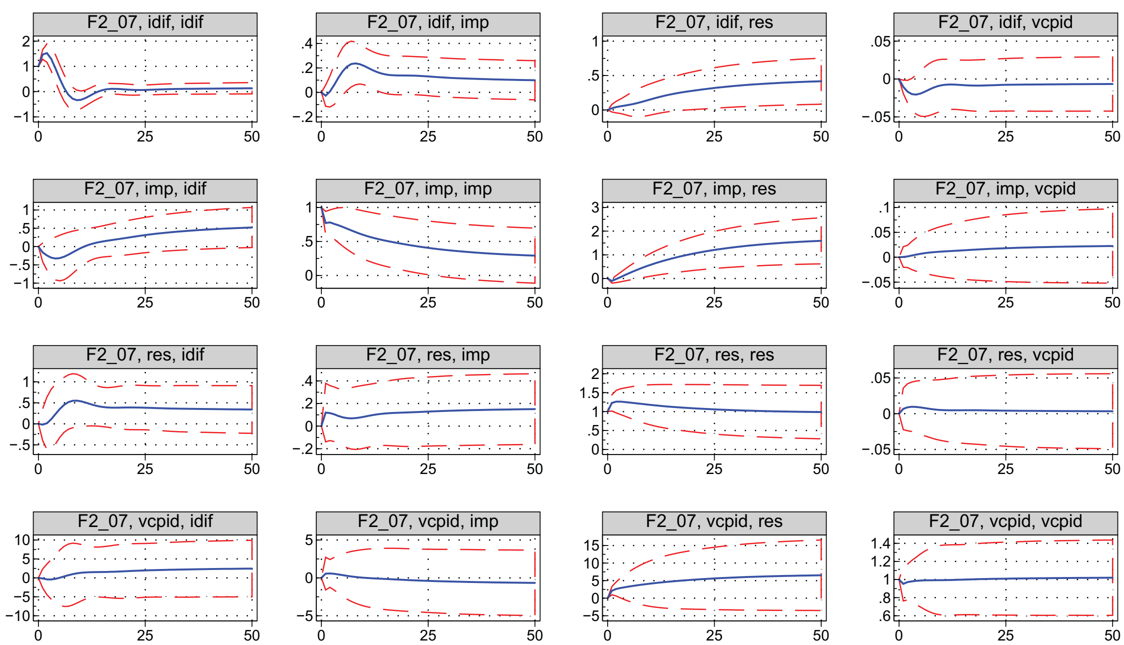
Subtitle: model, shock, response. Std(res)=8.57 Std(imp)=5.43 Std(idif)=6.05

Figure 3.3 IRF of VEC Model F2_04 (Full Sample)



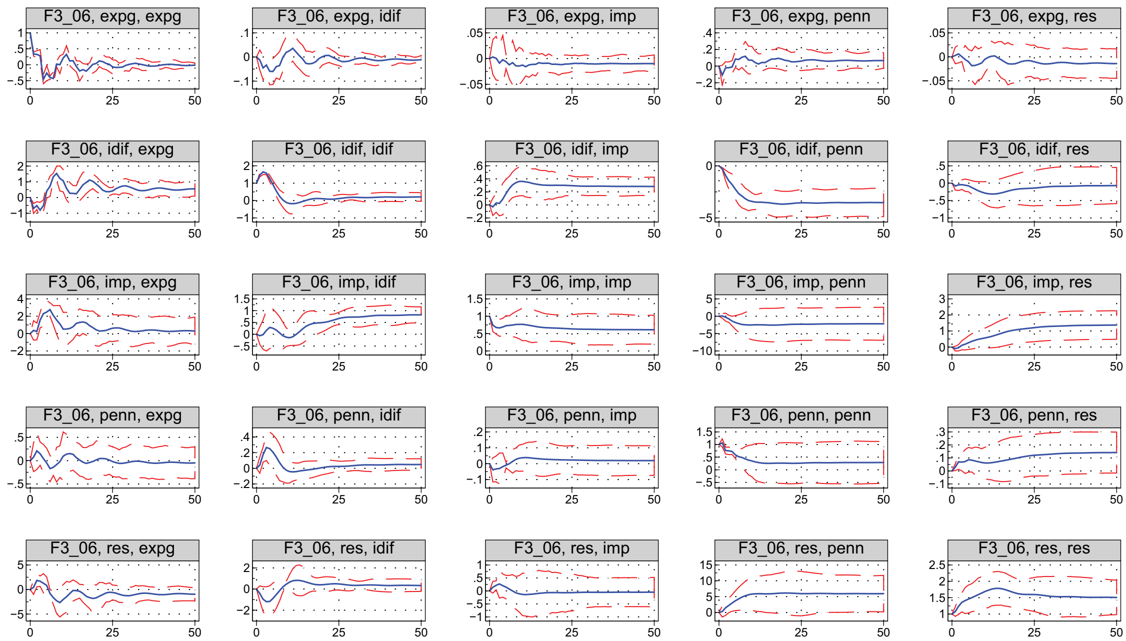
Subtitle: model, shock, response. Std(res)=8.57 Std(imp)=5.43 Std(vcpid)=.34

Figure 3.4 IRF of VEC Model F2_07 (Full Sample)



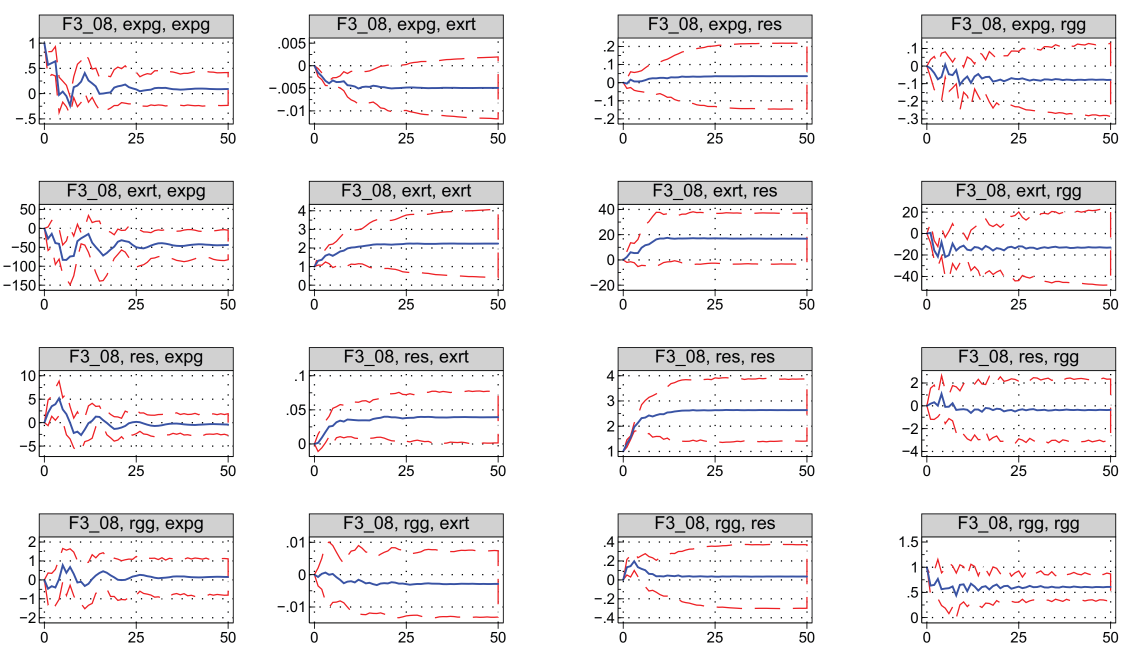
Subtitle: model, shock, response. Std(res)=8.57 Std(imp)=5.43 Std(idif)=6.05 Std(vcpid)=.34

Figure 3.5 IRF of VEC Model F3_06 (Full Sample)



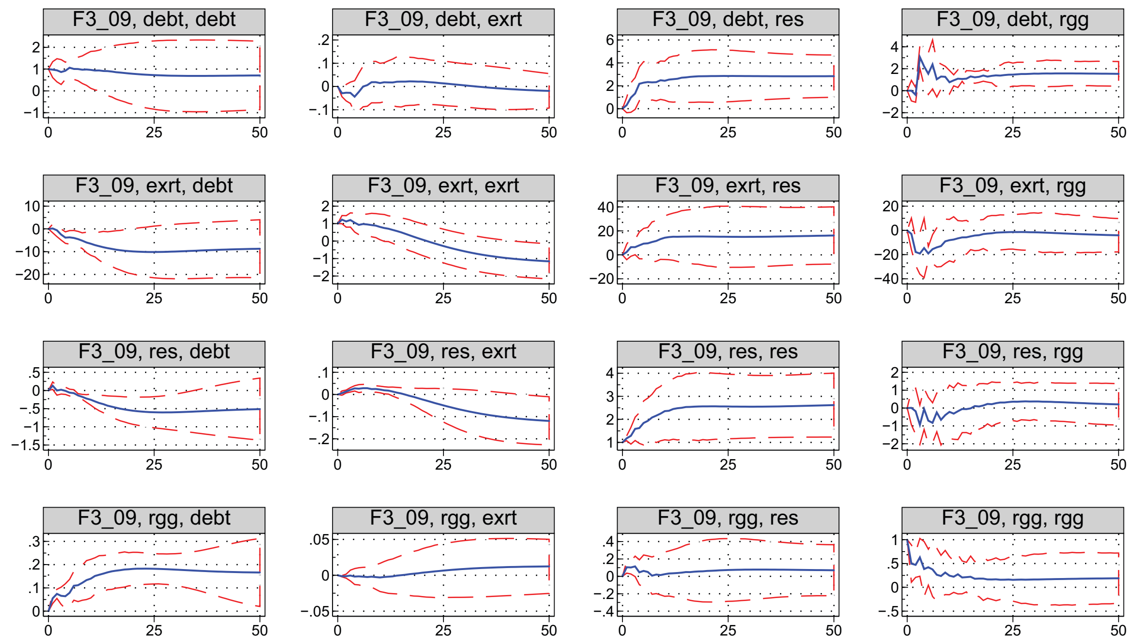
Subtitle: model, shock, response. Std(res)=8.57 Std(imp)=5.43 Std(idif)=6.05 Std(penn)=30.03 Std(expg)=13.11

Figure 3.6 IRF of VEC Model F3_08 (Full Sample)



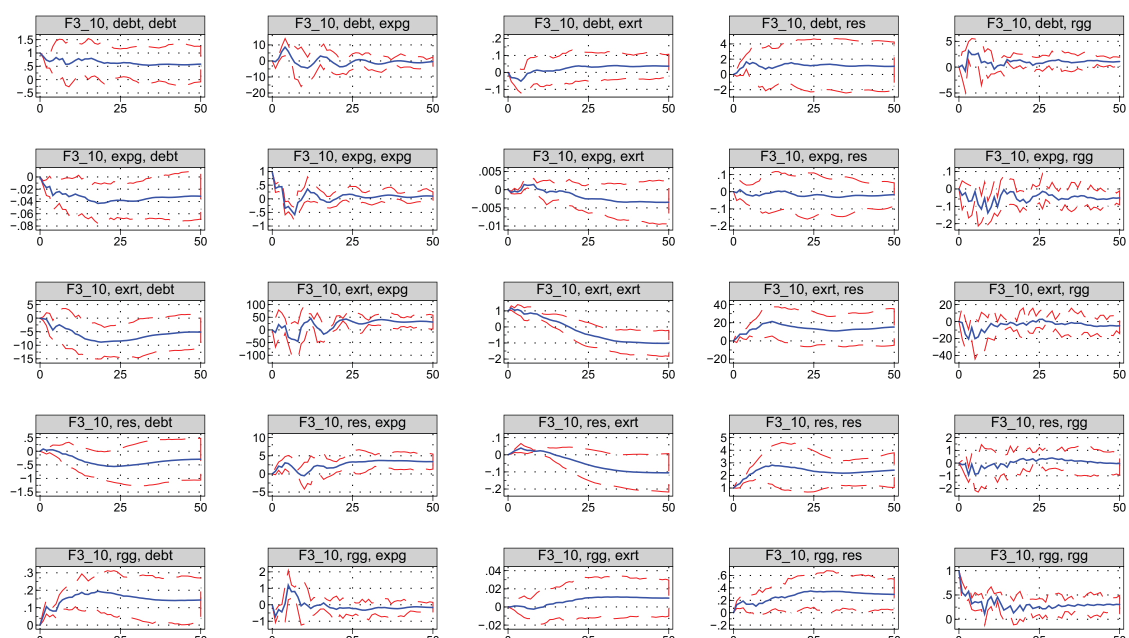
Subtitle: model, shock, response. Std(res)=8.57 Std(rgg)=3.52 Std(exrt)=.3 Std(expg)=13.11

Figure 3.7 IRF of VEC Model F3_09 (Full Sample)



Subtitle: model, shock, response. Std(res)=8.57 Std(rgg)=3.52 Std(debt)=2.08 Std(exrt)=.3

Figure 3.8 IRF of VEC Model F3_10 (Full Sample)



Subtitle: model, shock, response. Std(res)=8.57 Std(rgg)=3.52 Std(debt)=2.08 Std(exrt)=.3 Std(expg)=13.11

Figure 4   Prediction Errors by VEC Model (F1_01)

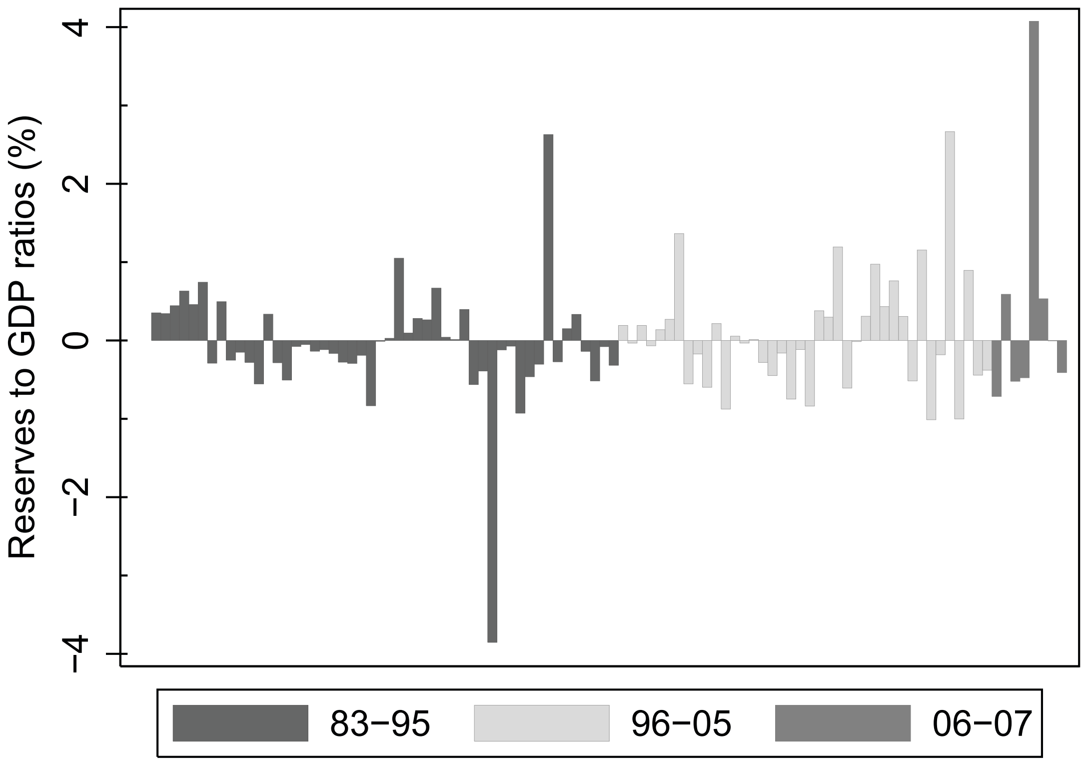
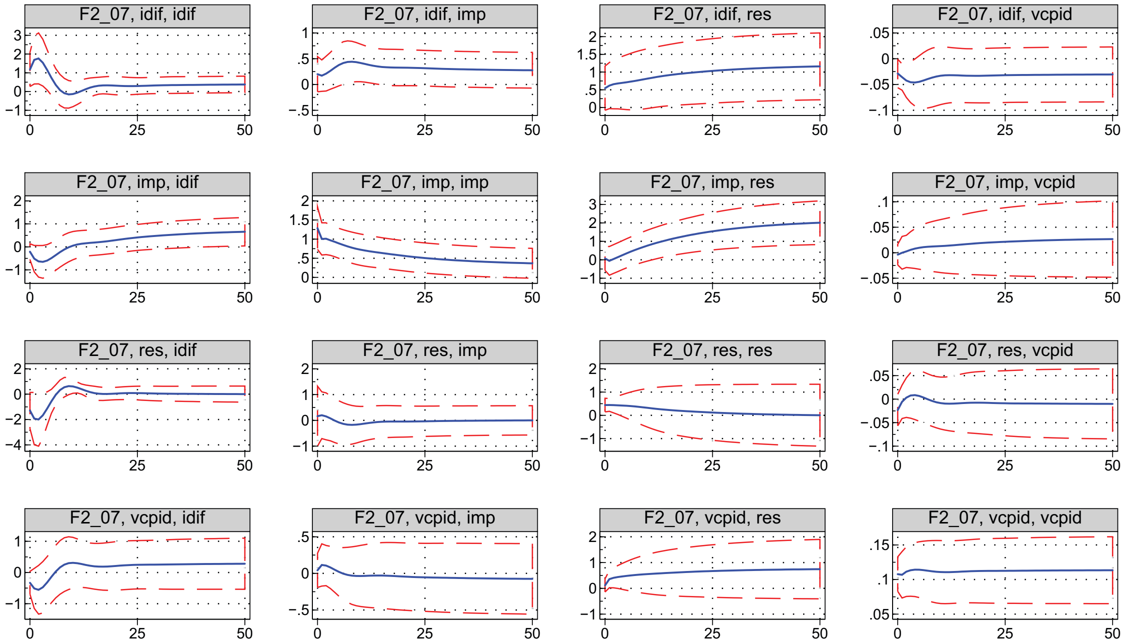
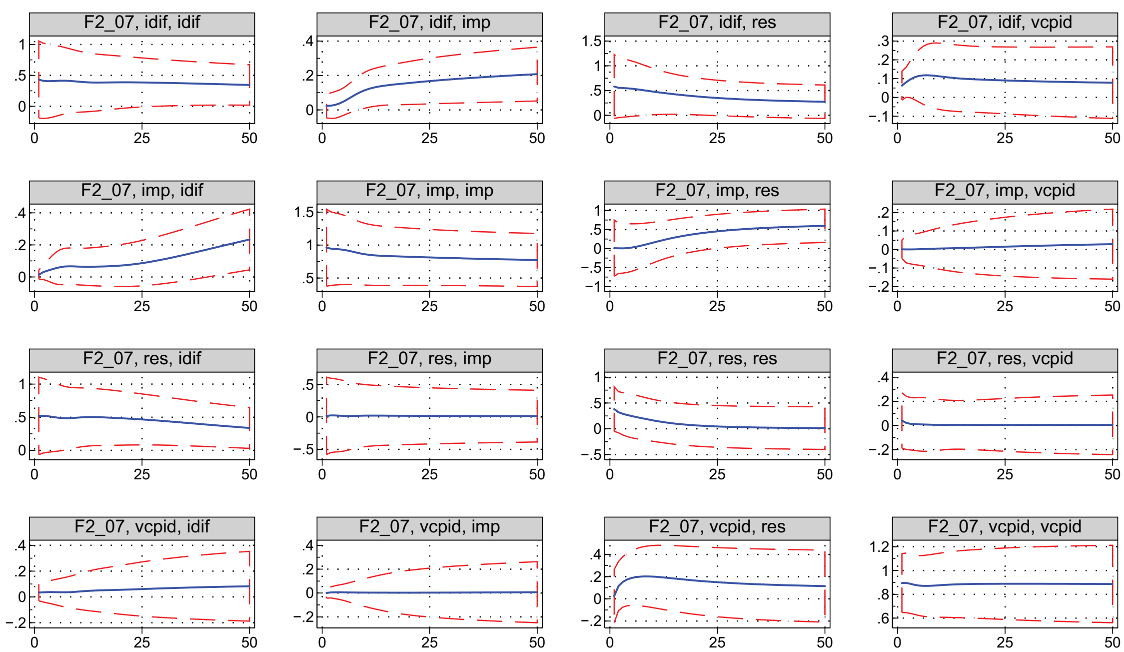


Figure 5.1.1 Structural IRF of VEC Model F2_07 (Full Sample)



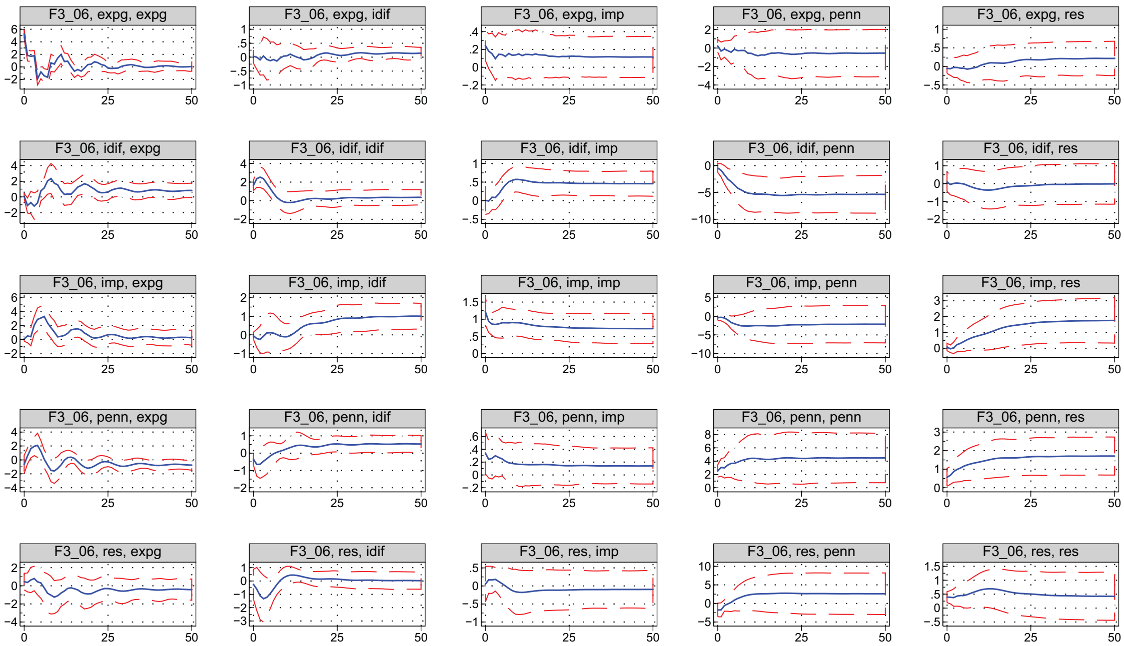
Subtitle: model, shock, response.

Figure 5.1.2 Forecast Error Variance Decomposition of VEC Model F2_07 (Full Sample)



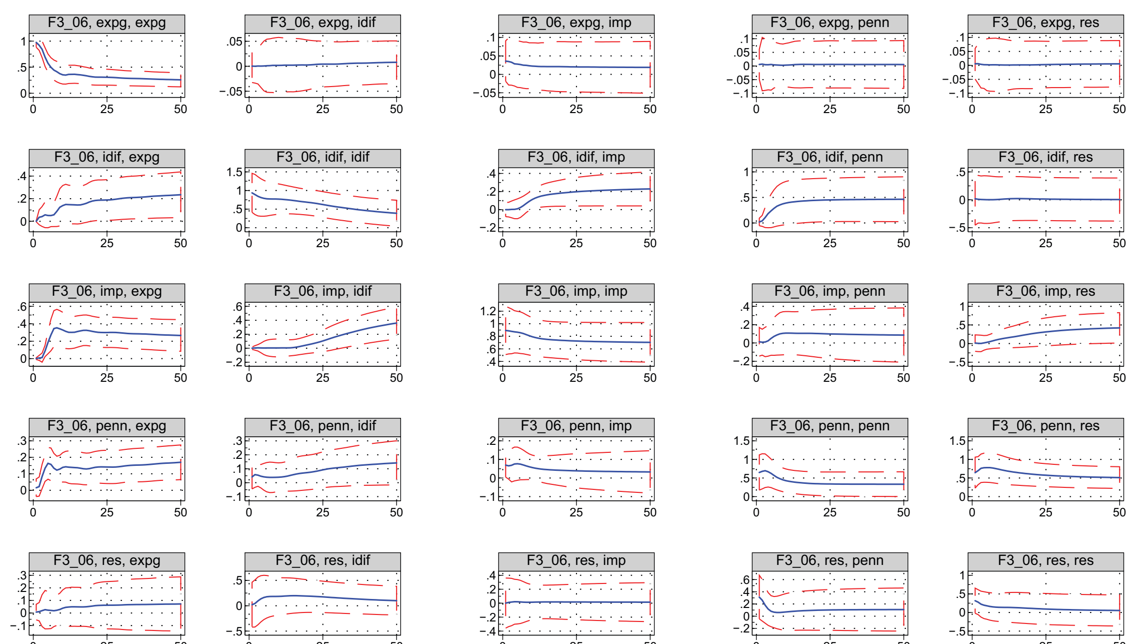
Subtitle: model, shock, response.

Figure 5.2.1 Structural IRF of VEC Model F3_06 (Full Sample)



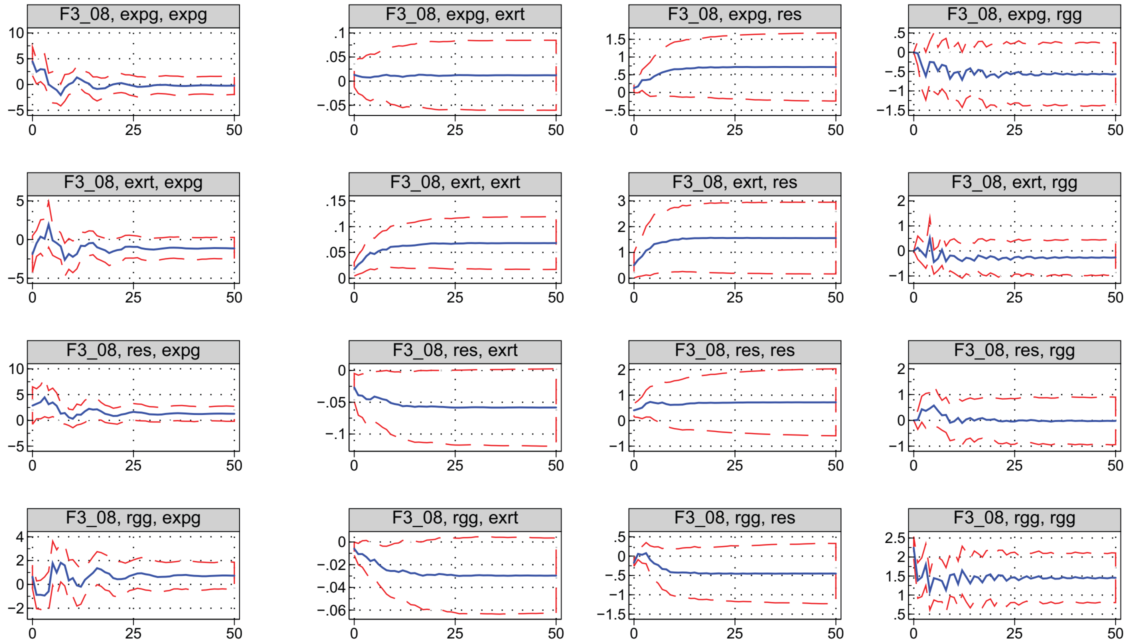
Subtitle: model, shock, response.

Figure 5.2.2 Forecast Error Variance Decomposition of VEC Model F3_06 (Full Sample)



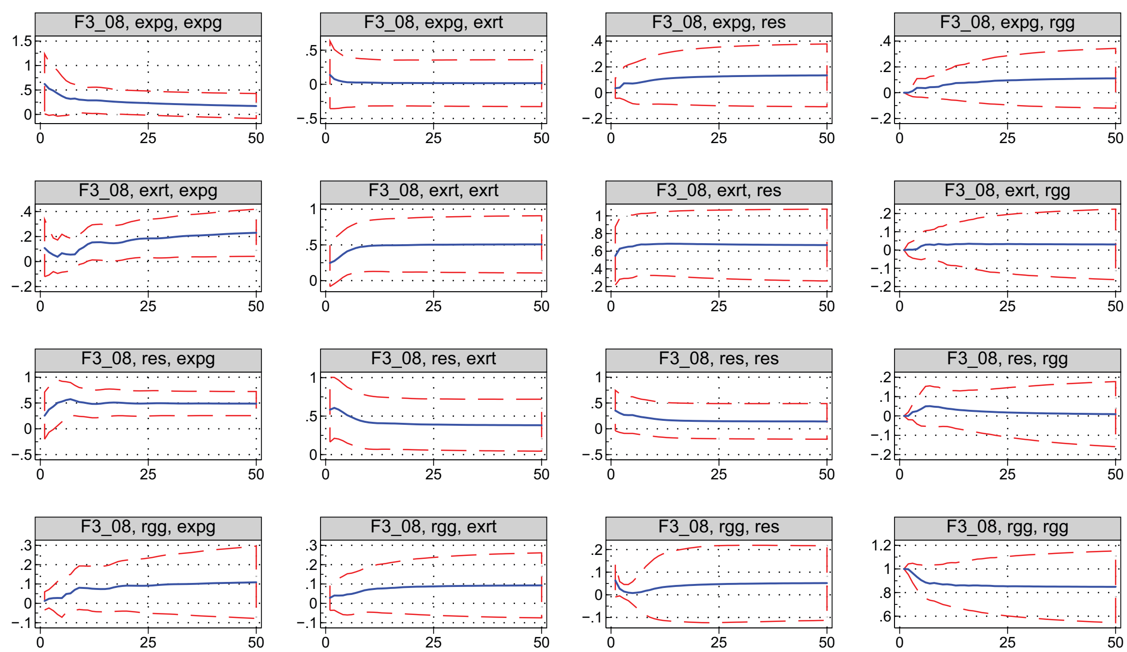
Subtitle: model, shock, response.

Figure 5.3.1 Structural IRF of VEC Model F3_08 (Full Sample)



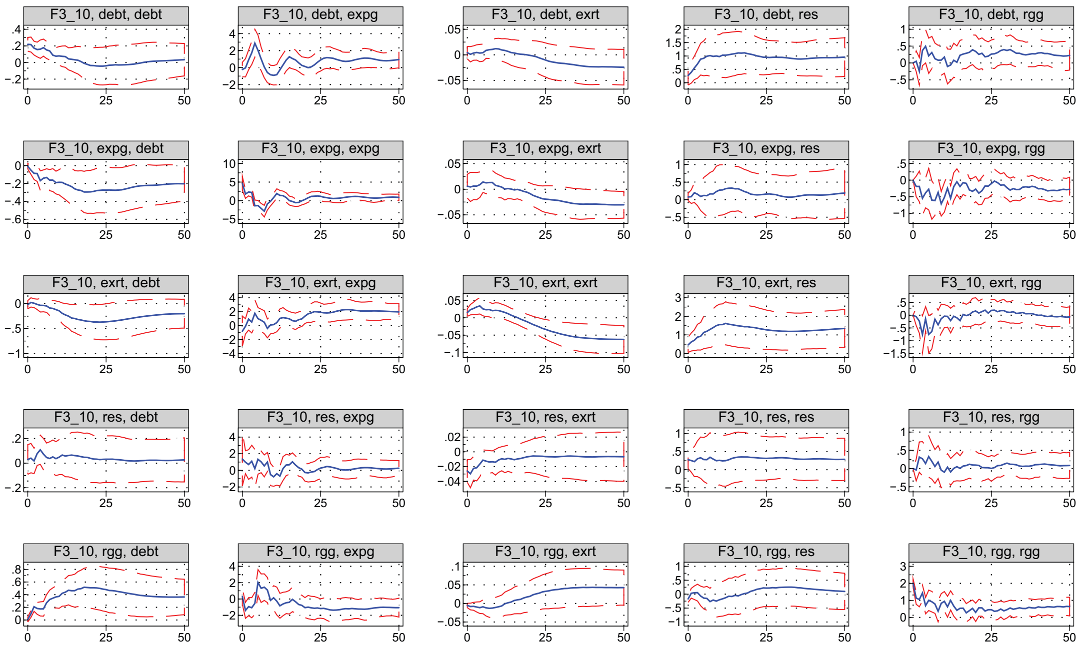
Subtitle: model, shock, response.

Figure 5.3.2 Forecast Error Variance Decomposition of VEC Model F3_08 (Full Sample)



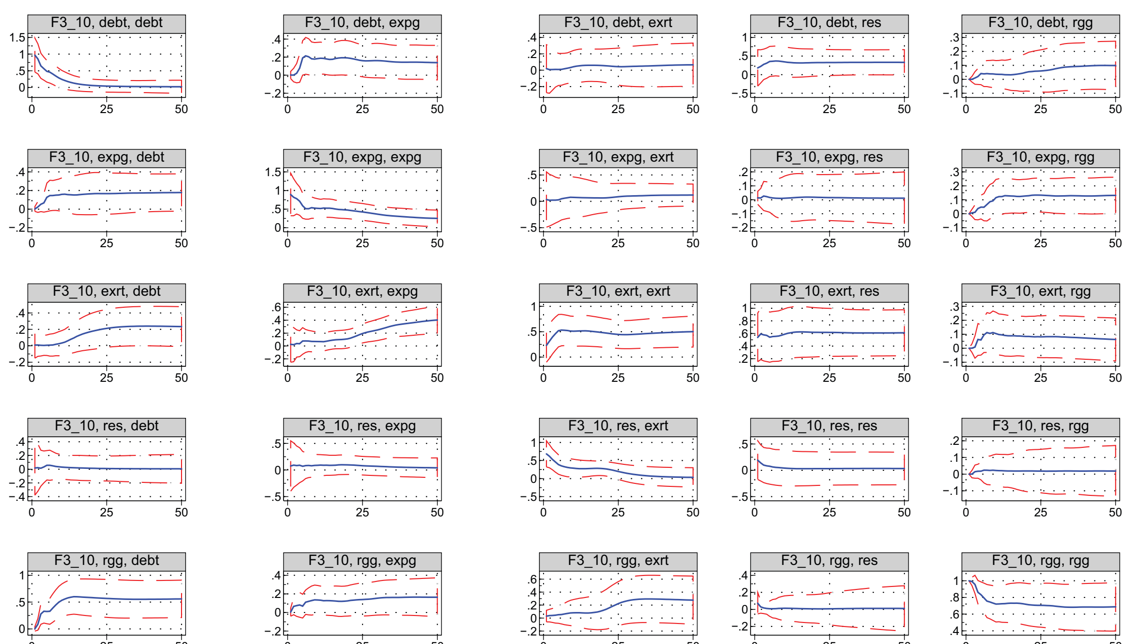
Subtitle: model, shock, response.

Figure 5.4.1 Structural IRF of VEC Model F3_10 (Full Sample)



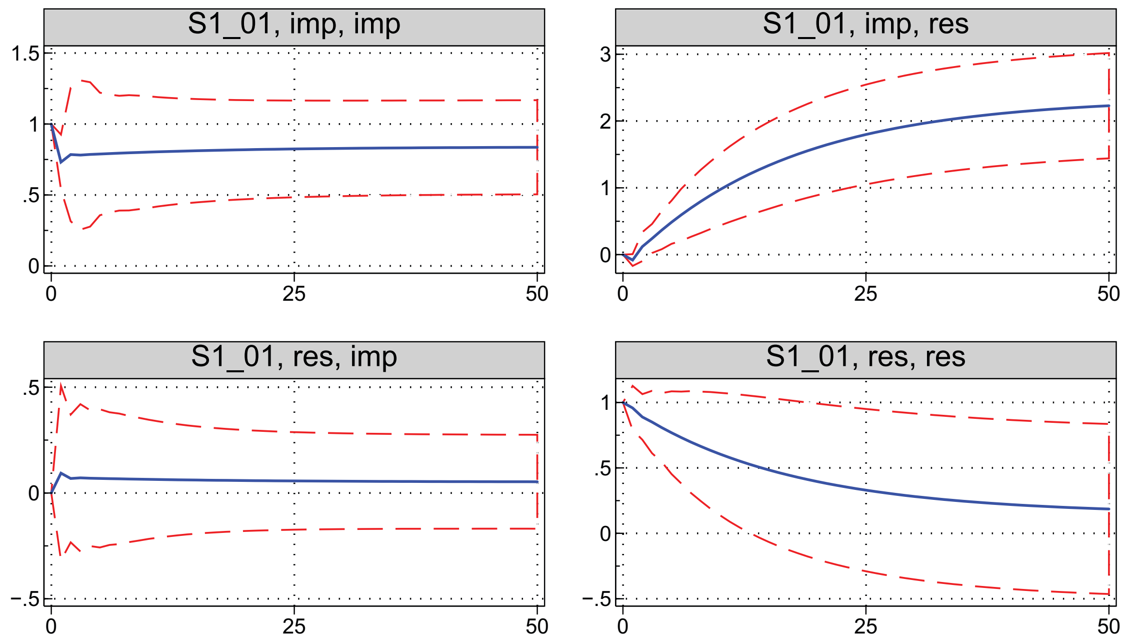
Subtitle: model, shock, response.

Figure 5.4.2 Forecast Error Variance Decomposition of VEC Model F3_10 (Full Sample)



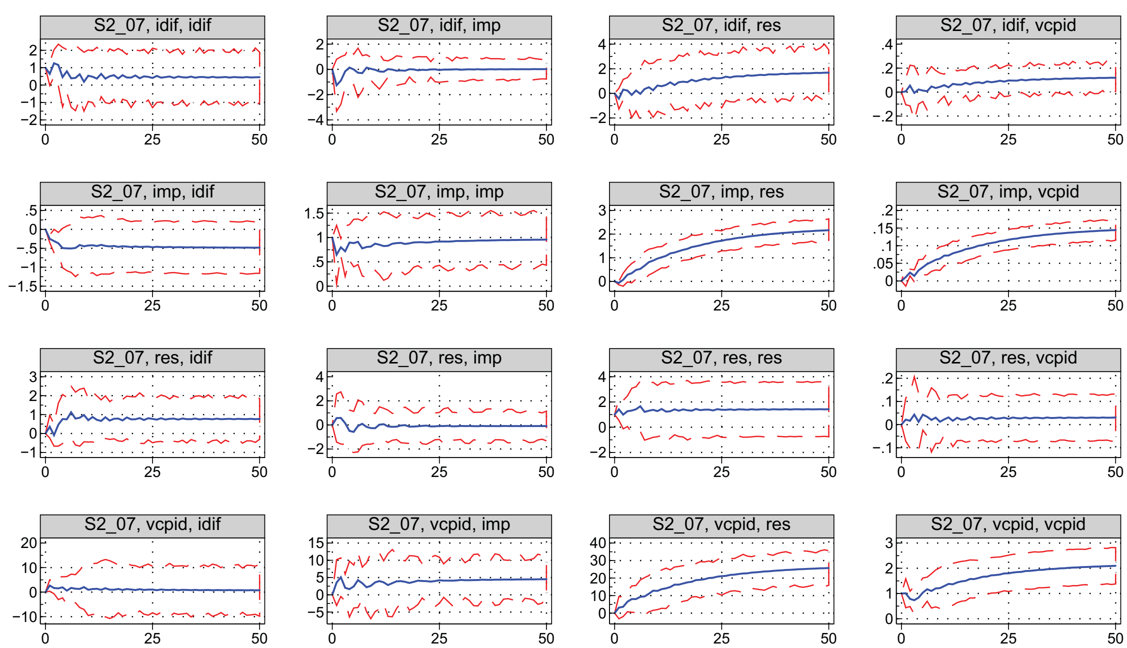
Subtitle: model, shock, response.

Figure 6.1 IRF of VEC Model S1_01 (Subsample)



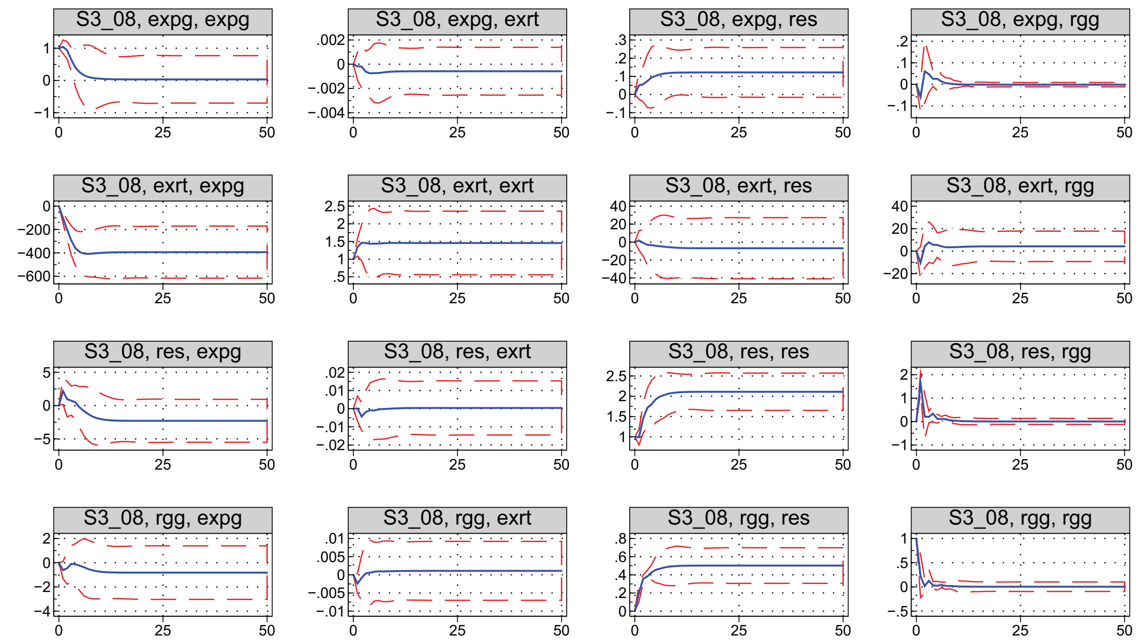
Subtitle: model, shock, response. Std(res)=7.05 Std(imp)=5.66

Figure 6.2 IRF of VEC Model S2_07 (Subsample)



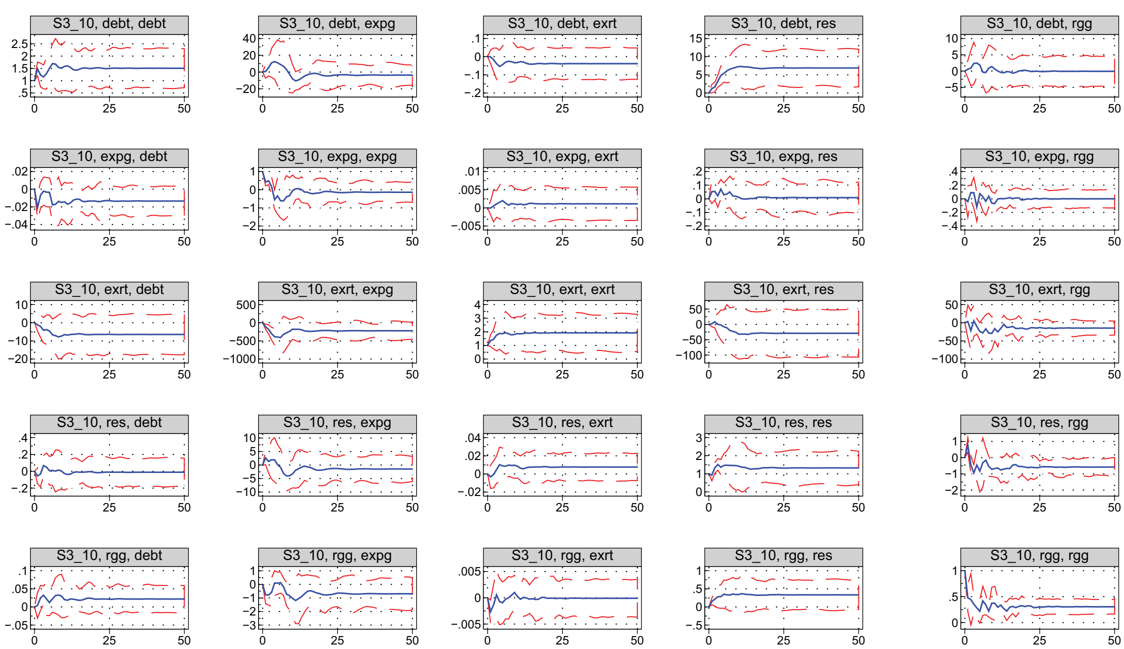
Subtitle: model, shock, response. Std(res)=7.05 Std(imp)=5.66 Std(idif)=1.61 Std(vcpid)=.34

Figure 6.3 IRF of VEC Model S3_08 (Subsample)



Subtitle: model, shock, response. Std(res)=7.05 Std(rgg)=2.28 Std(exrt)=.05 Std(expg)=14.36

Figure 6.4 IRF of VEC Model S3_10 (Subsample)



Subtitle: model, shock, response. Std(res)=7.05 Std(rgg)=2.28 Std(debt)=1.48 Std(exrt)=.05 Std(expg)=14.36

Figure 7.1 Prediction Errors by VEC Model (S1_01)

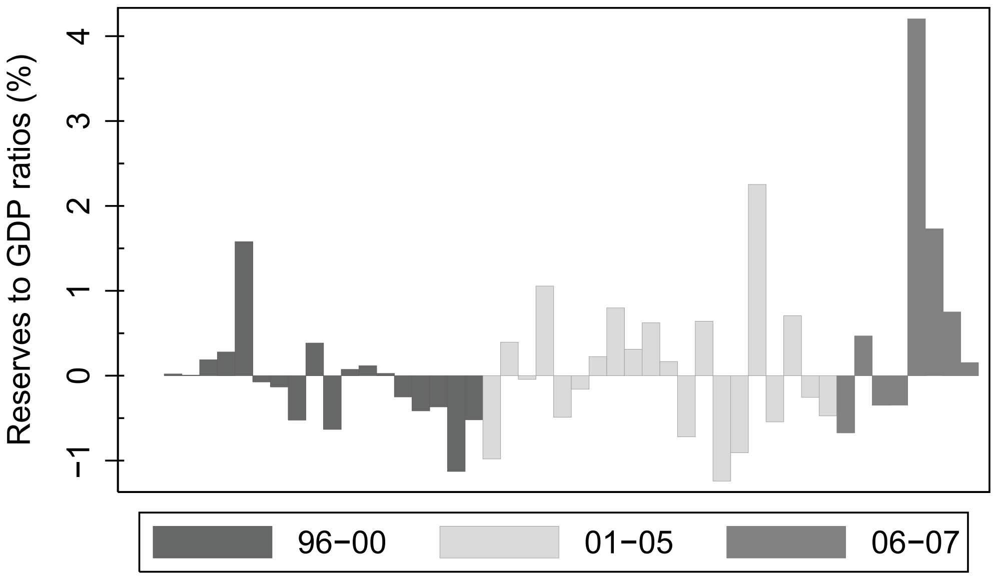


Figure 7.2 Prediction Errors by VEC Model (S3_10)

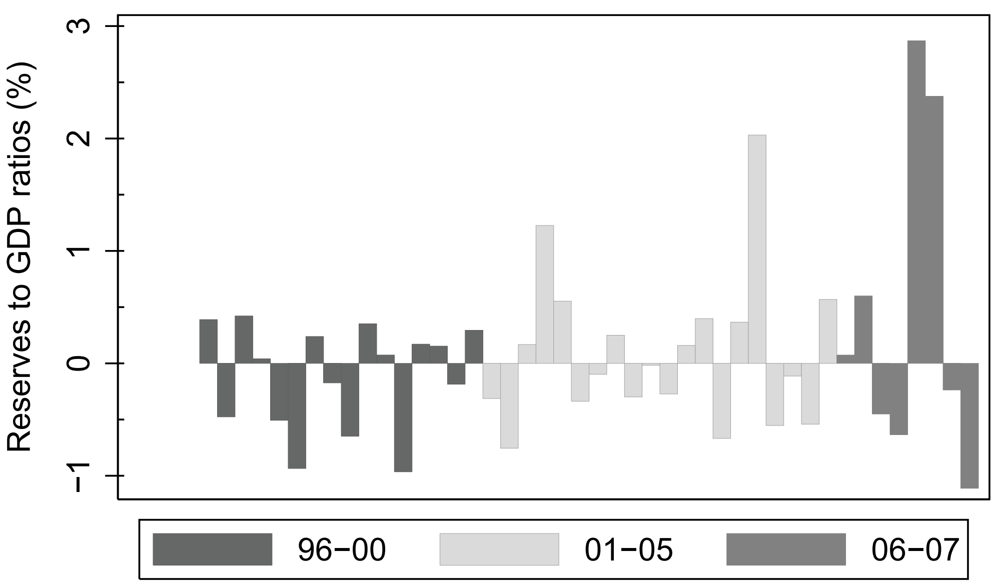
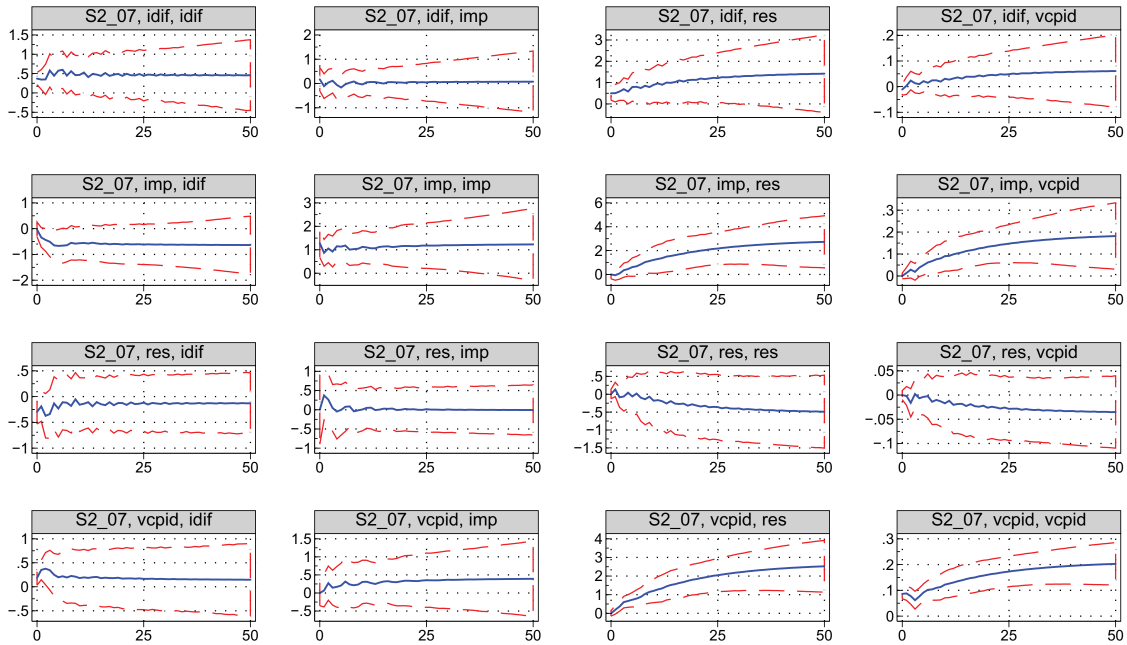
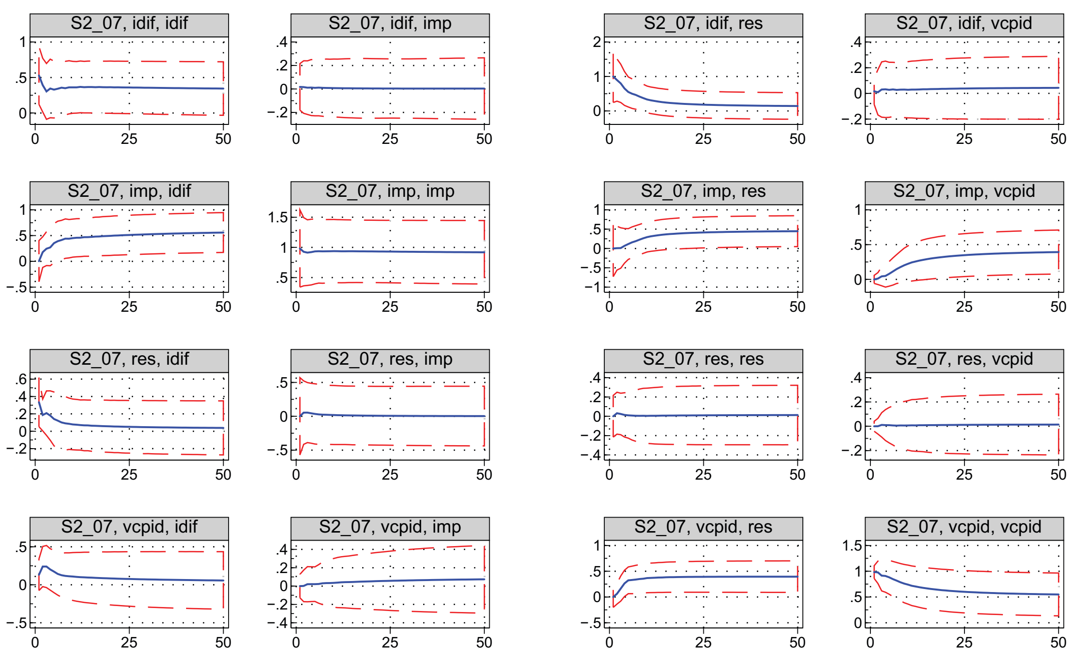


Figure 8.1.1 Structural IRF of VEC Model S2_07 (Subsample)



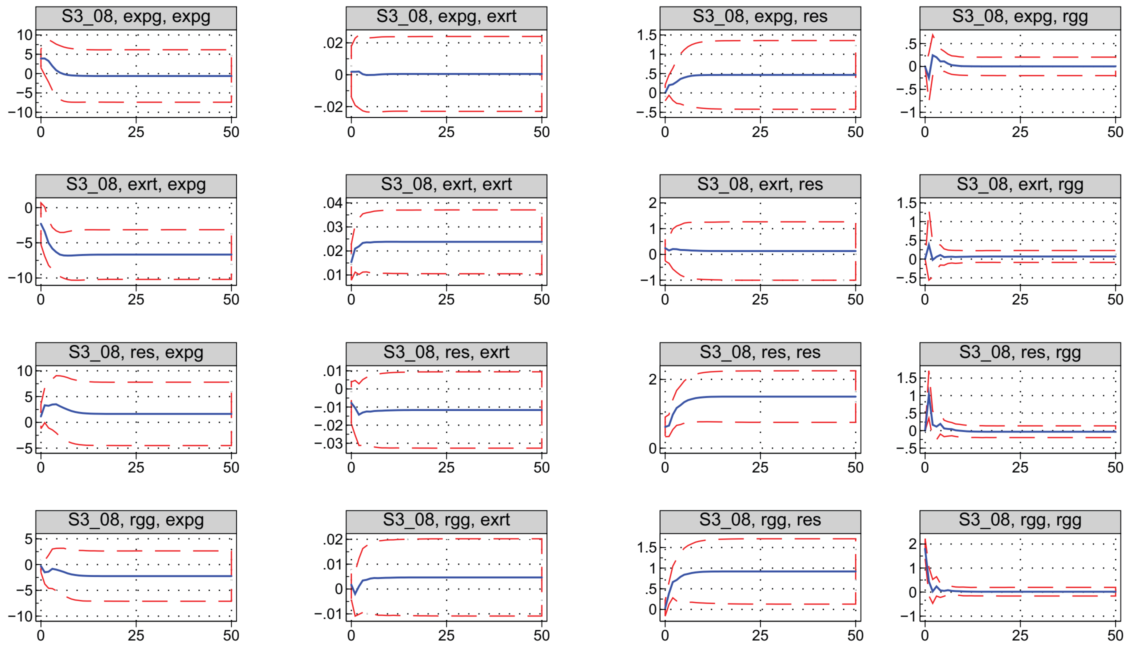
Subtitle: model, shock, response.

Figure 8.1.2 Forecast Error Variance Decomposition of VEC Model S2_07 (Subsample)



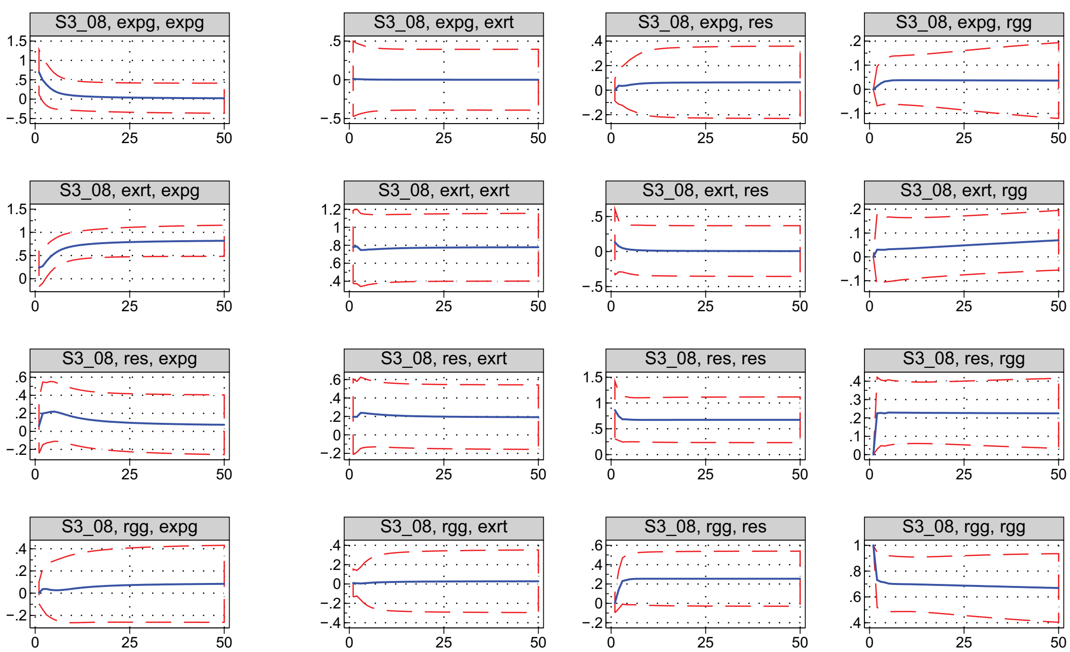
Subtitle: model, shock, response.

Figure 8.2.1 Structural IRF of VEC Model S3_08 (Subsample)



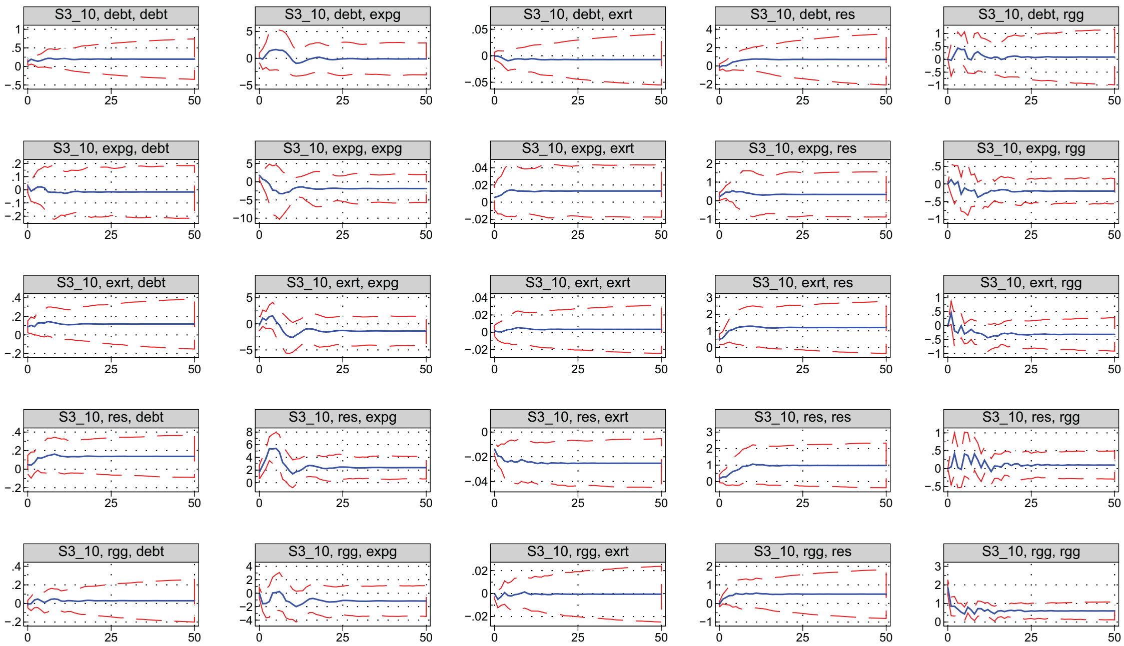
Subtitle: model, shock, response.

Figure 8.2.2 Forecast Error Variance Decomposition of VEC Model S3_08 (Subsample)



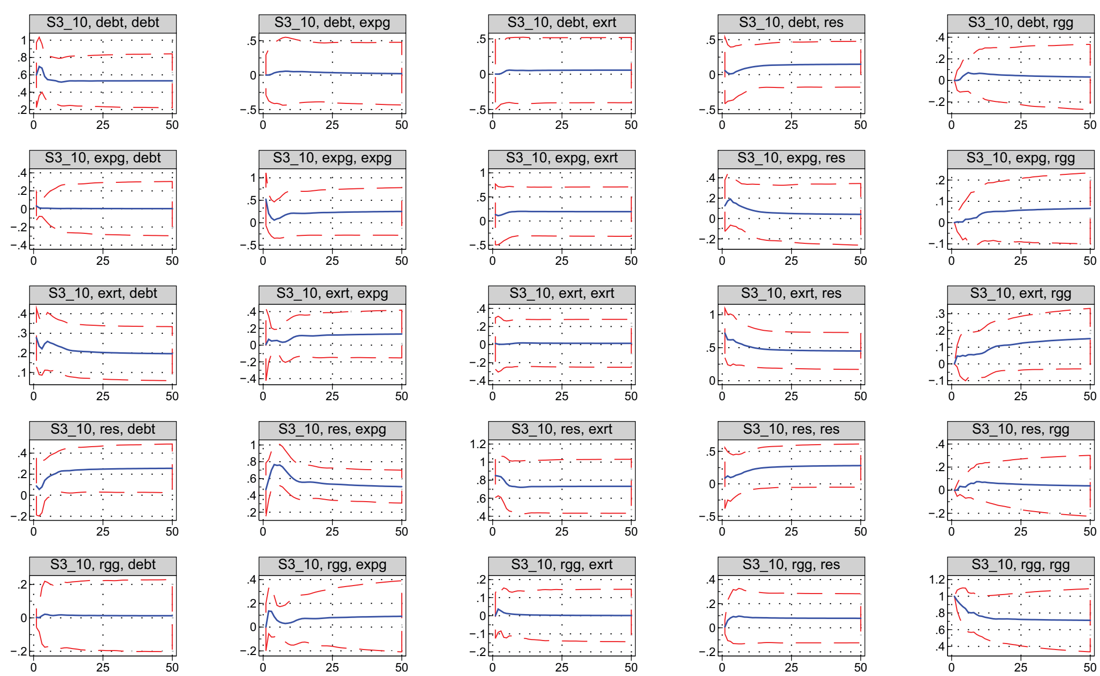
Subtitle: model, shock, response.

Figure 8.3.1 Structural IRF of VEC Model S3_10 (Subsample)



Subtitle: model, shock, response.

Figure 8.3.2 Forecast Error Variance Decomposition of VEC Model S3_10 (Subsample)



Subtitle: model, shock, response.

Appendix

Table A1 Data Sources and Regressors

Data sources:	
[1]	International Financial Statistics, IMF.
[2]	Monthly Bulletin of Statistics-China, July 1985 -, China Statistics Press.
[3]	China Quarterly Gross Domestic Product Estimates 1992-2001, China Statistics Press.
[4]	"Quarterly Real GDP Estimates for China and ASEAN4 with a Forecast Evaluation," Tilak Abeysinghe and Gulasekaran Rajaguru, Department of Economics, Working Paper No. 0404, National University of Singapore, 2003.
[5]	Table 1.7.4. Price Indexes for GDP, GNP and NNP, BEA.
[6]	Federal Reserve Statistical Release H.15, Board of Governers of the Federal Reserve System.

Regressors:	
debt	Total external debts to GDP ratio, %. Total external debts are deflated by the US CPI and GDP are deflated by the Chinese CPI. Sources: [1] [2] [3] [4] [5].
expg	Growth rate of real exports over the same period of last year, %. Nominal exports are deflated by the US CPI. Sources: [1] [5].
exrt	Real effective exchange rate, in logs. Source: [1].
idif	Real interest rate differential between China and the US, percent per annum. For China it is the saving deposit rate and for US it is the 6-month CD rate, both deflated by CPIs. Sources: [1] [2] [5] [6].
imp	Imports to GDP ratio, %. Imports are deflated by the US CPI and GDP are deflated by the Chinese CPI. Sources: [1] [2] [3] [4] [5].
m2gdp	M2 to GDP ratio. Sources: [1] [3] [4].
penn	Deviation of real effective exchange rate from the benchmark value based on the Balassa-Samuelson relation, %. Sources: [1] [2] [3] [4] [5]. Please refer to Table A3 for details of the estimation.
res	Reserves to GDP ratio, %. Reserves are deflated by the US CPI and GDP are deflated by the Chinese CPI. Sources: [1] [2] [3] [4] [5].
rgg	Real GDP growth rate of China relative to the rest of the world, %. The world GDP growth rate is the GDP-weighted average of the high incomes countries and China's top 30 trade partners. Sources: [1] [2] [3] [4] [5].
vcpid	Volatility of inflation rate. It is the standard deviation of detrended changes in inflation rate over the past three years (twelve quarters). Sources: [1] [2].
vexp	Volatility of exports to GDP ratio. It is the standard deviation of detrended changes in exports to GDP ratio over the past three years (twelve quarters). Exports are deflated by the US CPI and GDP are deflated by the Chinese CPI. Sources: [1] [2] [3] [4] [5].

Data are seasonally adjusted.

Table A2 Data Summary Statistics

Full Sample (1983Q1 - 2005Q4)

Obs = 92	Mean	Std.	Min	Max
debt	3.64	2.08	0.602	7.88
expg	13.9	13.1	-15.4	53.2
exrt	4.67	0.303	4.22	5.5
idif	-2.7	6.05	-22.7	5.46
imp	16.4	5.43	6.4	29.8
m2gdp	0.934	0.379	0.378	1.58
penn	1.4	30	-42.1	83.6
res	11.1	8.57	2.89	36.5
rgg	7.42	3.52	-2.26	19.2
vcpid	0.867	0.343	0.302	1.67
vexp	0.928	0.483	0.358	2.25

Obs = 92	*debt*	*expg*	*exrt*	*idif*	*imp*	*m2gdp*	*penn*	*res*	*rgg*	*vcpid*	*vres*
debt	1										
expg	0.13	1									
exrt	-0.63	-0.34	1								
idif	-0.04	-0.12	0.02	1							
imp	0.32	0.56	-0.56	0.12	1						
m2gdp	0.47	0.35	-0.52	0.41	0.81	1					
penn	-0.62	-0.36	1.00	0.00	-0.61	-0.57	1				
res	0.34	0.38	-0.31	0.38	0.83	0.89	-0.37	1			
rgg	0.13	0.16	-0.12	-0.17	0.13	-0.02	-0.12	-0.02	1		
vcpid	0.48	-0.03	-0.43	-0.30	-0.09	-0.23	-0.39	-0.19	0.01	1	
vexp	0.82	0.11	-0.61	-0.22	0.19	0.15	-0.58	0.07	0.20	0.64	1

Subample (1996Q1 - 2005Q4)

Obs =40	Mean	Std.	Min	Max
debt	4.71	1.48	3.04	7.46
expg	15.6	14.4	-13	36.8
exrt	4.59	0.05	4.5	4.68
idif	0.571	1.61	-3.27	3.28
imp	19.8	5.66	13.3	29.8
m2gdp	1.31	0.197	0.926	1.58
penn	-7.68	6.41	-21	3.1
res	19	7.05	11.7	36.5
rgg	6.88	2.28	0.785	13.6
vcpid	0.75	0.339	0.302	1.44
vexp	0.957	0.411	0.533	2.08

Obs = 40	*debt*	*expg*	*exrt*	*idif*	*imp*	*m2gdp*	*penn*	*res*	*rgg*	*vcpid*	*vres*
debt	1										
expg	-0.54	1									
exrt	0.11	-0.44	1								
idif	-0.26	-0.08	0.53	1							
imp	-0.71	0.69	-0.66	-0.10	1						
m2gdp	-0.96	0.64	-0.22	0.27	0.81	1					
penn	0.39	-0.58	0.94	0.40	-0.85	-0.49	1				
res	-0.58	0.58	-0.58	-0.02	0.89	0.69	-0.79	1			
rgg	-0.16	0.17	-0.34	-0.02	0.33	0.15	-0.38	0.28	1		
vcpid	0.80	-0.30	-0.17	-0.23	-0.33	-0.71	0.03	-0.08	0.04	1	
vexp	0.74	-0.41	-0.24	-0.51	-0.38	-0.77	0.02	-0.36	0.02	0.62	1

Table A3 Estimation of the Balassa-Samuel Relation

Regression: log(REER)_it = u_i + b * log(GDPPC_real)_it + e_it

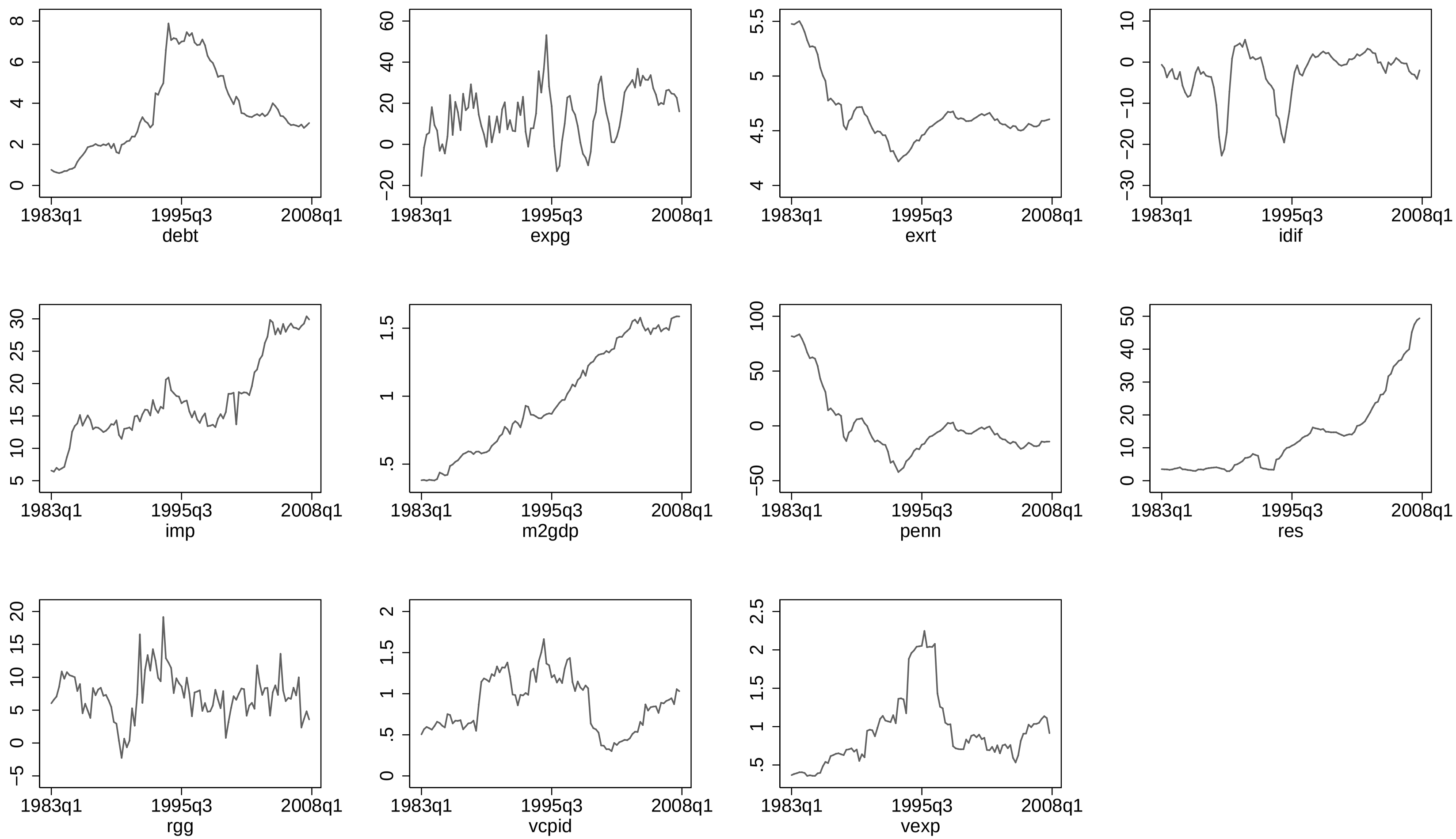
```
Fixed-effects (within) regression               Number of obs      =      3294
Group variable (i): ifscode                     Number of groups   =        44

R-sq:  within  = 0.0872                         Obs per group: min =        30
       between = 0.0397                                        avg =      74.9
       overall = 0.0100                                        max =       100

                                                F(1,3249)          =    310.52
corr(u_i, Xb)  = -0.7713                        Prob > F           =    0.0000

------------------------------------------------------------------------------
       log(REER)|      Coef.   Std. Err.      t    P>|t|     [95% Conf. Interval]
----------------+-------------------------------------------------------------
log(GDPPC_real) |   .1023224   .0058066    17.62   0.000     .0909373    .1137074
       constant |   4.773169   .0076999   619.90   0.000     4.758072    4.788266
----------------+-------------------------------------------------------------
        sigma_u |  .13001741
        sigma_e |  .11058102
            rho |  .58025974   (fraction of variance due to u_i)
------------------------------------------------------------------------------
F test that all u_i=0:        F(43, 3249) =    48.94           Prob > F = 0.0000
```

Table A4 Structural Estimation Results

	a1	a2	a3	a4	a5	a6	a7	a8	a9	a10	σ_1*	σ_2	σ_3	σ_4	σ_5
F2_07	-0.14	-0.5	-2.52	-0.37	2.92	0.05					1.12	1.26	2.77	0.11	
F3_06	-0.14	-0.12	-0.21	0.03	-0.36	-0.07	-0.05	0.61	4.45	0.26	0.80	1.22	1.70	5.02	5.34
F3_08	-24.24	0.04	0.11	-0.01	0.01	106.86					1.18	0.08	5.89	2.24	
F3_10	-30.62	0.02	-0.61	0.13	-0.004	-0.18	0.02	52.52	0.01	1.20	1.08	0.08	4.97	0.23	2.01
S2_07	-0.03	-1.25	2.73	-0.34	51.06	0.02					0.37	1.29	25.87	0.09	
S3_08	-14.89	0.01	0.01	-0.0005	-0.001	149.45					0.74	0.02	4.20	1.80	
S3_10	-211.52	0.55	-0.43	9.85	-1.34	11.73	0.32	121.57	-0.18	1.29	3.99	6.13	2.44	0.15	1.86

* σ's are the standard deviations of the structural shocks in *v* .

Figure A1   Graphs of Data Series (1983Q1−2007Q4)



CHAPTER TWO

The Impact of Trade Liberalization on the Trade Balance in Developing Countries

Jiandong Ju, Yi Wu and Li Zeng[1]

Using two recently constructed measures of trade liberalization dates, this paper studies the impact of trade liberalization on imports, exports, and overall trade balance for a large sample of developing countries. We find strong and consistent evidence that trade liberalization leads to higher imports and exports. However, in contrast to Santos-Paulino and Thirwall (2004), who find a robustly negative impact of trade liberalization on the overall trade balance, we find only mixed evidence of such a negative impact. In particular, we find little evidence of a statistically significant negative impact using our first measure of liberalization dates, which extends Li (2004). Using a second measure of liberalization dates compiled by Wacziarg and Welch (2003), we find some evidence that liberalization worsens the trade balance, but the evidence is not robust across different estimation specifications, and the estimated impact is smaller than that reported by Santos-Paulino and Thirwall (2004). [JEL F11, F14]

1. Introduction

Many developing countries have substantially liberalized their trade regime over the past three decades, either unilaterally or as part of multilateral initiatives. Nevertheless, trade barriers remain high in many developing countries. One of the concerns that attributes to the reluctance of many of these countries to liberalize their trade regime is the possible worsening of the trade balance.[2] This is the question we want to investigate in this study: did past liberalization episodes in developing countries lead to a deterioration of their trade balance?

On the theoretical ground, Ostry and Rose (1992) offer an extensive survey of the macroeconomic effects of trade tariffs based on different theoretical frameworks, including the income-expenditure approach, the monetary approach, and the intertemporal approach. The authors conclude that there is no clear conclusion about the effect of a tariff change on the trade balance. The effect depends on the behavior of real wages and exchanges rates, on

[1] Jiandong Ju is an associate professor in the Department of Economics at the University of Oklahoma. Yi Wu is an economist with the IMF Western Hemisphere Department. Li Zeng is an economist with the IMF Asia and Pacific Department. The authors would like to thank Maria Oliva Armengol, Beatriz A Maldonado-Bird, Azim Sadikov, Dustin Smith, and especially Tom Dorsey, Hans Peter Lankes, Brad McDonald, the editor, an anonymous referee for very helpful comments, and Xiangming Li for sharing her data set with us.

[2] Another common concern is the decline in tariff revenue—often a major source of revenue for developing countries. Baunsgaard and Keen (2005) find that low-income countries have mostly not been able to offset reductions in trade tax revenues by increasing their domestic tax revenues.

the values of a variety of elasticities, the degree of capital mobility, and whether the tariff shock is perceived as temporary or permanent.

Using a simple two period intertemporal trade model, we analyze the effect of trade liberalization on the import, the export, and the trade balance in a small country. The effects rely on the interactions among the *real income effect*, the *intratemporal substitution effect* between importing goods and exporting goods, and the *intertemporal substitution effect* across time periods. The *intertemporal substitution effect* is negligible, as the tariff reductions are permanent, and the small country takes the world prices and the interest rate as exogenous. Tariff reductions increase the real income and decrease the price of the import good. Thus, both the *real income effect* and the *intratemporal substitution effect* increase imports. The *intratemporal substitution effect* decreases the domestic consumption of the exportable good, while the *real income effect* increases it. Assuming the former effect dominates, trade liberalization will increase exports. As trade liberalization increases both exports and imports, the difference of these two, the trade balance, may increase or decrease due to tariff reductions. The impact of trade liberalization on the trade balance, therefore, needs to be investigated empirically.

One stream of the related empirical literature attempts to find out how trade liberalization affects a country's imports, and generally finds a positive impact (Melo and Vogt, 1984; Bertola and Faini, 1991; and Santos-Paulino, 2002a). There are also empirical researches focusing on the effects of trade liberalization on exports, where the findings are more mixed. Some of them show that countries which embarked on liberalization programs have improved their export performance (Ahmed, 2000; Thomas, Nash, and Edwards, 1991; and Santos-Paulino, 2002b) while others have found little evidence of such a relationship (Greenaway and Sapsford, 1994; Jenkins 1996).

For policy makers, the impact of trade liberalization on the overall balance would be the more important question. There have been however surprisingly few cross-country empirical studies on the subject. Ostry and Rose (1992) studied the impact of tariff changes on the trade balance using five different data sets, mostly data from Organization for Economic Cooperation and Development countries, and found no statistically significant effect. UNCTAD (1999) studied the effect of trade liberalization on the trade balance for 15 developing countries over the period of 1970 to 1995, and found a significant negative relationship. In a more recent paper, Santos-Paulino and Thirlwall (2004) studied the effect of trade liberalization on imports, exports and the overall trade balance using a sample of 22 developing countries for the period of 1972–97. They found that liberalization stimulated export growth but raised import growth by more, leading to a worsening of the overall trade balance.

One constraint researchers on the subject often face is the lack of systematic data measuring the dates of trade liberalization. Indeed, due to data limitation, most of the empirical studies on the subject are constrained to country case studies. In this paper, we use two recently compiled data sets establishing trade liberalization dates that cover a large sample of developing countries for a long period of time. In particular, our two samples cover 39 and 77 developing countries for the period of 1970–2004, and 1970–2001, respectively. Our study focuses on the impact of trade liberalization for developing countries, for which the policy relevance of this question remains especially high. We find strong evidence that trade liberalization leads to faster import and export growth. The evidence on the overall trade balance, however, is mixed. Using our first measure of trade liberalization dates, we find little evidence that trade liberalization worsens the trade balance. There is some evidence that liberalization leads to a deterioration of the trade balance when we use our second measure of liberalization dates, although the finding is not robust to alternative estimation specifications.

2. Theoretical Analysis

This section develops a two-period intertemporal trade model to analyze the effect of trade liberalization on the import, the export, and the trade balance. Since our empirical analysis investigates permanent tariff reductions in developing countries, we study a permanent tariff reduction in a small country in this theoretical analysis. The key insight relies on the interactions among the *real income effect*, the *intratemporal substitution effect* between importing goods and exporting goods, and the *intertemporal substitution effect* across time periods.

The life time utility function for the representative consumer in the home country is defined as

$$U = u(C_1) + \beta u(C_2),\ 0 < \beta < 1, \tag{1}$$

where C_t is the consumption in period t (t = 1,2) and β is a time-preference factor. With respect to technology, the home country specializes in producing a single good, labeled as good 1, and the foreign country specializes in producing good 2. Let the output of good 1 in period t be y_{t1}. The linear homogeneous production function for good 1 is $y_{t1} = F(k_t, L)$. where k_t and L are capital and labor used in production, respectively. We assume that labor supply, L, is fixed, and is normalized as 1 from here on. The capital stocks evolve according to

$$k_{t+1} = k_t + I_t, \tag{2}$$

where I_t is the investment in period t and the depreciation rate is assumed to be zero.

Consumption and investment are composite of foreign and domestic goods:

$$C_t + I_t = G(x_{t1}, x_{t2}),$$

where $G(x_{t1}, x_{t2}) = (x_{t1}^{\rho} + x_{t2}^{\rho})^{1/\rho}$ is an Armington aggregator and $0 < \rho \leq 1$. The elasticity of intratemporal substitution between foreign and domestic goods is $\sigma = 1/(1-\rho)$ and $1 < \sigma < \infty$. This setup is standard in the literature of international real business cycle (IRBC) (see Backus, Kehoe, and Kydland, 1992 and 1994 for more discussions). While the IRBC literature uses an infinite horizon model for calibrations, we use a two period model to get closed form solutions, in order to provide some intuition for our empirical investigations.

Similar to the argument in Ostry (1988), the consumer may be viewed as solving a two-stage optimization problem. In the first stage, the consumer chooses x_{t1} and x_{t2} to minimize her expenditure for a given level of consumption and investment, $C_t + I_t$. That is, she solves

$$\min_{x_{t1}, x_{t2}} E = p_{t1} x_{t1} + p_{t2} x_{t2} \text{ subject to}$$
$$G(x_{t1}, x_{t2}) \geq C_t + I_t$$

where p_{ti} is the domestic price of good *i*. Letting τ be an ad valorem tariff rate on imports, we have $p_{t1} = p_{t1}^*$ and $p_{t2} = (1+\tau) p_{t2}^*$ where p_{ti}^* is the world price. The solution to this problem yields the expenditure function

$$E(p_{t1}, p_{t2}, C_t + I_t) = q_t (C_t + I_t)$$

where $q_t = (p_{t1}^{1-\sigma} + p_{t2}^{1-\sigma})^{\frac{1}{1-\sigma}}$. To simplify the analysis, we assume that the world prices do not change. That is, $p_{1j}^* = p_{2j}^*$ for $j = 1,2$, and therefore we have $q_1 = q_2 = q$. Using the envelope theorem, we have

$$x_{ti} = \frac{\partial E(\cdot)}{\partial p_{ti}} = q_t^{\sigma} p_{ti}^{-\sigma} (C_t + I_t)$$

The intertemporal budget constraint for the consumer can be written as

$$q_1 (C_1 + I_1) + \frac{q_2 (C_2 + I_2)}{1+r} = p_{11} F(k_1) + \tau p_{12}^* x_{12} + \frac{p_{21} F(k_2) + \tau p_{22}^* x_{22}}{1+r}, \tag{3}$$

where r is the world interest rate which the small country takes as exogenous. The government redistributes the tariff revenue, $\tau p_{t2}^* x_{t2}$, back to the consumer in every period. Note that capital, k_2, accumulated in period 1 will be consumed at the end of period 2 and k_3 will be zero, implying that $I_2 = k_3 - k_2 = -k_2$. In the second stage, the consumer chooses C_1, I_1, and C_2 to maximize lifetime utility (1) subject to the intertemporal budget constraint (3) (k_1 is given by history and is not subject to choice on date 1).

Using equation (3) to substitute C_2 in (1), the two first order conditions for C_1 and I_1 are

$$\frac{\partial u(C_1)}{\partial C_1} = \beta\left(1+r\right)\frac{\partial u(C_2)}{\partial C_2} \ , \tag{4}$$

and

$$\frac{p_{21}}{Q}\frac{\partial F(k_2)}{\partial k_2} = r \ , \tag{5}$$

where $Q = q_2 - \tau p_{22}^{*} q_2^{\sigma} p_{22}^{-\sigma}$ is the aggregate price index, excluding the tariff revenue effect. Equation (4) is the standard Euler equation, and equation (5) states that the marginal value product of capital equals the interest rate. C_1, I_1, and C_2 are solved by equations (3), (4), and (5). The import value M_t, the export value X_t, and the trade balance TB_t are correspondingly written as

$$\begin{aligned} M_t &= p_{t2}^{*} x_{t2} = p_{t2}^{*} q_t^{\sigma} p_{t2}^{-\sigma}(C_t + I_t) \\ X_t &= p_{t1}^{*}(y_t - x_{t1}) = p_{t1}^{*}[y_t - q_t^{\sigma} p_{t1}^{-\sigma}(C_t + I_t)] \\ TB_t &= X_t - M_t \end{aligned}$$

Note that the intertemporal budget constraint (3) implies that $TB_1 + \frac{TB_2}{1+r} = 0$.

We are now ready to discuss the effect of trade liberalization. With some computations, we can show that $\frac{\partial Q}{\partial \tau} > 0$. Hence, the aggregate price index declines as tariff rate τ decreases. Equation (5) then indicates that k_2, and therefore I_1 must increase, since now the real price of the domestic product, $\frac{p_{21}}{Q}$, becomes higher. Rewriting the intertemporal budget constraint (3), we have

$$Q\left(C_1 + I_1\right) + \frac{Q\left(C_2 - k_2\right)}{1+r} = p_{11}F(k_1) + \frac{p_{21}F(k_2)}{1+r} \ . \tag{6}$$

The value of the right hand side of equation (6) increases as k_2 increases. Therefore, $C_1 + I_1$ must increase. The proof is straightforward: if $C_1 + I_1$ were smaller, then C_1 would be smaller since I_1 is larger, then C_2 would be smaller using equation (4), so that the value of the left hand side of equation (6) would decline, and that would be a contradiction.

When tariff rate τ is reduced, the real price of the domestic good and therefore the real income increases. This is labeled as the *real income effect,* which increases both consumption demand and investment demand. The *intertemporal substitution effect* across time periods is negligible. Even if C_1 declines, $C_1 + I_1$ must be higher after tariff reductions.

The effect of trade liberalization on the import value in current period is

$$\frac{\partial M_1}{\partial \tau} = p_{12}^*\left(C_1 + I_1\right)\frac{\partial\left(q_1^\sigma p_{12}^{-\sigma}\right)}{\partial \tau} + p_{12}^* q_1^\sigma p_{12}^{-\sigma}\frac{\partial\left(C_1 + I_1\right)}{\partial \tau}. \tag{7}$$

It is easy to show that the first derivative in the right hand side of expression (7) $\frac{\partial\left(q_1^\sigma p_{12}^{-\sigma}\right)}{\partial \tau} < 0$. This is called the *intratemporal substitution effect:* the tariff reduction reduces the price of the import good and therefore increases the import demand. As we have argued above, the *real income effect* implies that $\frac{\partial (C_1 + I_1)}{\partial \tau} < 0$. Thus both the *intratemporal substitution effect* and the *real income effect* increase the value of imports.

Noting that $y_1 = F(k_1)$ does not change, the effect of trade liberalization on the value of exports in current period is

$$\frac{\partial X_1}{\partial \tau} = -\left[p_{11}^*\left(C_1 + I_1\right)\frac{\partial\left(q_1^\sigma p_{11}^{-\sigma}\right)}{\partial \tau} + p_{11}^* q_1^\sigma p_{11}^{-\sigma}\frac{\partial\left(C_1 + I_1\right)}{\partial \tau}\right]. \tag{8}$$

Now the first derivative in the right hand side of (8) $\frac{\partial\left(q_1^\sigma p_{12}^{-\sigma}\right)}{\partial \tau} > 0$. That is, the *intratemporal substitution effect* decreases the domestic consumption of the exportable good and therefore increases the export value, while the *real income effect* does the opposite. Assuming the former effect dominates, we have $\frac{\partial X_1}{\partial \tau} < 0$. Hence, trade liberalization increases the export value.

As both X_1 and M_1 increase, the difference of these two, the trade balance, may increase or decrease due to tariff reductions. More precisely, with some computations we have

$$\frac{\partial TB_1}{\partial \tau} = \sigma q_1^{2\sigma-1} p_{11}^{-\sigma} p_{12}^{-\sigma}\left(-1 + \frac{p_{12}^*}{p_{11}^*\left(1+\tau\right)}\right),$$

and its sign may be positive or negative. Summarizing we have:

Proposition 1: *Tariff reductions increase the value of imports in the current period. If the intratemporal substitution effect dominates the real income effect, tariff reductions increase the value of exports in the current period. The effect of tariff reductions on the trade balance is ambiguous.*

Since $y_1 = F(k_1)$ does not change, our results also hold for the ratio of the import value, the export value, and the trade balance to GDP, which we will use as the dependent variables in our empirical study. As we will show next, the theoretical results we derived above are consistent with our empirical investigations.

3. Two Measures of Trade Liberalization Dates

Our first measure of trade liberalization dates is based on Li (2004), who has individually documented trade liberalization episodes in 45 countries between 1970 and 1995. We extended the liberalization measure for the 39 developing countries[3] in her data set to 2004 using the tariff data from the United Nations Conference on Trade and Development's Trade Analysis and Information System (TRAINS) database [supplemented by data from the IMF's Trade Policy Information Database (TPID)]. In doing so, a trade liberalization episode is identified if there is a continuous and accumulated tariff reduction by at least 35 percent (e.g., a tariff reduction from 15 to 9.75 percent).[4] However, once a country's overall tariff level reaches 10 percent or lower, we regard it as open and a further tariff cut, even by more than 35 percent, will no longer be considered as a liberalization episode.[5] The IMF's TPID database also rates a country's nontariff barrier level into three categories (open, moderate, and restrictive). In addition to looking at tariff reductions, we also take the reductions in nontariff barriers into consideration when defining a liberalization episode. However, it turns out that reductions in nontariff barriers are usually accompanied by large tariff cuts.

Table 1 reports our first measure of liberalization dates covering the period between 1970 and 2004, with the years of liberalization episodes highlighted (tariff reductions typically spread over several years). Two observations are worth mentioning. First, the period of 1985–95 seems to be the "opening-up decade" for developing countries. Almost all the countries in our sample experienced one or more episodes of liberalization during this period. Secondly, many countries experienced multiple episodes of liberalization (this is

[3]According to the World Bank's classification (http://go.worldbank.org/K2CKM78CC0).

[4]Ideally we would like to use the weighted average tariff, but often only the simple average tariff data are available.

[5]One example where this 10percent threshold is applied is Chile. Over the period from 1999–2004, Chile's simple average tariff rate was reduced from ten percent to five percent, which was a cut of 50 percent. However, since the 10 percent threshold was already met at the initial tariff level, this period is not treated as a liberalization episode.

the case for 20 of the 39 countries in the sample). Indeed, trade liberalization is still an ongoing process for many developing countries.

For countries that experienced multiple liberalization episodes, a subsequent liberalization is often implemented either because the earlier one was limited in scope or was later reversed (at least partially). We therefore define a trade liberalization dummy, which takes the value of one after the end of the last recorded liberalization episode for a country and zero beforehand.[6]

Our second measure of trade liberalization dates is from Wacziarg and Welch (2003). Wacziarg and Welch define the liberalization date as the date after which all of the Sachs and Warner (1995) openness criteria are continuously met. In particular, Wacziarg and Welch classify a country as closed if it displays at least one of the following characteristics: (1) average tariff rates of 40 percent of more; (2) nontariff barriers covering 40 percent or more of trade; (3) a black market exchange rate that is depreciated by 20 percent or more relative to the official exchange rate, on average; (4) a state monopoly on major export; and (5) a socialist economic system. However, data limitations often forced them to reply on country case studies of trade policy. One advantage of the Wacziarg-Welch data set is that it covers a substantially larger sample of developing countries. The Wacziarg-Welch liberalization dates are also reported in the last column of Table 1 (only for the overlapping countries).

We note in many cases the identified dates are very close across the two measures. For example, our first measure would identify 1992 as the year that Argentina liberalized its trade regime, compared with 1991 in Wacziarg and Welch (2003). For multiple liberalization episodes identified by our first measure, in several cases the Wacziarg-Welch date is closer to the first episode. For example, our first measure suggests that Chile had two episodes of liberalization, during 1974–79 and 1985–92, respectively. Thus our first liberalization dummy will be one starting from 1993. The Wacziarg-Welch liberalization measure, instead, identifies 1976 as the year after which the economy has been open. This misses the reversal afterwards and the second liberalization during 1985–92.[7] Finally, in a few cases, the identified liberalization dates are quite different across the two measures. For example, Li (2004) identifies a liberalization era lasting from 1985 to 1996 for Indonesia (average nominal tariff more than halved), while Wacziarg and Welch classify

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[6]We made one exception for China. China's (simple average) tariff was reduced from 39.7 percent in 1992 to 16.7 percent in 1997, and then from 15.4 percent in 2001 to 10.7 percent in 2003, and further to 9.8 percent in 2004. This is a 36 percent tariff reduction from 2001 to 2004. The classification will make the liberalization dummy zero for China for our sample period, and the analysis would miss the dramatic opening up and trade promotion that had happened during the 1990s. We therefore assign the liberalization dummy as one for China after 1998. Nevertheless, the regression results would be broadly similar even if we did not make such an exception.

[7]Chile's uniform tariff was raised to 20 percent in 1983, then to 35 percent in 1984. During 1985–92, the uniform tariff rate was reduced to 15 percent, while the average tariff dropped from 36 to 12 percent. Nontariff barriers were also lowered (Li, 2004).

Indonesia as open from 1970. Nevertheless, the two measures are significantly and positively correlated, with a correlation coefficient of 0.57 (for countries in which they overlap).

Table 2a tabulates the average import-, export-, and trade-balance-to-GDP ratios using our first measure of trade liberalization for the periods before and after liberalization. Reported at the bottom of the table are cross-country averages. In general, countries not only import but also export more after they liberalized their trade regimes. The cross-country average import-to-GDP ratio increased from 23.8 to 30.6 percent, with 33 countries seeing their import-to-GDP ratio increased versus four countries experiencing a decline. The average export-to-GDP ratio increased from 19.5 to 24.1 percent, with the ratio increased in 28 countries and reduced in nine countries. The average increase in exports however is smaller than that of imports, as the average trade deficit slightly increased from 4.3 to 6.5 percent. However, the picture is not uniform across countries: 22 countries experienced a deterioration of the trade balance after liberalization, and 15 countries actually had an improved trade balance.

Table 2b reports the summary statistics using the Wacziarg-Welch measure of trade liberalization dates.[8] The average import-to-GDP ratio increased from 25.1 percent before liberalization to 29.9 percent afterwards. 47 of the 62 developing countries that experienced trade liberalization during the period had higher import-to-GDP ratios. The average export-to-GDP ratio increased from 18.5 to 20.4 percent, with 40 countries experiencing an increase in the average ratio and 22 countries a decrease. Finally, the average trade deficit increased from 6.5 to 9.5 percent, with 41 out of 62 countries experienced a worsening of their trade balance.

Tables 2a and 2b are nevertheless only simple summary statistics. To pin down the partial impact of trade liberalization on the trade balance, one needs regression analysis to control for other factors that also affect the trade balance, which we do in the next section.

4. Regression Analysis

Specification and Data

We follow Santos-Paulino and Thirwall (2004) to use trade balance over GDP as the dependent variable and estimate the following dynamic panel equation:

$$\frac{TB}{GDP}_{it} = \alpha + \beta_1 \frac{TB}{GDP}_{it-1} + \beta_2 lib_{it} + \beta_3 \hat{y}_{it} + \beta_4 \hat{y}^*_{it} + \beta_5 \hat{reer}_{it} + \beta_6 \hat{TOT} + \beta_7 fisr + u_i + v_{it},$$

[8]We excluded former Soviet Union and former Yugoslavia countries due to substantially shorter time series.

where *TB* denotes the trade balance (the lagged dependent variable is included in the equation to control for adjustment dynamics); *lib* is the trade liberalization dummy; $\hat{y}_{it}$ and $\hat{y}^*_{it}$ are domestic and foreign real GDP growth respectively; $\hat{reer}_{it}$ and $\hat{TOT}$ denote the change in (log) real exchange rate and terms of trade respectively. We also include fiscal-balance-to- GDP ratio (*fisr*) to control for the impact of government fiscal policy on the trade balance. Finally, u_i represents time in varying country-specific effects, and v_{it} is a well-behaved disturbance term.

Trade, GDP, and fiscal balance data are from the IMF's International Financial Statistics (IFS) database. Terms of trade data are from the IMF's World Economic Outlook database. Foreign (real) GDP growth is the weighted growth rates of a country's export market countries, where the weight is the market country's 1990 share of the home country's total exports. Bilateral trade data used to calculate the weights are from the IMF's Direction of Trade Statistics database. Finally, the real exchange rate is calculated as a geometric weighted average of bilateral real exchange rates between home country and its trading partners:

$$reer_i = \prod_j \left(\frac{E_{i,us} CPI_i}{E_{j,us} CPI_j} \right)^{W_{ij}},$$

where *i* indicates home country and *j* indicates trading partner countries. $E_{i,us}$ is the nominal exchange rate of country *i* in U.S. dollar per local currency unit, and W_{ij} is the share of country *j* in country *i*'s total trade with its major trading partners. Countries whose trade share in home country is larger than 10 percent are included as major trading partners in calculating *reer* except China, because of incomplete consumer price index (CPI) data (both CPI and bilateral exchange rate data are from the IFS). An increase in *reer* indicates a real appreciation.

Before studying the impact of trade liberalization on the overall trade balance, we first analyze its impact on imports and exports separately. The standard trade equation would use the log of import and export volume as the dependent variable to derive income and price elasticities. This, however, will dramatically reduce our sample size due to missing import/export price data for many countries. Because income and price elasticities are not our primary interests, we use the import-and–export-to-GDP ratio (in log)[9] as the dependent variable in the import and export analyses to maintain our sample size and for consistency between import/export regressions and the trade balance regressions (where trade balance over GDP is the dependent variable).

Impact of Trade Liberalization on Imports

[9]Using the ratios in level yields broadly similar results.

The regression results using our first measure of liberalization dates are reported in Table 3a. The sample covers 39 countries with 1,202 observations. Column one reports the fixed effects panel regression as a benchmark. The trade liberalization dummy is positive and significant at the 1 percent level, indicating that liberalization leads to higher import growth. In addition, higher domestic growth also leads to a higher import-to-GDP ratio, suggesting an income elasticity larger than one. Both real exchange rate appreciation and improved terms of trade (through lower import prices) lead to lower imports (in value), suggesting a price elasticity lower than one.[10] Finally, the positive sign on the fiscal balance is a bit puzzling, as we would expect that an improvement in the fiscal balance lowers the import demand.

However, under the dynamic panel setting fixed effects estimates, even if the country fixed effects assumption is correct, will be consistent only if the time series dimension of the panel goes to infinity. We therefore use the system generalized method of moments (GMM) developed in Blundell and Bond (1998) to get consistent estimates.[11] As a robustness check, we report both one-step and two-step estimates. The two-step procedure involves the additional computation of an optimal weight matrix but is theoretically more efficient. We first follow the standard procedure to use all available lags of the dependent variable and the exogenous regressors in levels dated *t*–2 to all earlier years as instruments in the estimation.[12] However, too many instruments can "overfit" endogenous variables and bias coefficient estimates, as well as weaken Hansen test of instrument validity (Ziliak, 1997; Bowsher, 2002), and it has been suggested that shorter lags of instruments be used (Arellano, 2003; Roodman, 2007). We therefore also report GMM estimates only using lags dated *t*–2 and *t*–3 as instruments (labeled as GMM(2, 3) in the tables). The GMM estimates are reported in columns (2)–(5) of Table 3a.

The results are broadly similar to the fixed effects regression[13] except that the fiscal balance now becomes insignificant and domestic GDP growth becomes insignificant when shorter lags are used as instruments. In all specifications, trade liberalization is shown to lead to higher imports. The Arellano-Bond test confirms the absence of second order correlation of the disturbance term required for consistency, and the Hansen test also does not reject the null hypothesis of joint validity of instruments.[14]

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[10]Developing countries' imports could be more inelastic if the share of imports of intermediate inputs is high.

[11]The Stata program is from Roodman (2006).

[12]This is for the transformed (first-difference) equation. The contemporaneous first difference is used as the instrument in the levels equation.

[13]We note that the fixed effects estimate of the lagged dependent variable is smaller than the GMM estimates as one would expect (Bond, 2002).

[14]A very high *p*-value for the Hansen test, however, is often a sign of instrument proliferation weakening its ability to detect the problem.

Table 3b reports the import regressions using the Wacziarg-Welch measure of trade liberalization dates which covers a larger sample of 77 developing countries (62 of which "opened up" during the sample period) with 2,039 observations. The results are broadly similar to those reported in Table 3a except that the fiscal balance now becomes negative as expected, although insignificant. The trade liberalization dummy is positive and significant at the 1 percent level in all specifications. The estimated coefficients are larger than those reported in Table 3a. For example, for one-step GMM (2, 3), the coefficient on the trade liberalization dummy is 0.074 vs. 0.047 in Table 3a.

Impact of Trade Liberalization on Exports

The regression results for exports are reported in Tables 4a and 4b, for the two measures of trade liberalization dates, respectively. The pattern of coefficients is broadly as expected and consistent across the two measures: higher foreign growth and terms of trade improvement lead to higher exports, and real exchange rate appreciation lowers exports.

The trade liberalization dummy is positive and significant either at the 5 or 10 percent level in all regressions except in the fixed effects regression when the Wacziarg-Welch trade liberalization dates are used. This suggests that developing countries not only import more after liberalizing their trade regime, but also export more. We observe, however, that the coefficients on the trade liberalization dummy from the export regressions tend to be smaller than those from the import regressions. For example, for one-step GMM (2, 3), the coefficients from the export regressions are 0.030 and 0.036 for the two measures of liberalization dates, respectively, while the corresponding coefficients from the import regressions are 0.047 and 0.074, respectively. This indicates that liberalization may lead to higher import growth than export growth, possibly leading to a deterioration in the overall trade balance.[15]

Impact of Trade Liberalization on the Trade Balance

In this section we study the impact of trade liberalization on the overall trade balance. The regression results using the Li measure of liberalization dates are presented in Table 5a. Among the control variables, domestic GDP growth is negative and significant. Foreign GDP growth is positive although only significant in the fixed effects and one-step GMM regressions. The change in real effective exchange rate is negative although insignificant. This is not too surprising given that it is negative in both the import and export regressions. The change in terms of trade is consistently positive and significant. Finally, the fiscal balance is positive as expected, although only significant in the one-step GMM regressions.

The liberalization dummy is negative and significant in the fixed effects regression. However, it becomes insignificant in all the GMM regressions although it remains

[15]Krueger (1978) suggests that there is evidence that import flows respond more rapidly than exports to trade liberalization, causing temporary trade imbalances.

negative. Since GMM yields consistent estimates, the evidence here gives little support to the claim that that trade liberalization has a negative and significant impact on the overall trade balance.

Table 5b reports the results using the Wacziarg-Welch measure of trade liberalization dates. The results for the control variables are again broadly as expected. Higher domestic GDP growth leads to a deterioration of the trade balance, while higher foreign GDP growth improves a country's trade balance. Real exchange rate appreciation also tends to lead to a deterioration in the trade balance, although for the GMM regressions the coefficient is only significant when the shorter list of instruments are used. There is strong evidence across different specifications that positive terms of trade shocks improve the trade balance. For the fiscal balance, although the coefficient is always positive as expected, it is only significant in the fixed effects regression.

In contrast to the results in Table 5a, the trade liberalization dummy is negative and significant in all specifications except in the standard two-step GMM estimation. For example, the one-step GMM (2,3) estimate of the trade liberalization dummy is -1.30, suggesting an immediate worsening of the trade balance-to-GDP ratio of 1.3 percent after liberalization, which we note is substantially smaller than the estimates (–2.52 and –3.57) reported in Santos-Paulino and Thirlwall (2004).

Finally, we re-run the trade balance regressions using the Wacziarg-Welch liberalization dates, but limit the sample to the 39 countries in the Li data set. The results (not reported) are broadly similar to those reported in Table 5b. In particular, the trade liberalization dummy is negative and significant in all specifications except in the standard two-step GMM. This suggests the difference between Tables 5a and 5b is more likely from the difference in the measure of liberalization dates than from the difference in country coverage.

In summary, unlike in the import and export analyses, where we get consistent results across the two measures of liberalization dates, in the analysis of liberalization's impact on the overall trade balance, we get different results depending on the measure used. There is little evidence that liberalization worsens the overall trade balance using the Li measure, but some evidence of a negative impact when the Wacziarg-Welch measure is used.

5. Concluding Remarks

It is a common concern among developing countries that trade liberalization could lead to a deterioration of their trade balance. Despite the importance of the question, cross-country empirical studies on the subject have been scarce. In a recent paper, Santos-Paulino and Thirwall (2004), using a data set of 22 developing countries for the period of 1976–98, find strong evidence of such a negative impact. This paper studied the impact of trade liberalization on imports, exports, and the trade balance for developing countries using two recently compiled measures of trade liberalization dates that cover a much larger sample of developing countries and for longer time periods.

In a simple theoretical model we show that trade liberalization increases both exports and imports, while it has ambiguous effects on the trade balance. Consistent with the theoretical results, we find robust and consistent evidence using both measures that trade liberalization in developing countries promotes both imports and exports. The results, however, are mixed for the impact on the overall balance depending on the liberalization measure used. Using an extended Li (2004) measure of liberalization dates, we find little evidence of a statistically significant negative impact of liberalization on the overall trade balance. There is, however, some evidence that liberalization worsens the trade balance when the Wacziarg-Welch liberalization dates are used, although the evidence is not robust across different estimation specifications. And even in this case, the estimated impact is smaller than that reported by Santos-Paulino and Thirwall (2004).

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Table 1. Trade Liberalization Episodes (Measure of Liberalization Dates Based on Li, 2004)

Country	70	71	72	73	74	75	76	77	78	79	80	81	82	83	84	85	86	87	88	89	90	91	92	93	94	95	96	97	98	99	00	01	02	03	04	W-W open year
Argentina	0	0	0	0	0	0	x	x	x	x	0	0	0	0	0	0	0	0	x	x	x	x	0	0	0	0	0	0	0	0	0	0	0	0	0	1991
Benin	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	0	0	0	0	0	0	0	0	0	0	0	1990
Brazil	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	x	0	0	0	0	0	0	0	0	0	0	0	1991
Cameroon	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	0	0	0	0	0	0	0	0	0	0	1993
Chile	0	0	0	0	x	x	x	x	x	x	0	0	0	0	0	x	x	x	x	x	x	x	x	0	0	0	0	0	0	0	0	0	0	0	0	1976
China	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	x	0	0	0	0	0	x	0	n.a.
Colombia	0	0	0	x	x	x	x	x	x	x	0	0	0	0	0	x	x	x	x	x	0	0	x	0	0	0	0	0	0	0	0	0	0	0	0	1986
Costa Rica	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	1986
Ecuador	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	x	x	0	0	0	0	0	0	0	0	0	0	0	0	1991
The Gambia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	1985
Ghana	0	0	0	0	0	0	0	0	0	0	0	0	0	x	0	0	x	x	x	x	x	x	x	0	0	0	0	0	0	0	0	0	0	0	0	1985
Guatemala	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	1988
Guinea-Bissau	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	0	0	0	0	0	0	0	0	0	0	0	0	x	0	0	0	0	1987
Guyana	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	0	0	0	0	0	0	0	x	x	x	0	0	0	1988
Honduras	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	0	0	0	0	0	0	0	0	0	0	0	0	1991
India	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	0	0	x	x	x	x	x	x	x	0	0	0	0	0	0	0	n.a.
Indonesia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	x	x	x	x	x	x	x	0	0	0	0	0	0	0	0	1970
Jamaica	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	0	0	0	0	0	x	x	0	0	0	0	1989
Kenya	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	x	0	0	0	0	0	0	x	x	x	x	x	1993
Malaysia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	0	0	0	0	0	0	0	0	0	1963
Mali	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	0	0	0	0	0	0	0	0	0	0	0	0	0	1988
Mauritania	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	0	0	0	0	0	0	0	0	0	0	x	0	0	0	0	1995
Mexico	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	1986
Morocco	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	x	x	0	0	0	0	0	0	0	x	0	0	0	0	0	0	0	1984
Nepal	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	0	0	0	0	0	0	0	0	0	0	0	1991
Nicaragua	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	0	0	0	x	x	0	0	0	0	0	0	0	0	0	0	0	0	1991
Nigeria	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	n.a.
Pakistan	0	0	x	x	x	x	x	x	x	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	x	x	0	0	x	x	x	x	x	0	0	2001
Paraguay	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	1989
Peru	0	0	0	0	0	0	0	0	0	x	x	x	0	0	0	0	0	0	0	0	0	x	0	0	0	0	0	0	0	0	0	0	0	0	0	1991
Philippines	0	0	0	0	0	0	0	0	0	0	0	x	x	x	0	0	x	0	0	0	0	x	x	x	x	x	x	x	x	x	0	0	0	0	0	1988
Sri Lanka	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	0	0	0	0	x	0	0	0	0	0	0	1991
Thailand	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	0	0	0	x	x	x	0	0	x	x	x	x	x	x	Always
Tunisia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	x	x	x	0	0	0	0	0	0	0	0	0	0	0	1989
Turkey	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	x	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	1989
Uganda	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	0	0	0	0	0	0	0	1988
Uruguay	0	0	0	0	x	x	x	x	x	x	x	x	0	0	0	0	0	0	0	0	0	x	x	x	x	0	0	0	0	0	0	0	0	0	0	1990
Venezuela	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	0	0	0	0	0	0	0	0	0	0	0	0	1996
Zambia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	x	x	x	x	x	0	0	0	0	0	0	0	0	1993

Sources: (i) Li (2004); (ii)the United Nations Conference on Trade and Development's Trade Analysis and Information System (TRAINS) database; (iii) the IMF's Trade Policy Information Database (TPID); (iv) Wacziarg and Welch (2003).

Table 2a. Import, Export, and Trade Balance to GDP Ratios Before and After Trade Liberalization (Extended Li Trade Liberalization Measure, 1970–2004)

	Imports/GDP (%)		Exports/GDP (%)		Trade Balance/GDP (%)	
Country	Before lib.	After lib.	Before lib.	After lib.	Before lib.	After lib.
Argentina	6.2	9.1	8.1	11.5	2.0	2.4
Benin	27.6	28.8	10.3	18.6	-17.4	-10.2
Brazil	7.5	9.1	8.1	9.7	0.6	0.6
Cameroon	16.8	15.6	15.5	18.1	-1.3	2.5
Chile	18.6	24.4	18.8	24.4	0.3	0.0
China	11.5	20.1	11.4	22.4	-0.1	2.3
Colombia	12.2	15.2	11.7	13.2	-0.4	-2.0
Costa Rica	31.3	35.7	24.3	29.5	-7.0	-6.2
Ecuador	17.9	21.9	21.2	22.6	3.3	0.7
Gambia, The	52.1	54.5	26.8	8.4	-25.4	-46.2
Ghana	22.7	42.2	22.0	26.5	-0.7	-15.7
Guatemala	17.3	24.6	15.4	14.0	-2.0	-10.6
Guinea-Bissau	37.9	29.7	9.5	28.6	-28.4	-1.1
Guyana	69.9	81.0	62.9	71.2	-7.0	-9.9
Honduras	29.7	44.7	25.8	24.8	-3.9	-19.8
India	6.9	11.4	5.4	9.2	-1.5	-2.2
Indonesia	15.6	24.0	22.0	34.3	6.4	10.3
Jamaica	40.5	42.2	24.1	14.5	-16.4	-27.7
Kenya	26.3	…	17.0	…	-9.3	…
Malaysia	47.0	83.0	52.4	97.7	5.4	14.7
Mali	22.2	29.6	10.4	20.5	-11.8	-9.0
Mauritania	31.6	32.6	36.3	26.7	4.7	-6.0
Mexico	8.9	24.1	9.0	22.0	0.0	-2.1
Morocco	24.8	32.2	15.3	20.8	-9.4	-11.3
Nepal	15.4	28.8	5.9	10.6	-9.5	-18.3
Nicaragua	34.8	41.6	21.2	14.9	-13.5	-26.7
Nigeria	23.4	23.8	29.4	40.3	5.9	16.6
Pakistan	17.5	17.3	11.6	14.2	-5.9	-3.1
Paraguay	13.3	31.1	9.3	15.1	-4.0	-15.9
Peru	11.7	13.6	13.8	12.1	2.1	-1.5
Philippines	27.0	49.0	19.8	46.0	-7.3	-2.9
Sri Lanka	31.1	38.9	22.4	29.7	-8.8	-9.1
Thailand	32.5	…	28.2	…	-4.3	…
Tunisia	35.2	43.3	22.0	30.7	-13.2	-12.6
Turkey	10.5	21.0	5.6	13.6	-4.9	-7.5
Uganda	13.4	22.5	12.7	8.5	-0.7	-14.0
Uruguay	14.4	17.5	13.8	13.8	-0.6	-3.6
Venezuela	17.8	15.6	24.8	27.3	7.0	11.7
Zambia	27.7	31.9	35.6	26.5	8.0	-5.4
Average	23.5	30.6	19.3	24.1	-4.2	-6.5
Before<After	33		28		15	
Before>After	4		9		22	

Source: author's calculation based on data from the IMF's International Financial Statistics (IFS) database.

Table 2b. Import, Export, and Trade Balance to GDP Ratios Before and After Trade Liberalization (Wacziarg-Welch Trade Liberalization Measure, 1970–2001)

	Imports/GDP (%)		Exports/GDP (%)		Trade balance/GDP (%)	
Country	Before lib.	After lib.	Before lib.	After lib.	Before lib.	After lib.
Average	25.1	29.9	18.5	20.4	-6.5	-9.5
Before<After	47		40		21	
Before>After	15		22		41	

Source: author's calculation based on data from the IMF's International Financial Statistics (IFS) database.

Table 3a. Trade Liberalization and Imports (Extended Li Trade Liberalization Measure, 1970–2004)

Dependent variable: Imports/GDP (in log)	(1) Fixed effects	(2) GMM (one-step)	(3) GMM (2, 3) (one-step)	(4) GMM (two-step)	(5) GMM (2, 3) (two-step)
Lagged dependent variable	0.778***	0.897***	0.854***	0.883***	0.859***
	(0.018)	(0.023)	(0.037)	(0.072)	(0.046)
Trade liberalization	0.082***	0.037***	0.047***	0.041*	0.043**
	(0.013)	(0.010)	(0.011)	(0.024)	(0.017)
Domestic GDP growth	0.005***	0.005**	0.004	0.005*	0.004
	(0.001)	(0.002)	(0.002)	(0.003)	(0.002)
Change in real effective	-0.115***	-0.141***	-0.127***	-0.135**	-0.124**
exchange rate	(0.026)	(0.050)	(0.049)	(0.054)	(0.054)
Changes in terms of	-0.001***	-0.001***	-0.002**	-0.002**	-0.002**
trade	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
Fiscal balance/GDP	0.003*	0.001	0.002	0.001	0.003
	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)
No. of countries	39	39	39	39	39
No. of observations	1,202	1,202	1,202	1,202	1,202
Arellano-Bond test for AR(2) in first differences (p-value)		0.36	0.37	0.36	0.37
Hansen test of joint validity of instruments (p-value)		1.00	1.00	1.00	1.00

Note: *, **, and *** represent significant at 10, 5, and 1 percent level. Robust standard errors are in parenthesis, with robust standard errors for the two-step estimates calculated using the Windmeijer (2005) correction.

Table 3b. Trade Liberalization and Imports (Wacziarg-Welch Trade Liberalization Measure, 1970–2001)

Depedent variable: Imports/GDP (in log)	(1) Fixed effects	(2) GMM (one-step)	(3) GMM (2, 3) (one-step)	(4) GMM (two-step)	(5) GMM (2, 3) (two-step)
Lagged dependent variable	0.767***	0.811***	0.793***	0.812***	0.787***
	(0.014)	(0.029)	(0.037)	(0.038)	(0.039)
Trade liberalization	0.061***	0.069***	0.074***	0.070***	0.077***
	(0.012)	(0.016)	(0.019)	(0.021)	(0.020)
Domestic GDP growth	0.003***	0.003*	0.002	0.002	0.001
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
Change in real effective	-0.124***	-0.127**	-0.120**	-0.126**	-0.118**
exchange rate	(0.022)	(0.050)	(0.049)	(0.052)	(0.051)
Changes in terms of	-0.001***	-0.001**	-0.001**	-0.001**	-0.001**
trade	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Fiscal balance/GDP	-0.001	-0.001	-0.001	-0.001	-0.001
	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
No. of countries	77	77	77	77	77
No. of observations	2,039	2,039	2,039	2,039	2,039
Arellano-Bond test for AR(2) in first differences (p-value)		0.43	0.44	0.43	0.46
Hansen test of joint validity of instruments (p-value)		1.00	0.93	1.00	0.93

Note: *, **, and *** represent significant at 10, 5, and 1 percent level. Robust standard errors are in parenthesis, with robust standard errors for the two-step estimates calculated using the Windmeijer (2005) correction.

Table 4a. Trade Liberalization and Exports (Extended Li Trade Liberalization Measure, 1970–2004)

Dependent variable: Exports/GDP (in log)	(1) Fixed effects	(2) GMM (one-step)	(3) GMM (2, 3) (one-step)	(4) GMM (two-step)	(5) GMM (2, 3) (two-step)
Lagged dependent variable	0.855***	0.924***	0.874***	0.919***	0.882***
	(0.015)	(0.021)	(0.043)	(0.025)	(0.050)
Trade liberalization	0.040***	0.025**	0.030**	0.026**	0.033*
	(0.013)	(0.012)	(0.015)	(0.013)	(0.018)
Foreign GDP growth	0.007*	0.011***	0.010***	0.013***	0.011***
	(0.004)	(0.003)	(0.004)	(0.004)	(0.004)
Change in real effective	-0.225***	-0.217***	-0.199**	-0.207**	-0.219**
exchange rate	(0.028)	(0.085)	(0.095)	(0.086)	(0.093)
Changes in terms of	0.003***	0.002***	0.002***	0.002***	0.002***
trade	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
No. of countries	39	39	39	39	39
No. of observations	1,203	1,203	1,203	1,203	1,203
Arellano-Bond test for AR(2) in first differences (p-value)		0.84	0.87	0.87	0.88
Hansen test of joint validity of instruments (p-value)		1.00	1.00	1.00	1.00

Note: *, **, and *** represent significant at 10, 5, and 1 percent level. Robust standard errors are in parenthesis, with robust standard errors for the two-step estimates calculated using the Windmeijer (2005) correction.

Table 4b. Trade Liberalization and Exports (Wacziarg-Welch Trade Liberalization Measure, 1970–2001)

Dependent variable: Exports/GDP (in log)	(1) Fixed effects	(2) GMM (one-step)	(3) GMM (2, 3) (one-step)	(4) GMM (two-step)	(5) GMM (2, 3) (two-step)
Lagged dependent variable	0.821***	0.895***	0.860***	0.897***	0.856***
	(0.013)	(0.016)	(0.032)	(0.027)	(0.038)
Trade liberalization	0.019	0.028**	0.036**	0.031**	0.042**
	(0.012)	(0.013)	(0.016)	(0.014)	(0.018)
Foreign GDP growth	0.005	0.008***	0.007**	0.008***	0.007**
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Change in real effective	-0.226***	-0.224***	-0.214***	-0.226***	-0.214***
exchange rate	(0.023)	(0.070)	(0.072)	(0.073)	(0.074)
Changes in terms of	0.003***	0.003***	0.003***	0.003***	0.003***
trade	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
No. of countries	77	77	77	77	77
No. of observations	2,055	2,055	2,055	2,055	2,055
Arellano-Bond test for AR(2) in first differences (p-value)		0.82	0.84	0.83	0.86
Hansen test of joint validity of instruments (p-value)		1.00	0.91	1.00	0.91

Note: *, **, and *** represent significant at 10, 5, and 1 percent level. Robust standard errors are in parenthesis, with robust standard errors for the two-step estimates calculated using the Windmeijer (2005) correction.

Table 5a. Trade Liberalization and the Trade Balance (Extended Li Trade Liberalization Measure, 1970–2004)

Dependent variable: Trade balance/GDP	(1) Fixed effects	(2) GMM (one-step)	(3) GMM (2, 3) (one-step)	(4) GMM (two-step)	(5) GMM (2, 3) (two-step)
Lagged dependent var.	0.694***	0.883***	0.787***	0.919***	0.742***
	(0.020)	(0.035)	(0.053)	(0.099)	(0.071)
Trade liberalization	-0.951***	-0.395	-0.588	-4.591	-0.259
	(0.307)	(0.292)	(0.375)	(4.403)	(1.530)
Domestic GDP growth	-0.185***	-0.162**	-0.143*	-0.177***	-0.188**
	(0.032)	(0.071)	(0.085)	(0.066)	(0.083)
Foreign GDP growth	0.156*	0.187***	0.167**	0.048	0.118
	(0.090)	(0.065)	(0.079)	(0.218)	(0.139)
Change in real effective	-0.810	-0.957	-0.839	-0.356	-1.106
exchange rate	(0.682)	(1.643)	(1.594)	(1.673)	(1.325)
Changes in terms of	0.086***	0.089***	0.083***	0.080***	0.079***
trade	(0.009)	(0.026)	(0.026)	(0.023)	(0.024)
Fiscal balance/GDP	0.036	0.090*	0.116*	0.10	0.143*
	(0.039)	(0.053)	(0.071)	(0.069)	(0.082)
No. of countries	39	39	39	39	39
No. of observations	1,202	1,202	1,202	1,202	1,202
Arellano-Bond test for AR(2) in first differences (p-value)		0.90	0.88	0.81	0.88
Hansen test of joint validity of instruments (p-value)		1.00	1.00	1.00	1.00

Note: *, **, and *** represent significant at 10, 5, and 1 percent level. Robust standard errors are in parenthesis, with robust standard errors for the two-step estimates calculated using the Windmeijer (2005) correction.

Table 5b. Trade Liberalization and the Trade Balance (Wacziarg-Welch Trade Liberalization Measure, 1970-2001)

Dependent variable: Trade balance/GDP	(1) Fixed effects	(2) GMM (one-step)	(3) GMM (2, 3) (one-step)	(4) GMM (two-step)	(5) GMM (2, 3) (two-step)
Lagged dependent variable	0.637***	0.842***	0.752***	0.839***	0.754***
	(0.017)	(0.029)	(0.042)	(0.035)	(0.042)
Trade liberalization	-1.260***	-0.844**	-1.300***	-0.843	-1.249***
	(0.294)	(0.334)	(0.484)	(1.059)	(0.466)
Domestic GDP growth	-0.106***	-0.079**	-0.077*	-0.083**	-0.085**
	(0.022)	(0.037)	(0.040)	(0.038)	(0.037)
Foreign GDP growth	0.217***	0.214***	0.237***	0.211***	0.227***
	(0.070)	(0.070)	(0.073)	(0.081)	(0.076)
Change in real effective	-1.887***	-1.910	-2.033*	-1.953	-2.235**
exchange rate	(0.553)	(1.188)	(1.175)	(1.324)	(1.136)
Changes in terms of	0.076***	0.082***	0.079***	0.082***	0.078***
trade	(0.007)	(0.016)	(0.015)	(0.017)	(0.015)
Fiscal balance/GDP	0.071***	0.04	0.05	0.04	0.05
	(0.022)	(0.040)	(0.049)	(0.041)	(0.048)
No. of countries	77	77	77	77	77
No. of observations	2,039	2,039	2,039	2,039	2,039
Arellano-Bond test for AR(2) in first differences (p-value)		0.75	0.76	0.75	0.78
Hansen test of joint validity of instruments (p-value)		1.00	0.98	1.00	0.98

Note: *, **, and *** represent significant at 10, 5, and 1 percent level. Robust standard errors are in parenthesis, with robust standard errors for the two-step estimates calculated using the Windmeijer (2005) correction.

CHAPTER THREE

Can Oil Prices Really Help to Forecast Real Effective Exchange Rates?

Li Zeng[1]

This paper empirically tests the proposition that domestic real oil prices can help to forecast real effective exchange rates. For most countries in our sample, little evidence was found to support such a proposition. There were two exceptions, Japan and Norway, where we did see some consistent signs in favor of the oil-price exchange rate model. But the statistics favoring the model are mostly insignificant. We also find that the relations between oil prices and real exchange rates are not stable over time. The conjecture that the oil-price exchange rate model can produce better forecasts on exchange rates when oil prices are more volatile is not consistent with the pattern seen in our subsample results. [JEL F31, F37]

1. Introduction

Since Meese and Rogoff (1983a and 1983b) presented striking evidence that some empirical exchange rate models with nice in-sample fit failed to forecast more accurately than a naïve random walk model, out-of-sample predictive power has become an important criterion in judging the performance of exchange rate models.[2] As a piece of pioneering work in the field now known as *commodity currencies*, Amano and van Norden (1998) showed that for three advanced economies (the United States, Japan and Germany), the domestic real oil prices could help to project their real effective exchange rates.[3] Such a finding was attributed by the authors to the long-run relation between these countries' domestic real oil prices and their real effective exchange rates, which was in turn argued as the result of the close correlations between the oil prices and these countries' terms of trade.

There is, however, an eyebrow-raising part in this otherwise very neat study, that is, its interpretation on the difference between the out-of-sample forecast performance by the oil-price exchange rate model for the two forecasting periods investigated in the paper. Amano and van Norden (1998) evaluated the oil-price model's forecast performance for two out-of-sample periods, starting in 1985 and 1989, respectively. Consistent evidence favoring

[1] I am grateful to Professor Maurice Obstfeld for his great guidance, and would also like to thank Professor Andrew Rose for very helpful discussions.

[2] For a recent study on the out-of-sample forecast performance by the monetary exchange rate model, please see Cerra and Saxena (2008).

[3] Notable studies on commodity currencies include Chen and Rogoff (2003), Cashin, Cespedes and Sahay (2004), and Chen, Rogoff and Rossi (2009), among others.

the oil-price model over the simple random walk model was reported for the post-1985 period, but only mixed results were found for the other. Interestingly, such difference was interpreted by the authors as evidence supporting the oil-price model. Their argument was that the relative performance of the oil-price exchange rate model would be better when oil prices were more volatile, yet the oil prices had entered a relative tranquil period since the late 1980s, causing the model to perform poorly for the post-1989 period.

Was the relative stability in oil prices the true reason behind the oil-price model's poor forecast performance for the post-1989 period? Or was it merely by luck that Amano and van Norden (1998) identified a period (the post-1985 one) when the oil-price model happened to perform better than the random walk model? If there is anything positive coming out of the current global economic crisis, it offers economists many good opportunities of *natural experiments*. In the past few years, we have witnessed roller-coaster type of movements in oil prices. If the arguments by Amano and van Norden (1998) were indeed correct, that is, oil prices could help to forecast exchange rates and the oil-price model would perform better when oil prices were more volatile, then we should again expect good forecast performance by the oil-price model. Is it really the case? With such curiosity, we want to revisit this topic and check whether oil prices can really help to forecast exchange rates.

Although Amano and van Norden (1998) emphasized that it was only for the three oil-importing advanced economies that they identified the long-run relations between domestic oil prices and real effective exchange rates, we will try to check the forecast ability of the oil-price exchange rate model for a more diversified group of countries. In addition to the United States (US), Japan and Germany in the original study, our sample also includes the United Kingdom (UK), Canada, Norway, Mexico and Turkey. In contrast to the US, Japan and Germany, who are mainly oil importers, the international trade in oil products by the UK and Canada is more balanced. Norway, on the other hand, is an oil-exporting high income country. Mexico, another oil-exporting country, and Turkey, another oil-importing country, are also included in the sample to allow us to have a peek on how the oil-price exchange rate model might work for different types of emerging market economies.[4]

This paper extends the original work by Amano and van Norden (1998) not only in the sense that a larger sample of countries and a longer time horizon are being studied, but also that a more thorough investigation is being carried out by looking at various alternative specifications of the oil-price model and evaluating their out-of-sample forecast performance using different criteria. The econometric specifications estimated by this

[4] Acknowledgedly this paper is far from exhausting the countries that data would allow us to study. However, our selection of sample countries, especially the two emerging market economies, Mexico and Turkey, is not entirely random. One factor that may lead the oil-price model to fail for the emerging market economies is the heavy intervention on the foreign exchange market by the authorities. To minimize the impact of such noise, we purposely choose two emerging market economies whose exchange rates have been allowed to float relatively freely for a substantial period of time. For more details on the history of countries' exchange rate arrangements, please see Reinhart (2002).

paper include the vector autoregression (VAR) model, in both levels and first-order difference, and the vector error correction (VEC) model. Two types of out-of-sample forecasts are examined. One is *true* dynamic out-of-sample forecasts, where both future exchange rates and future oil prices are forecasted. This type of forecasts are what typically produced by time-series models such as VAR and VEC. The other type of exchange rate forecasts are based on *ex post* oil price information. The purpose of studying of this second type of forecasts is to address the concern that it might be the poor forecastability of the oil prices, rather than their lack of predictive power on exchange rates, that leads to bad out-of-sample projections on exchange rates. Since such forecasts contain *ex post* information, they actually give the oil-price model extra edge in its comparison with the random walk models. To explore the potential nonlinearity in the relation between oil prices and exchange rates, we include higher order terms of the regressors in one of the specifications. There might as well be doubts that omission of other important determinants of exchange rates by the oil-price model would affect its forecast performance. Thus we will check the robustness of our evaluation results by adding additional control variables, such as growth rates of exports, to the model. Since Amano and van Norden (1998) suggested terms of trade as the linking channel between oil prices and exchange rates, we will make a detour and check the predictive power of terms of trade on exchange rates too. Additionally, the stability of the relation between oil prices and exchange rates is investigated through the subsample study, which also helps us to verify whether the oil-price model indeed performs better when oil prices are more volatile.

In evaluating the oil-price model's out-of-sample forecast performance, three different criteria (U-statistics) are adopted by the paper. The first one is the Theil's U-statistic, which compares the root mean squared forecast errors (RMSEs) of the oil-price exchange rate model to those of the simple random walk model. The other two criteria are also ratio statistics between the oil-price exchange rate model and the random walk model. But instead of the RMSEs, they are based on the mean absolute forecast errors (MAEs) and the proportions of forecasts that correctly predict the direction of change in the exchange rate, respectively.

The answers we find to the title question of this paper are not exactly the same across all the sample countries. For most of them, contrary to the results of Amano and van Norden (1998), little evidence is found to support the oil-price exchange rate model as a better forecasting tool than the random walk model. The two exceptions are Japan and Norway, where we do see some consistent, yet mostly weak, results favoring the oil-price exchange rate model. The subsample results of this paper suggest that the relations between domestic real oil prices and real effective exchange rates in the sample countries are quite unstable. Time-varying patterns in the forecast results are seen not only for the six countries where the oil-price model receives little support, but also for Japan and Norway. Besides, the conjecture that the oil-price model would perform better when oil prices are more volatile does not seem to be consistent with the pattern seen in our subsample results.

The rest of the paper is organized as the following. Section 2 discusses data and model specifications, explains the evaluation criteria for out-of-sample forecasts, and reports the benchmark results. Sections 3 checks the robustness of the evaluation results by looking at various alternative specifications and scenarios, including subsample analysis. Section 4 summarizes our main findings and concludes the paper.

2. Can Oil Prices Really Help to Forecast Exchange Rates?

The focus of this paper is to check whether the oil-price exchange rate model proposed by Amano and van Norden (1998) can help to produce better out-of-sample forecasts than naïve models such as a random walk. To stay close to the original study, the Morgan Guaranty real effective exchange rate (REER) series are used for most of our sample countries. The only exception is Germany, because the Morgan Guaranty REER series for the Deutsche Mark has discontinued since the introduction of the Euro in 1999. Instead the REER series from the International Financial Statistics is used for Germany. The domestic real oil prices for each country are obtained by first converting the US dollar prices of the West Texas intermediate crude oil into local currencies and then deflating them using domestic consumer price indices. For most countries, the monthly REER and oil price series used in our estimations cover the period from 1973M1 to 2009M7. The samples for Germany and Mexico are slightly shorter, because the REER series for these two countries became available only in 1975M1. More details of the data and the summary statistics are provided in the appendix, and the REER and oil price series are graphed in Figure-1.

The benchmark specification of the oil-price exchange rate model in this paper is a two-variable, exchange rate and oil price, VAR system in levels. This is different from Amano and van Norden (1998), where the model was estimated in its VEC format. The VEC format, which is a constrained version of the VAR specification, would be more appropriate if the data series in the model are nonstationary but cointegrated with each other. However, for the oil-price exchange rate model, neither do we have such a priori knowledge, nor can the unit root (or stationarity) and cointegration tests lead us to such conclusions with enough confidence.[5] Hence, VAR in levels, the nonconstrained version of the model, is used as our benchmark specification. Nonetheless, to check the robustness of our findings, we will estimate the model again in the next section using both the VEC specification and the VAR specification in first-order difference. The latter would be most

[5] For the data series, we carried out three unit root tests and one stationarity test. They are the Augmented Dickey-Fuller test, the Phillips-Perron test, the DF-GLS test, and the Kwiatkowski-Phillips-Schmidt-Shin stationarity test. Conflicting results often arise, not only from different tests, but also from the same test with different lag choices. Similarly, the cointegration tests often lead us to ambiguous results too. For instance, in some cases, with the same choice of lag orders, the Johansen (1995) test would suggest two cointegrating vectors between exchange rates and oil prices at a 95 percent confidence level, which implies that both variables being tested are stationary, yet none at the 99 percent confidence level. Since we will check the robustness of our findings by estimating alternative specifications (VAR, in both levels and first-order difference, and VEC), the unit root, stationarity, and cointegration test results are not reported.

appropriate if both the oil price and exchange rate series are I(1) processes yet not cointegrated with each other.[6]

Three different criteria are adopted in this paper to evaluate the forecast performance of the oil-price exchange rate model. The first one is the Theil's U-statistic (or U-ratio), which compares the oil-price model to the simple random walk model. It is the ratio between the root mean squared forecast errors (RMSEs) of the two models. If the oil-price model performs better, the Theil's U-statistic will be less than unity. Following Meese and Rogoff (1983a), this measure is regarded as the principal criterion for our evaluations. The second measure is similar to the Theil's U-statistic, but with the RMSEs replaced by the mean absolute forecast errors (MAEs). The MAE U-statistic is more meaningful in situations where the exchange rates have fat-tailed distributions or the estimation results are heavily affected by outlier observations, because it is the median rather than the mean that a MAE-minimizing method tries to project. The statistical significance of these two U-statistics are gauged by the Diebold-Mariano (1995) test. The last criterion is still a ratio. But instead of the RMSEs or the MAEs, it is based on the proportions of forecasts that correctly predict the direction of change in the exchange rate. Since the simple random walk model always predicts no change in future exchange rates, it has to be modified before being used as the comparison benchmark for this measure. Taking its spirit of "same as the last period", we adopt a simple rule to forecast the direction of future exchange rate changes, that is, future exchange rate changes will be in the same direction as the change in the last observed period. This simple forecast rule forms the comparison benchmark for the last evaluation criterion. For simplicity of future reference, we call this rule the modified random walk model and the last criterion directional U-statistic. To be consistent with previous two U-statistics, the directional U-statistic is calculated in the way that a value larger than unity implies a poor performance by the oil-price exchange rate model relative to the modified random walk model. The statistical significance of the directional U-statistic is obtained from the studentized *t*-test.[7]

To evaluate the forecast ability of the oil-price exchange rate model, we will recursively produce out-of-sample exchange rate projections for each period starting in 1988M1, for up to 24 months ahead. For example, in the case of the US, the first forecast estimation uses data from 1973M1 to 1987M12 and produces out-of-sample exchange rate projections for 1988M1 to 1989M12. Estimations toward the late end of the data coverage produce fewer out-of-sample forecasts, because there is no actual information available yet to compare some of the long-term projections. For instance, again in the case of the US, the last

[6] In fact, even if the data series contain unit roots, consistent estimates of the VAR coefficients can still be obtained with the classical methods. For more discussions on VAR versus VEC, please see Fabio Canova (2007).

[7] Measures like the directional U-statistic are often thought of as loose metrics, because the sizes of the forecast errors are totally ignored. They are sometimes called utility-based criteria. For more details on related discussions, please see, for example, Leitch and Tanner (1991) and Engel (1994).

forecast estimation uses data from 1973M1 to 2009M6 and produces only one out-of-sample projection for 2009M7.

The benchmark evaluation outcomes are reported in Table-1. Instead of providing the statistics for all forecast horizons, we select three short term ones (at 1-3 months), three medium term ones (at 6, 9 and 12 months) and two long term ones (at 18 and 24 months). In this table Japan appears to be the strongest case supporting the oil-price exchange rate model, because it is the only country with more than half of the U-statistics below unity. However, it is apparent that even this strongest case is indeed quite weak. The exact number of U-statistics smaller than one for Japan is 14, exceeding half of the total only by two, and six of them are directional ones. More importantly, among these 14 statistics, only one is significant, which is the directional U-statistic at the 18 months horizon. There are two other countries, Norway and Mexico, whose results can also be interpreted as slightly in favor of the oil-price exchange rate model. The reason is that both of these countries have more than half of the Theil's U-statistics smaller than one, and the Theil's ratio, as mentioned earlier, is regarded as the principal criterion for our evaluations. Nonetheless, like in the case of Japan, the supports received by the oil-price model from these two countries are actually quite weak, because none of the U-ratios is statistically significant.
The results for the other five countries are clearly against the oil-price exchange rate model, although the statistical significance of the evidence varies from one country to another. Canada has the strongest results, with 22 of its 24 U-statistics above unity, including all 8 Theil's U-statistics. Furthermore, 9 of these 22 statistics are significant and 4 of them are the Theil's U-ratios. The numbers of U-statistics above unity are 21, 20, 17 and 20 for the US, Germany, the UK and Turkey, respectively, which again include most of the Theil's U-ratios. However, unlike the case of Canada, none of the Theil's U-ratios above one for these countries is statistically significant.

If we compare across different forecast horizons, it appears that the U-statistics below unity mostly fall into the short-term and long-term ranges. Among the 55 U-statistics smaller than one for all countries, 20 are short-term ones and 26 are long-term ones. Only 9 of them are for the medium-term forecasts. However, such a pattern should be observed with caution, because most of the U-ratios are statistically insignificant. The results in Table-1 also seem to indicate that the oil-price exchange rate model works better for the oil-exporting countries in our sample. As discussed earlier, the outcomes for both *pure* oil-exporting countries, Norway and Mexico, are slightly in favor of the oil-price model. Some words of caution should be said here. First, the evidence supporting the oil-price model is actually quite weak for both countries. Secondly, even if the evidence for Norway and Mexico were consistent and significant, the observations for two countries are far from sufficient to establish the result as a general rule for all oil-exporting countries.

As a summary for this section, very weak evidence supporting the oil-price exchange rate model is found for three countries in our sample, Japan, Norway and Mexico. The evaluation outcomes for the other five economies show that the forecast performance of the oil-price model is actually worse than the simple random model, and some of the evidence found for Canada is statistically significant. Comparisons across different forecast horizons and across the sample countries seem to suggest that the oil-price exchange rate model forecasts better for the short term and the long term than for the medium term, and better for *pure* oil-exporting countries than for other types of economies. Yet, since most of the evaluation results are statistically insignificant, all the observations should be taken with caution and we will check their robustness in the next section.

3. Are We Missing Anything?

Our benchmark evaluation results in the previous section do not seem to lend much support to the claim by Amano and van Norden (1998) that oil prices could help to forecast exchange rates. In this section, we will check the robustness of our findings by investigating alternative specifications and scenarios.

3.1 Underspecification of the VAR model?

In Table-1 the lag orders of the VAR models for each country were selected based on the Schwarz's Bayesian Information Criterion (SBIC).[8] Since the SBIC tends to suggest short lags, to make sure that the evaluation outcomes do not suffer from underspecification of the models, we repeat the estimations in Table-1 but now use the likelihood ratio statistics to choose the lag orders. Among the various selection criteria we look at, lag order choices based on the likelihood ratio statistics are usually the longest. The new results are reported in Table-2.

The patterns of changes, compared to the results in Table-1, are the same across all the sample countries. Including more lags in the VAR specification makes the forecast performance of the oil-price exchange rate model even worse. For instance, for both Norway and Mexico, the numbers of U-statistics above unity are now 22, which include all the Theil's U-ratios. The oil-price model also loses its edge in Japan. Exactly half of the U-statistics are now smaller than one, but only 2 of them are the Theil's U-ratios. Similar changes also happened to the other five countries. Either they have more U-ratios above one now, or more U-ratios above one become statistically significant, or both.

According to the results in Table-2, underspecification of the VAR models does not seem to be the source of poor forecast performance for the oil-price exchange rate model.[9] The same exercise of checking alternative specifications with more lagged terms is also carried out for most estimations in the coming subsections. As seen here, adding additional lags to the SBIC selections usually makes the forecasts of the oil-price model worse. Thus those results are not reported hereafter.

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[8] Instead of choosing a lag order for each one of the recursive estimations, we selected the lag order of the VAR model for each country only once, using the entire sample.

[9] We als tried to select the lag orders following other criteria, such as the Akaike's information criterion. The conclusions are broadly the same.

3.2 Misspecification?

The next concern we try to address is the misspecification of the oil-price exchange rate model, which can arise from nonstationarity of the data series. If the exchange rate and oil price series are both nonstationary but cointegrated with each other, then a VEC specification would be most appropriate. If the two series are both I(1) processes yet not cointegrated, then theoretically a VAR specification in first-order difference should be considered. Since the statistical test results are ambiguous, we check the robustness of our findings by estimating both alternative specifications.

Table-3 reports the evaluation results for the VEC specification of the oil-price exchange rate model. Compared to the benchmark in Table-1, some mixed changes are seen here. For the US, Japan, the UK, Mexico and Turkey, the evidence turned less favorable for the oil-price model. The most striking case is Mexico. Previously we interpreted the results of Mexico in Table-1 as slightly in favor of the oil-price model. However, after the VEC specification was adopted, all 24 U-statistics for Mexico became larger than one in Table-3. A similar but less dramatic shift also happened to Japan, another country whose benchmark outcome supported the oil-price model. With the VEC specification, 16 of Japan's U-statistics are now above unity, including 5 Theil's ratios. For the US, although the total number of U-statistics above unity is one fewer than in Table-1, 5 of such Theil's U-ratios are now statistically significant, as compared to none in Table-1. For the UK and Turkey, the UK sees its number of U-ratios above unity increase from 17 in Table-1 to 20 in Table-3, and Turkey now has 5 more U-statistics significantly above unity.

Under the VEC specification, the forecast performance of the oil-price exchange rate model improved for Germany, Canada and Norway. In the case of Norway, the U-statistics supporting the oil-price model increased from 12 to 17, although as before, none of them is significant. Canada was the strongest case against the oil-price model under the benchmark specification. In contrast, in Table-3, Canada has an even split between the U-statistics supporting and against the oil-price model. It should be noted, however, since there are 7 Theil's U-ratios above unity, such results are still in favor of the random walk model. Germany is the only country among the three whose results actually switched sides. There are 16 U-ratios smaller than one for Germany in Table-3, in contrast to merely 4 in Table-1. The support Germany's results provide to the oil-price model is still quite weak though, because among those 16 statistics, only one directional U-ratio is significant.

The evaluation results when the oil-price model is estimated as a VAR in first-order difference are shown in Table-4, which seem to be weaker than those in Tables-1 and 3. In this table, Canada becomes the only country where the oil-price model edges out the random walk model. Even for Canada, if we take a closer look at its results, we can see that the advantage of the oil-price model is actually extremely small, because there are 6 Theil's U-statistics above one, and several MAE U-statistics below unity are actually very close to one. For all the other countries except the US, the forecasts based on VAR in first-order difference are worse than the benchmark. Especially, Japan, Norway and Mexico, the three countries whose benchmark results supported the oil-price model, are now on the side of the random walk model. For the US, some improvements were brought about the new specification, but not sufficient to change it to a case supporting the oil-price model.

Two points are worth noticing in summarizing this subsection. The first one is that there is no single specification that suits best for all countries. The overall results based on the specification of VAR in levels appear to be the strongest among all three alternatives we tried. For this reason, we will stick to it in the rest of the paper. The second point is that even with the best specification for each individual country chosen for that country, the general evidence is still too weak to provide convincing supports the claim that oil prices can help to forecast exchange rates.

3.3 Bad forecasts on oil prices?

The out-of-sample forecasts we have examined so far are all *true* dynamic forecasts, which contain no *ex post* information at all. In other words, we project not only future exchange rates but also future oil prices. This may lead to a question about the root of the poor forecasts on exchange rates by the oil-price model. Is it really because oil prices lack predictive power on exchange rates, or is it actually because the model cannot produce good forecasts on future oil prices, which in turn makes the forecasts on exchange rates bad? Table-5 provides justifications to such concerns. In this table forecasts on oil prices by the oil-price exchange rate model and those by the random walk model are compared to each other.[10] It is apparent that the former are generally not as accurate as the latter. This suggests that from the perspective of theoretical model testing, we might want to project the out-of-sample exchange rates for the oil-price model using *ex post* oil price information. If such forecasts are good, it then implies that oil prices are actually helpful to forecasting exchange rates, and the true problem with the oil-price model is its bad forecasts on future oil prices.

As discussed in the previous subsection, we will continue to estimate the oil-price exchange rate model as a VAR in levels. It is worth noting that since we do not forecast the out-of-sample oil prices, an equivalent way to proceed is to estimate a single equation OLS of exchange rates, on contemporary oil prices and the lags of both variables, and then forecast them out recursively using the actual out-of-sample oil prices.

Does the *ex post* oil price information help? The evaluation outcomes are shown in Table-6. The most interesting cases in this table are Japan and Norway, whose benchmark results were slightly in favor of the oil-price model. We can see that replacing projected oil prices with actual ones brings further improvements to the results for these two countries. The numbers of U-statistics below unity for both countries are now 23, as against 14 and 12 in Table-1, which provides additional supports to the oil-price exchange rate model. On the other hand, the outcome of Mexico tells a completely different story. Against the expectation that using *ex post* oil prices should improve the model's forecast performance,

[10] For a related study on this topic, please see Chen, Rogoff and Rossi (2009).

it makes Mexico one of the strongest examples against the oil-price model. In Table-6, 22 of Mexico's 24 U-statistics are now above one, including all 8 Theil's U-ratios. For the other countries, there are no significant changes in the patterns of their evaluation results. Nonetheless, this means that the benefits of hindsight did not translate into better forecast performance, which should be interpreted as a negative sign for the oil-price exchange rate model.

What we see in this subsection is a division among the sample countries. For most of them, *ex post* oil prices did not bring the expected improvements to the forecasts of the oil-price exchange rate model, casting further doubts on theory of Amano and van Norden (1998). Two countries, Japan and Norway, are different. The improvements in their evaluation results are consistent with the conjectures that oil prices can help to forecast exchange rates and the dynamic exchange rate forecasts of the model are affected by its poor projections on future oil prices.

3.4 Nonlinearity?

If there exists nonlinearity in the relations between oil prices and exchange rates, ignoring it may cause bias in the estimation and forecast results. In this subsection, we will explore such possibility by including nonlinear regressors in the estimations. More specifically, we will add the second and third order terms of the original regressors to the estimation equations. Like in the previous subsection, we will use *ex post* oil prices for forecasting, which again makes the estimation of a VAR model equivalent to a single equation OLS.
The results are reported in Table-7. Compared to Table-1, the patterns of changes are very clear and similar across all of the sample countries. That is, introducing higher order terms of the regressors either brought little improvements to the forecast performance of the oil-price exchange rate model, or actually made things even worse. Although the evidence in this table cannot entirely exclude the possibility of nonlinearity in the relations between oil prices and exchange rates, it does show that attempts to capture the nonlinearity between oil prices and exchange rates, if there is any, in simple ways as used here can hardly bring any gains to the forecast performance of the oil-price exchange rate model.

3.5 Omitted variables?

All the specifications of the oil-price exchange rate model we have estimated so far are essentially two-factor systems, oil prices and exchange rates. Some other important determinants of exchange rates have been ignored. Is such simplification causing the model to underperform? To provide some answers to this question, we will add additional control variables to the oil-price exchange rate model and re-evaluate its forecast performance.

Specifically, we will try to control two additional factors in the new models. One is the industrial output growth of the economies, and the other is their export growth. We experimented different ways to include these factors into the models. For instance, we first tried to add them individually, then we also estimated the models with both factors added simultaneously. Yet, reported in Tables-8 and 9 are the evaluation results based on two other specifications, both trying to capture the Balassa-Samuelson effect. The models in Table-8 include only one additional regressor, which is equal to the difference between the export growth and industrial output growth for the country under study minus the same difference for the US.[11] In this specification, export growth can be thought of as a proxy for the productivity growth in the tradable sectors, while industrial output growth is a proxy for the nontradable sectors. On the other hand, the models in Table-9 include four more regressors on the basis of the benchmark, export growth and industrial output growth of the country under study and those of the US. These models can be thought of as unrestricted versions of those in Table-8.

Although the Balassa-Samuelson effect has received wide empirical support in the literature, the regressors trying to capture it do not seem to bring much improvement to the forecast performance of the oil-price exchange rate model. In Table-8, with the exception of Mexico, all countries' results are qualitative the same as those in Table-1. For Mexico, once again, the results turned against the oil-price model. The evaluation outcomes in Table-9 are worse than those in Table-8 for most countries. In particular, Japan and Norway's results become against the oil-price model. The only country whose outcome improved slightly under the unrestricted specification is Turkey, but the change is too small to have any qualitative impact. One may wonder how the unrestricted models could produce worse forecasts than the restricted ones. Part of the answer is that, unlike in-sample fit, there is no guarantee in the first place that the out-of-sample forecast results of unrestricted models would be better. In the case of our VAR models, the unrestricted models have a heavier burden of forecasts. They need to project the future paths for six variables, as compared to three for the restricted models.

Ideally, in evaluating the predictive power of oil prices, we would like to control other factors that have influence on exchange rates. However, what this subsection shows is that, the simplification of ignoring other potential determinants of exchange rates does not necessarily affect the evaluation results of the oil-price exchange rate model in a negative way.

3.6 Terms of trade?

In explaining why domestic real oil prices could help to forecast a country's real exchange rates, Amano and van Norden (1998) suggested terms of trade as the linking channel. On the one hand, terms of trade would affect a country's real exchange rates; on the other hand, they showed that domestic real oil prices and terms of trade were highly correlated for the countries in their study.

A natural question to ask then is why Amano and van Norden (1998) did not focus on terms of trade, the factor directly affecting exchange rates. Their explanation was that

[11] For the US, Japan's export and industrial output growth are used instead.

compared to terms of trade, shocks to oil prices were more exogenous to the economies.[12] Although exogeneity of explanatory variables is important for structural analysis, it is not very critical if our primary goal is to forecast by estimating reduced form VAR or VEC models. Thus, we will make a short detour in this subsection and check how helpful terms of trade are in forecasting a country's real exchange rates.

Since monthly terms of trade information is not available for Norway and Mexico, we will study only six sample countries here. Their oil price and terms of trade (inverse) series are plotted in Figure-2. As found by Amano and van Norden (1998), domestic oil prices and real exchange rates are highly negatively correlated for the US, Japan and Germany. Their correlation coefficients are -0.65, -0.66, -0.83, respectively. Turkey has a similar high negative correlation between the two series. Its correlation coefficient is –0.59. The situations are different for the UK and Canada. Their correlation coefficients are 0.04 and 0.5, respectively.

We re-estimated the benchmark models with oil prices replaced by terms of trade. The results are reported Table-10. Similar to what we see in Table-1, there is little evidence showing that terms of trade can help to forecast exchange rates. The best results are those for Japan, with 19 U-statistics below unity. Yet again, none of them is statistically significant. The only other country whose results are in slight favor of the terms-of-trade model is the US. Although its total number of U-statistics below unity is just 10, they include more than half of the Theil's ratios. But like Japan, none of the statistics below unity is significant.

The results of this subsection cast more doubts on the theory proposed by Amano and van Norden (1998). Terms of trade, the linking channel between oil prices and exchange rates, do not seem to have high predictive power on exchange rates themselves. This probably explains why, although oil prices and terms of trade have maintained high correlations for some of our sample countries, we could not find strong evidence showing that oil prices are helpful to forecasting exchange rates.[13]

3.7 Stability?

Finally, we are going to look at the subsample results. There are two main purposes of the subsample analysis. First, the forecast period we have examined covers a range of more than twenty years. Over such a long of period time, structural breaks in the relations between oil prices and exchange rates become very natural concerns. By looking at the subsample results, we would like to test the stability of such relations. The second purpose

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[12] The exogeneity assumption on oil prices may no longer be valid. Please see, for example, Caballero, Farhi and Gourinchas (2008)

[13] In analysis not reported here, we also tried to replace the projected terms of trade with *ex post* information. The evaluation results did not change significantly.

is to check the claim by Amano and van Norden (1998) that the oil-price exchange rate model would forecast better when oil prices were more volatile.

The entire forecast period (1988M1-2009M7) is divided into four intervals. The first three each cover five years, 1988M1-92M12, 1993M1-97M12 and 1998M1-2002M12. The last interval is a bit wider, covering 2003M1 to 2009M7. The cut of subsample periods is largely arbitrary, in the sense that it is not based on significant economic or political events, for instance, the reunion of Germany and the introduction of the Euro. The rationale behind this is the assumption that the structural breaks in the relations between oil prices and exchange rates, if exist, take place gradually. In addition, this way of cutting sample involves little subjective judgments, thus helps to prevent data mining. Instead of using all the information available prior to the forecasting points, we adopt a moving window for the estimations and forecasts in this subsection. For example, for the US, the 1-month ahead forecast for 1988M1 is based on the estimation using data from 1973M1 to1987M12, while the same forecast for 1998M1 is based on information of the period 1983M1-1997M12. The purpose of using moving-windows is also to reduce the impact of potential structural breaks on our forecasts.

There are some interesting patterns in the subsample results, which are reported in Tables-11.1-11.8. First to notice is that there is no consistent good performance over the subsample periods by the oil-price exchange rate model for any of the sample countries. For Japan, the oil-price model beats the random walk model in the first three periods. In fact, its advantage is quite big in the first period, where we see five Theil's ratios significantly below unity. However, the results for the last subsample period shows exactly the opposite, where all Theil's ratios are above one and 7 of them are statistically significant. This clearly suggests that some structural breaks must have taken place between the last two subsample periods.[14]

For the two oil-exporting countries, Norway and Mexico, the oil-price exchange rate model performed particularly well in the period 1993M1-97M1. For instance, in the case of Norway, 22 of the U-statistics in this period are smaller than one, and more than half of them are significant. However, for both countries, this good performing period for the oil-price model actually comes in between two periods during which it performed really badly. Again, if we look at Norway's results for 1998M1-2002M12, things turned completely around from the previous period. All of the U-statistics are now larger than one, and more than half of them are significant. These results indicate that the relations between oil prices and exchange rates in these two countries are quite unstable over time.

Fluctuating patterns can also be seen in the subsample results of Germany, the UK, Canada and Turkey. For instance, in the case of Turkey, the oil-price model performed better than

[14] One caveat is that what we look at is the performance of the oil-price exchange rate model relative to the random walk model. There is the possibility that it is the variations in the performance of the random walk model, rather than the oil-price model, that leads to the pattern we see in the results.

the random walk model in the first and third subsample periods, but underperformed in the second and last ones. The only country showing a consistent pattern in the subsample results is the US. But in this case, it is the random walk model who outperformed the oil-price model consistently.

The over time variations in the performance of the oil-price exchange rate model allow us to verify whether it forecasts better when oil prices are more volatile. The oil price volatilities, the volatility rank and the performance rank of the oil-price model for each subsample period are reported in Table-12.[15] Better forecasts in periods with higher oil price volatilities is not exactly the pattern we see in Table-12. For example, the model shows the best performance in the periods with the lowest volatilities in oil prices for the US and Germany. In the case of Japan, the model has the worse results when the oil prices are most volatile. In fact, if we calculate the correlation between the oil price volatility rank and the model's performance rank, the coefficient is -0.15. The conjecture that the oil-price model performs better when oil prices are more volatile finds little support in our results.

The key findings of this subsection are the following. First, the relations between oil prices and exchange rates are not stable over time. The forecast performance of the oil-price model differs substantially from one period to another. With the exception of Japan, where the oil-price model outperforms the random walk model in three subsample periods, the oil-price model is outperformed by the random walk model in most subsample periods for other countries. Secondly, the conjecture that the oil-price model can produce better forecasts on exchange rates when oil prices are more volatile does not seem to hold in our subsample results.

4. Summary

This paper empirically tests the proposition proposed by Amano and van Norden (1998) that domestic real oil prices can help to forecast real effective exchange rates. The test is carried out through out-of-sample forecast evaluations for a selected group of countries. The benchmark specification of the oil-price exchange rate model in this paper is VAR in levels, and the robustness of the results is checked by estimating two alternative specifications, VEC and VAR in first-order difference. In the evaluation process, we examined not only dynamic forecasts, which contained no *ex post* information at all, but also those based on actual out-of-sample oil prices. The study on the second type of forecasts takes into consideration that the oil-price model might not be able to produce good forecasts on oil prices, and is to test the predictive power of oil prices on exchange rates from a more theoretical point of view. We explored the potential nonlinearity in the relations between oil prices and exchange rates by including higher order terms of the regressors. The sensitivity of the results to the issue of omitted variables was also addressed by including additional control variables in the model. In particular, we reported

[15] The performance rank of the oil-price exchange rate model is based on the number of U-statistics below unity and how many of them are significant.

in the paper the evaluation outcomes for two specifications intended to capture the Balassa-Samuelson effect, one in restricted form and the other in unrestricted form. Since Amano and van Norden (1998) suggested terms of trade as the linking channel between oil prices and exchange rates, we also had a quick check on their predictive power. Finally, through the subsample analysis, we tested the stability of the relations between oil prices and exchange rates, and verified whether the oil-price model could forecast better when oil prices were more volatile.

The overall message conveyed by this paper is a negative one. For most countries in our sample, little evidence was found to support the oil-price exchange rate model as a better forecasting tool than the random walk model. There were two exceptions, Japan and Norway, where we did see some consistent signs in favor of the oil-price model. However, even for these two countries, the statistics favoring the oil-price model were mostly insignificant, making the supports extremely weak.

In the subsample analysis, we found that the forecast performance of the oil-price exchange rate model differed substantially from one period to another. Such results suggest that the relations between oil prices and exchange rates are not stable over time. The conjecture that the oil-price model can produce better forecasts on exchange rates when oil prices are more volatile is not consistent with the pattern seen in our subsample results.

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Data Appendix

A. Data sources

1. MorganMarkets database, JPMorgan.
2. International Financial Statistics, International Monetary Fund.
3. Global Financial Database (GFD), Global Financial Data, Inc.
4. Author's calculations.

B. Summary statistics

		Data range	Obs.	Mean	Std. Dev.	Min	Max
US	REER	1973M1-2009M7	439	91.4	8.9	76.0	119.2
	Real oil prices	1973M1-2009M7	439	42.4	19.5	13.5	124.9
	TOT	1973M1-2009M7	439	105.7	6.3	89.2	131.9
	Industrial production growth	1973M1-2009M7	439	2.2	4.6	-13.3	12.2
	Export growth	1973M1-2009M7	439	5.2	8.7	-21.0	30.5
Japan	REER	1973M1-2009M7	439	88.5	10.4	66.7	118.5
	Real oil prices	1973M1-2009M7	439	4988	2850	1352	14574
	TOT	1973M1-2009M7	439	122.0	23.8	67.3	201.9
	Industrial production growth	1973M1-2009M7	439	1.8	7.2	-35.7	17.1
	Export growth	1973M1-2009M7	439	6.4	10.3	-46.9	38.3
Germany	REER	1975M1-2009M7	415	107.9	8.8	91.1	133.5
	Real oil prices	1975M1-2009M7	415	69.4	33.2	21.3	162.3
	TOT	1975M1-2009M6	414	98.9	5.9	82.3	109.4
	Industrial production growth	1975M1-2009M7	415	1.4	5.2	-24.6	17.3
	Export growth	1975M1-2009M6	414	5.2	8.6	-27.3	34.1
Turkey	REER	1973M1-2009M7	439	115.1	20.5	68.0	217.4
	Real oil prices	1973M1-2009M7	439	69.9	28.6	22.5	140.3
	TOT	1982M1-2009M7	331	106.1	9.1	87.3	130.3
	Industrial production growth	1986M1-2009M7	283	4.2	8.6	-23.8	22.1
	Export growth	1983M1-2009M7	319	11.4	14.3	-24.3	63.6
UK	REER	1973M1-2009M7	439	88.7	11.1	61.8	104.4
	Real oil prices	1973M1-2009M7	439	27.1	11.8	8.0	62.2
	TOT	1973M1-2009M7	439	95.6	4.9	78.0	103.6
	Industrial production growth	1973M1-2009M7	439	0.9	4.3	-12.8	22.6
	Export growth	1973M1-2009M7	439	4.5	10.0	-22.8	89.0
Canada	REER	1973M1-2009M7	439	115.0	12.5	86.6	143.4
	Real oil prices	1973M1-2009M7	439	50.3	21.5	16.4	132.2
	TOT	1973M1-2009M7	439	92.8	5.2	84.8	110.5
	Industrial production growth	1973M1-2009M7	439	2.6	4.8	-15.1	19.0
	Export growth	1973M1-2009M7	439	3.9	9.2	-25.7	30.8
Mexico	REER	1975M1-2009M7	439	96.0	15.1	54.2	134.1
	Real oil prices	1975M1-2009M7	439	501.6	218.6	173.3	1281.5
	Industrial production growth	1981M1-2009M6	342	2.3	5.9	-14.7	14.2
	Export growth	1975M1-2009M7	439	-5.6	25.5	-62.1	87.5
Norway	REER	1973M1-2009M7	439	96.0	15.1	54.2	134.1
	Real oil prices	1973M1-2009M7	439	282.0	126.4	97.9	669.4
	Industrial production growth	1973M1-2009M7	439	3.4	6.3	-29.0	53.9
	Export growth	1973M1-2009M7	439	7.2	20.6	-50.0	95.2

Figure-1. Domestic Real Oil Prices and Real Effective Exchange Rates

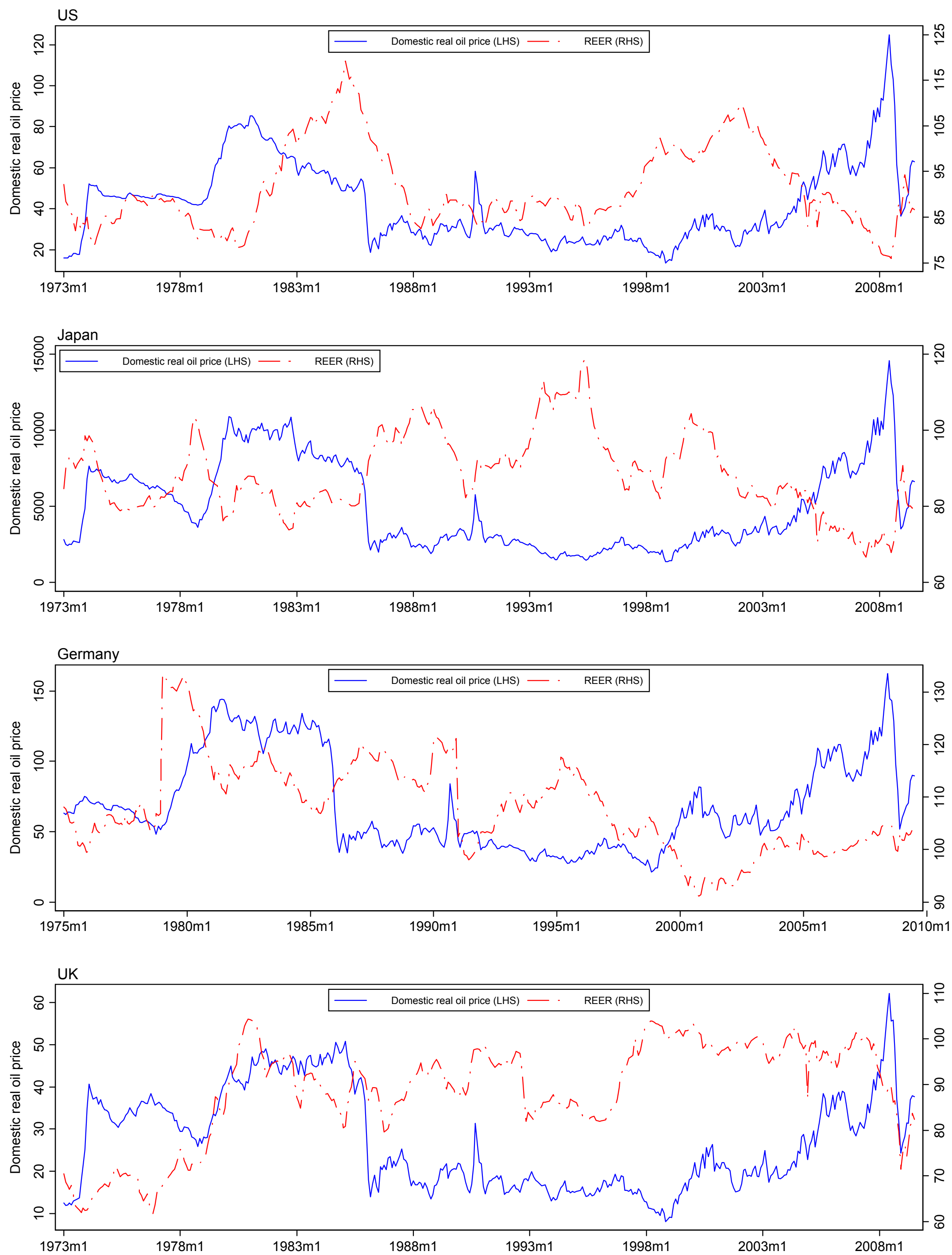

Figure-1. Domestic Real Oil Prices and Real Effective Exchange Rates (cont.)

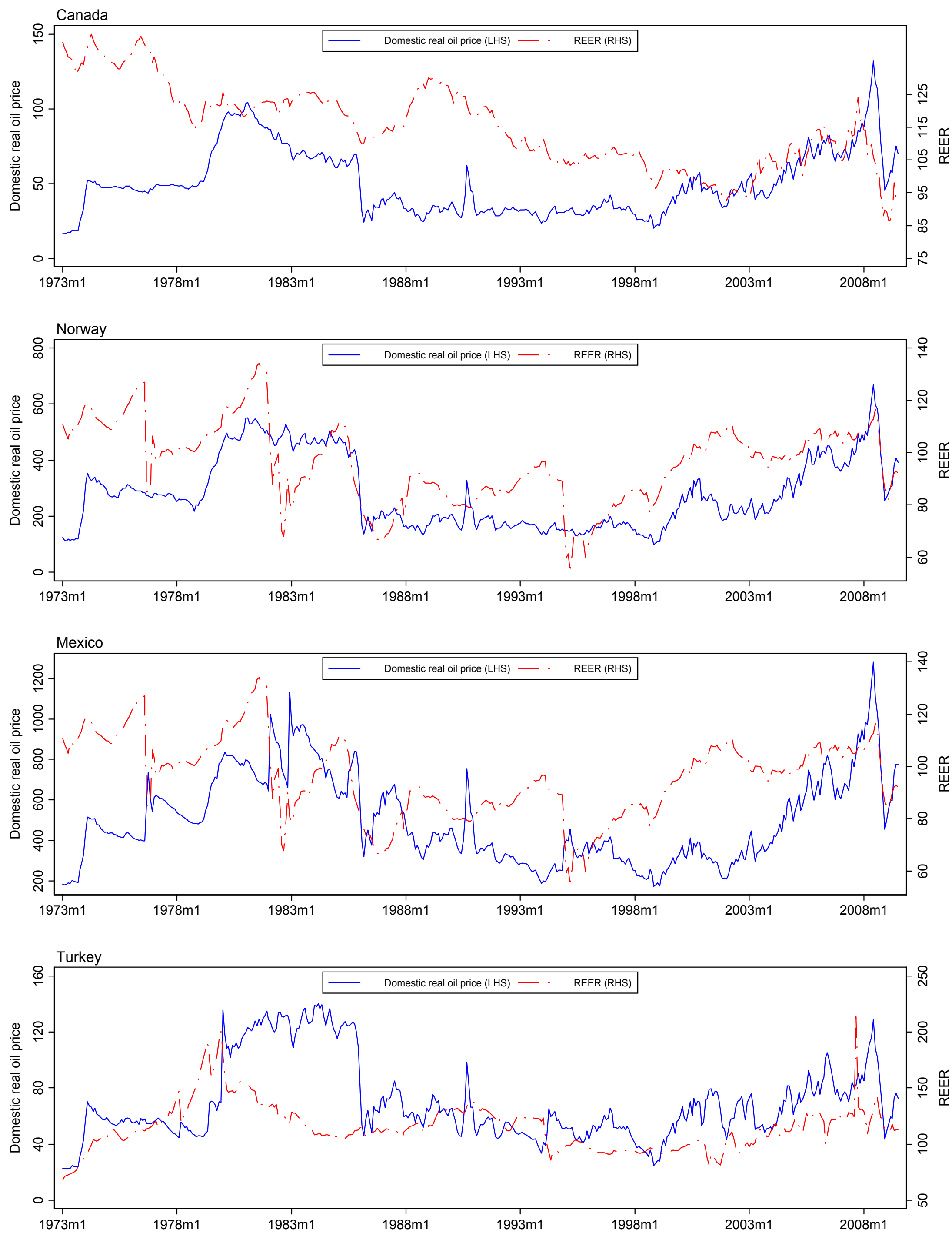

Figure-2. Domestic Real Oil Prices and Terms of Trade

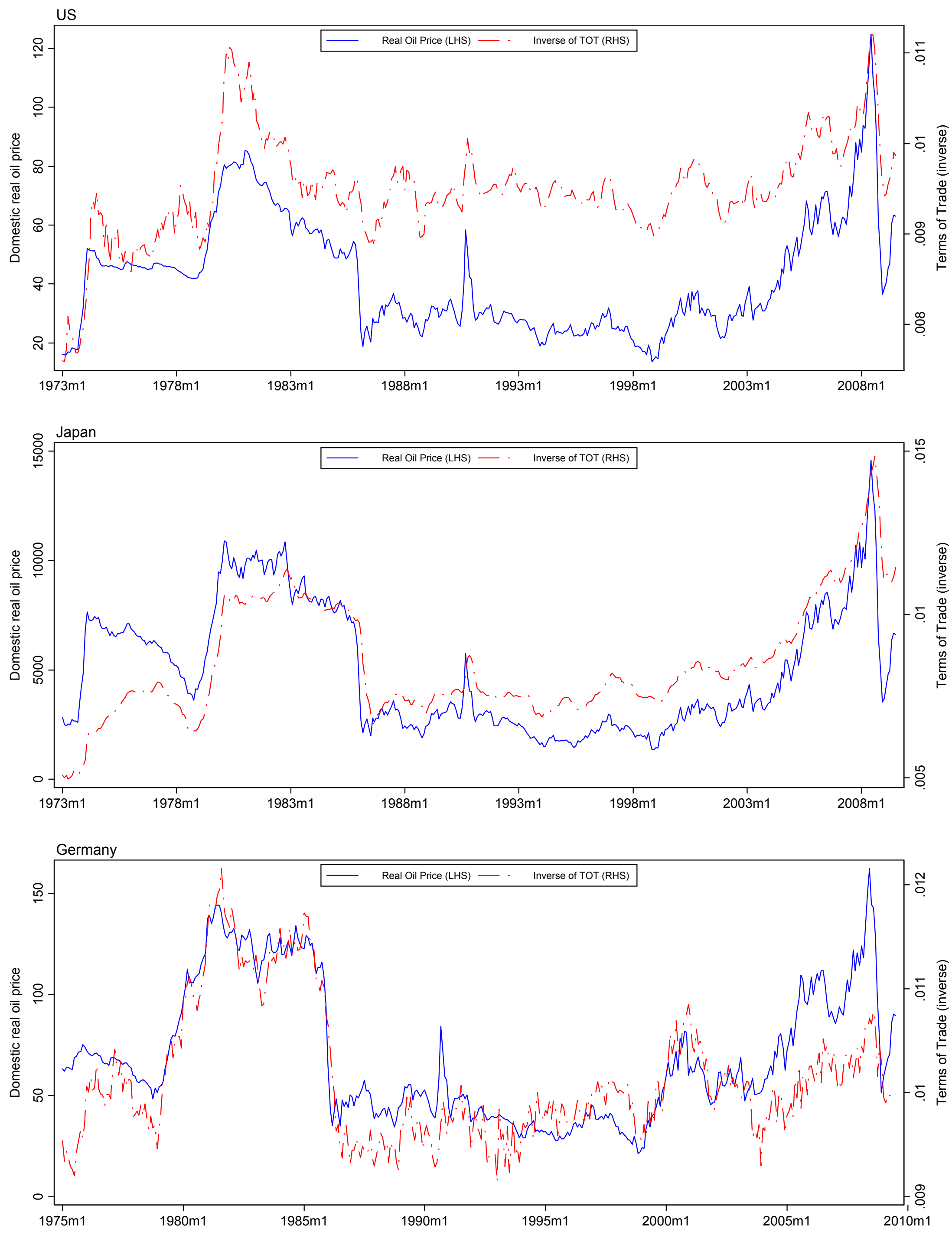

Figure-2. Domestic Real Oil Prices and Terms of Trade (cont.)

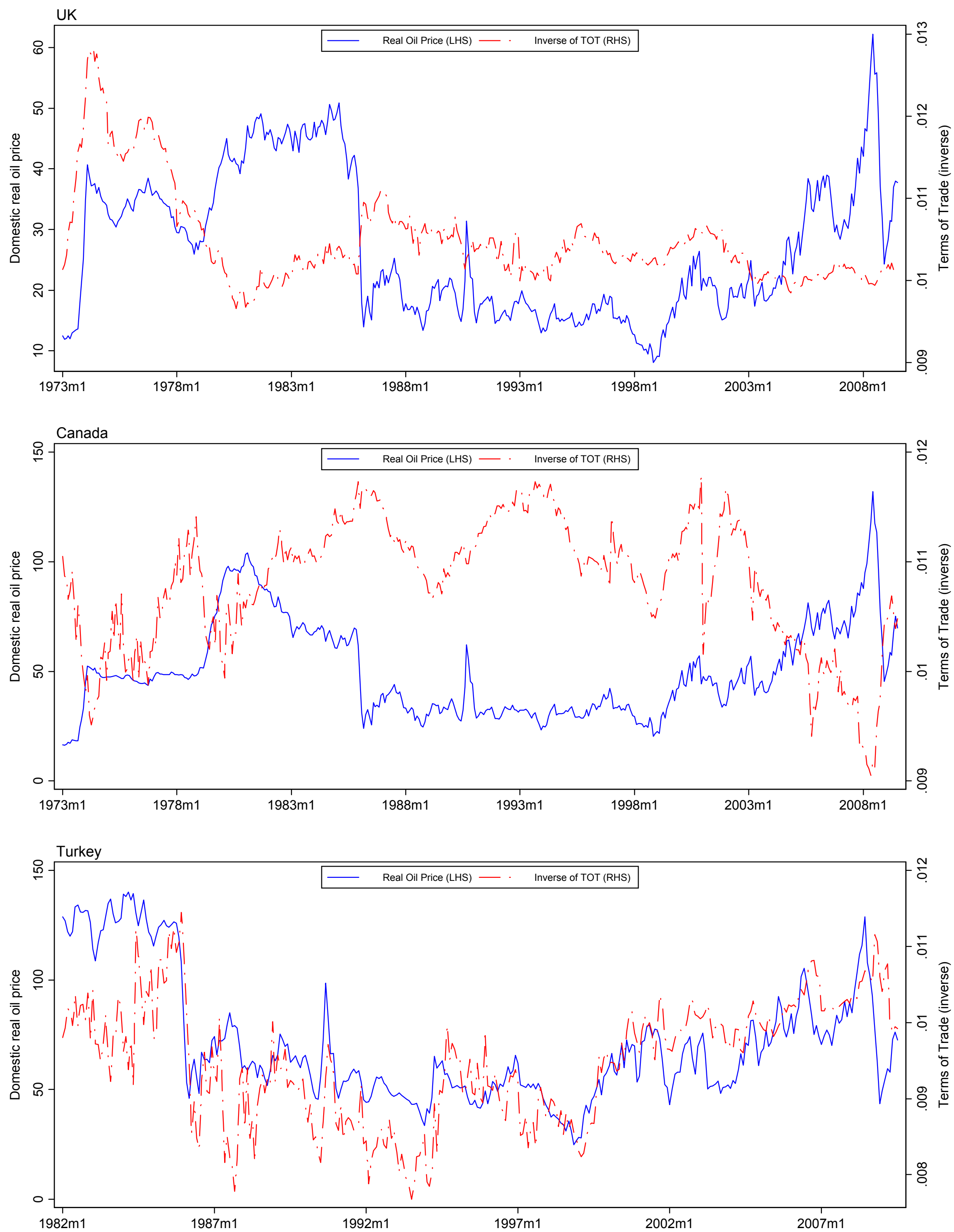

Table - 1. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Full Sample, Based on Schwarz's Bayesian Information Criterion)

Forecast horizon	US			Japan			Germany			UK		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.989	0.948	1.131	0.978	0.966	1.065	1.018	1.064	1.228	1.079	1.001	1.019
	(0.77)	(0.14)	(0.01)	(0.61)	(0.38)	(0.2)	(0.13)	(0.02)	(0.02)	(0.39)	(0.98)	(0.55)
2	1.009	0.995	1.079	0.995	1.011	0.993	1.032	1.082	1.107	1.055	1.015	1.027
	(0.82)	(0.9)	(0.23)	(0.9)	(0.79)	(0.91)	(0.16)	(0.04)	(0.24)	(0.41)	(0.63)	(0.57)
3	1.012	1.022	1.075	1.003	1.031	0.993	1.040	1.073	1.135	1.038	1.017	1.021
	(0.8)	(0.62)	(0.35)	(0.94)	(0.45)	(0.93)	(0.2)	(0.2)	(0.13)	(0.44)	(0.54)	(0.71)
6	1.062	1.106	1.236	1.029	1.083	1.000	1.048	1.061	0.938	1.019	1.030	0.965
	(0.35)	(0.12)	(0.03)	(0.54)	(0.19)	(1)	(0.34)	(0.48)	(0.45)	(0.45)	(0.32)	(0.61)
9	1.095	1.141	1.259	1.014	1.069	0.993	1.057	1.057	0.960	1.015	1.029	0.952
	(0.25)	(0.12)	(0.01)	(0.83)	(0.45)	(0.93)	(0.37)	(0.61)	(0.61)	(0.62)	(0.43)	(0.51)
12	1.122	1.184	1.264	0.984	1.046	0.916	1.055	1.052	0.979	1.012	1.010	0.926
	(0.18)	(0.06)	(0.02)	(0.84)	(0.66)	(0.29)	(0.45)	(0.66)	(0.8)	(0.74)	(0.8)	(0.3)
18	1.162	1.208	1.200	0.916	0.973	0.849	1.080	1.042	0.965	0.992	1.005	0.825
	(0.13)	(0.08)	(0.07)	(0.37)	(0.81)	(0.04)	(0.4)	(0.76)	(0.66)	(0.86)	(0.94)	(0.01)
24	1.169	1.227	1.314	0.887	0.906	0.898	1.141	1.038	1.060	0.992	1.044	0.884
	(0.17)	(0.09)	(0)	(0.26)	(0.41)	(0.18)	(0.25)	(0.81)	(0.47)	(0.88)	(0.61)	(0.13)
No. of Stat. < 1	1	2	0	5	3	6	0	0	4	2	0	5
(of which sig.)	0	0	0	0	0	1	0	0	0	0	0	1
No. of Stat. >= 1	7	6	8	3	5	2	8	8	4	6	8	3
(of which sig.)	0	3	6	0	0	0	0	2	1	0	0	0

Forecast horizon	Canada			Norway			Mexico			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.013	0.977	1.000	0.936	0.960	1.071	0.990	1.056	1.250	1.111	1.068	1.170
	(0.48)	(0.31)	(1)	(0.13)	(0.38)	(0.2)	(0.83)	(0.27)	(0)	(0.28)	(0.36)	(0.01)
2	1.017	1.006	1.007	0.957	0.993	1.060	0.986	1.051	1.197	1.072	1.053	1.144
	(0.24)	(0.76)	(0.87)	(0.28)	(0.89)	(0.35)	(0.78)	(0.41)	(0.02)	(0.27)	(0.19)	(0.06)
3	1.012	1.005	0.978	0.970	0.996	1.027	0.998	1.053	1.214	1.065	1.074	1.109
	(0.41)	(0.74)	(0.67)	(0.39)	(0.95)	(0.68)	(0.95)	(0.34)	(0.02)	(0.2)	(0.08)	(0.17)
6	1.023	1.014	1.081	1.012	1.026	1.087	0.998	1.014	1.113	1.039	1.081	0.935
	(0.15)	(0.44)	(0.23)	(0.75)	(0.75)	(0.31)	(0.93)	(0.8)	(0.22)	(0.14)	(0.19)	(0.41)
9	1.035	1.058	1.036	1.028	1.017	1.108	1.023	1.021	1.125	1.033	1.088	0.884
	(0.06)	(0.03)	(0.67)	(0.59)	(0.87)	(0.21)	(0.57)	(0.74)	(0.17)	(0.45)	(0.34)	(0.13)
12	1.040	1.060	1.229	1.012	0.987	1.046	1.012	0.988	1.070	1.055	1.120	0.863
	(0.07)	(0.07)	(0.01)	(0.83)	(0.9)	(0.59)	(0.79)	(0.86)	(0.43)	(0.37)	(0.31)	(0.06)
18	1.055	1.062	1.194	0.991	0.953	1.014	0.993	0.948	1.061	1.086	1.177	0.808
	(0.08)	(0.15)	(0.04)	(0.9)	(0.68)	(0.86)	(0.89)	(0.5)	(0.48)	(0.36)	(0.29)	(0.01)
24	1.067	1.076	1.330	0.969	0.910	0.978	0.973	0.929	1.015	1.213	1.280	1.107
	(0.09)	(0.13)	(0)	(0.71)	(0.48)	(0.78)	(0.64)	(0.4)	(0.85)	(0.11)	(0.17)	(0.22)
No. of Stat. < 1	0	1	1	5	6	1	6	3	0	0	0	4
(of which sig.)	0	0	0	0	0	0	0	0	0	0	0	2
No. of Stat. >= 1	8	7	7	3	2	7	2	5	8	8	8	4
(of which sig.)	4	2	3	0	0	0	0	0	3	0	1	2

Table - 2. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Full Sample, Based on Likelihood Ratio Statistics)

Forecast horizon	US			Japan			Germany			UK		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.027	0.990	1.116	1.022	0.996	1.065	1.026	1.076	1.208	1.146	1.115	1.061
	(0.53)	(0.84)	(0.03)	(0.63)	(0.93)	(0.26)	(0.06)	(0.01)	(0.02)	(0.14)	(0.08)	(0.26)
2	1.036	1.025	1.027	1.020	1.019	0.946	1.051	1.084	1.090	1.119	1.104	1.013
	(0.42)	(0.61)	(0.66)	(0.64)	(0.7)	(0.37)	(0.09)	(0.04)	(0.29)	(0.09)	(0.02)	(0.81)
3	1.032	1.034	1.075	1.029	1.038	0.925	1.057	1.084	1.162	1.124	1.104	1.007
	(0.53)	(0.56)	(0.32)	(0.52)	(0.5)	(0.25)	(0.11)	(0.12)	(0.07)	(0.06)	(0.04)	(0.91)
6	1.088	1.124	1.248	1.085	1.141	0.912	1.030	1.049	0.900	1.088	1.105	1.038
	(0.2)	(0.14)	(0.02)	(0.23)	(0.12)	(0.25)	(0.58)	(0.57)	(0.19)	(0.08)	(0.04)	(0.63)
9	1.136	1.161	1.369	1.055	1.125	0.944	1.029	1.046	0.923	1.078	1.094	1.120
	(0.12)	(0.16)	(0)	(0.56)	(0.32)	(0.46)	(0.68)	(0.69)	(0.3)	(0.17)	(0.14)	(0.15)
12	1.177	1.213	1.376	1.026	1.091	0.873	1.025	1.046	0.927	1.069	1.079	1.132
	(0.12)	(0.11)	(0)	(0.83)	(0.54)	(0.08)	(0.77)	(0.71)	(0.34)	(0.28)	(0.25)	(0.15)
18	1.269	1.261	1.156	0.949	0.983	0.832	1.039	1.026	0.879	1.034	1.085	1.134
	(0.1)	(0.15)	(0.14)	(0.71)	(0.92)	(0.02)	(0.71)	(0.86)	(0.08)	(0.7)	(0.4)	(0.18)
24	1.310	1.342	1.229	0.902	0.873	0.904	1.087	1.019	1.007	1.025	1.088	1.057
	(0.11)	(0.11)	(0.03)	(0.49)	(0.45)	(0.2)	(0.48)	(0.9)	(0.93)	(0.81)	(0.41)	(0.5)
No. of Stat. < 1	0	1	0	2	3	7	0	0	4	0	0	0
(of which sig.)	0	0	0	0	0	2	0	0	1	0	0	0
No. of Stat. >= 1	8	7	8	6	5	1	8	8	4	8	8	8
(of which sig.)	1	0	5	0	0	0	2	2	2	3	4	0

Forecast horizon	Canada			Norway			Mexico			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.036	1.020	1.061	1.070	1.113	1.170	1.056	1.124	1.170	1.177	1.161	1.204
	(0.29)	(0.49)	(0.3)	(0.37)	(0.11)	(0.01)	(0.44)	(0.06)	(0.01)	(0.26)	(0.16)	(0.01)
2	1.029	1.037	1.029	1.068	1.130	1.053	1.048	1.104	1.082	1.092	1.104	1.063
	(0.21)	(0.16)	(0.65)	(0.21)	(0.08)	(0.42)	(0.42)	(0.11)	(0.2)	(0.2)	(0.02)	(0.38)
3	1.035	1.028	1.023	1.083	1.115	1.020	1.060	1.088	1.027	1.083	1.143	1.060
	(0.22)	(0.34)	(0.75)	(0.05)	(0.11)	(0.78)	(0.14)	(0.1)	(0.7)	(0.13)	(0.02)	(0.46)
6	1.059	1.059	1.039	1.072	1.085	0.986	1.045	1.066	1.015	1.020	1.109	0.928
	(0.06)	(0.06)	(0.62)	(0.09)	(0.36)	(0.84)	(0.18)	(0.39)	(0.84)	(0.7)	(0.18)	(0.39)
9	1.073	1.107	1.074	1.076	1.083	1.000	1.047	1.068	1.075	1.085	1.155	0.884
	(0.03)	(0.01)	(0.44)	(0.25)	(0.48)	(1)	(0.38)	(0.5)	(0.35)	(0.16)	(0.12)	(0.13)
12	1.078	1.115	1.136	1.078	1.072	1.007	1.041	1.054	1.054	1.128	1.199	0.875
	(0.03)	(0.03)	(0.11)	(0.33)	(0.56)	(0.92)	(0.53)	(0.62)	(0.51)	(0.12)	(0.13)	(0.09)
18	1.119	1.138	1.173	1.081	1.068	1.007	1.022	1.026	1.022	1.186	1.316	0.847
	(0.03)	(0.06)	(0.07)	(0.43)	(0.63)	(0.93)	(0.79)	(0.83)	(0.79)	(0.1)	(0.08)	(0.04)
24	1.151	1.180	1.317	1.060	1.024	0.985	1.003	0.987	0.957	1.374	1.459	1.174
	(0.04)	(0.05)	(0)	(0.58)	(0.87)	(0.85)	(0.98)	(0.92)	(0.58)	(0.02)	(0.04)	(0.06)
No. of Stat. < 1	0	0	0	0	0	2	0	1	1	0	0	4
(of which sig.)	0	0	0	0	0	0	0	0	0	0	0	2
No. of Stat. >= 1	8	8	8	8	8	6	8	7	7	8	8	4
(of which sig.)	5	5	2	2	1	1	0	2	1	2	4	2

Table - 3. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VEC, Full Sample)

Forecast horizon	US			Japan			Germany			UK		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.998	0.949	1.075	0.981	0.971	1.072	1.002	1.012	1.119	1.078	0.997	1.013
	(0.95)	(0.12)	(0.12)	(0.66)	(0.42)	(0.15)	(0.86)	(0.55)	(0.16)	(0.4)	(0.94)	(0.62)
2	1.024	0.995	1.034	1.001	1.022	1.037	1.001	1.017	1.058	1.052	1.007	0.969
	(0.51)	(0.87)	(0.58)	(0.98)	(0.56)	(0.59)	(0.95)	(0.55)	(0.48)	(0.45)	(0.84)	(0.34)
3	1.030	1.018	0.979	1.010	1.037	1.015	0.999	1.007	1.102	1.031	1.002	0.960
	(0.44)	(0.65)	(0.75)	(0.75)	(0.32)	(0.85)	(0.97)	(0.85)	(0.22)	(0.55)	(0.96)	(0.32)
6	1.101	1.110	1.097	1.041	1.077	1.042	0.986	0.957	0.888	1.007	1.012	0.945
	(0.04)	(0.07)	(0.25)	(0.33)	(0.16)	(0.66)	(0.72)	(0.46)	(0.13)	(0.85)	(0.76)	(0.31)
9	1.153	1.192	1.205	1.033	1.075	1.071	0.972	0.954	0.900	1.000	1.020	1.029
	(0.01)	(0.01)	(0.02)	(0.53)	(0.35)	(0.44)	(0.61)	(0.56)	(0.15)	(0.99)	(0.65)	(0.64)
12	1.208	1.275	1.311	1.012	1.066	0.985	0.955	0.956	0.908	1.006	1.038	1.015
	(0)	(0)	(0)	(0.85)	(0.46)	(0.86)	(0.49)	(0.59)	(0.21)	(0.9)	(0.42)	(0.83)
18	1.301	1.408	1.465	0.962	1.029	0.921	0.944	0.938	0.841	1.014	1.066	1.114
	(0)	(0)	(0)	(0.59)	(0.76)	(0.34)	(0.47)	(0.54)	(0.02)	(0.81)	(0.29)	(0.15)
24	1.364	1.532	1.740	0.945	0.982	0.923	0.958	0.923	0.959	1.012	1.038	1.024
	(0)	(0)	(0)	(0.45)	(0.85)	(0.34)	(0.64)	(0.5)	(0.59)	(0.86)	(0.6)	(0.72)
No. of Stat. < 1	1	2	1	3	2	3	6	5	5	0	1	3
(of which sig.)	0	0	0	0	0	0	0	0	1	0	0	0
No. of Stat. >= 1	7	6	7	5	6	5	2	3	3	8	7	5
(of which sig.)	5	5	4	0	0	0	0	0	0	0	0	0

Forecast horizon	Canada			Norway			Mexico			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.016	0.978	0.994	0.939	0.967	1.107	1.004	1.090	1.289	1.122	1.066	1.154
	(0.42)	(0.34)	(0.87)	(0.11)	(0.33)	(0.07)	(0.93)	(0.04)	(0)	(0.28)	(0.41)	(0.01)
2	1.023	1.003	0.941	0.960	0.977	1.075	1.009	1.085	1.153	1.087	1.051	1.135
	(0.21)	(0.89)	(0.17)	(0.2)	(0.56)	(0.25)	(0.83)	(0.14)	(0.06)	(0.27)	(0.23)	(0.05)
3	1.022	0.996	0.958	0.971	0.975	1.070	1.025	1.078	1.195	1.087	1.072	1.118
	(0.27)	(0.83)	(0.41)	(0.27)	(0.49)	(0.34)	(0.37)	(0.12)	(0.04)	(0.21)	(0.07)	(0.11)
6	1.034	1.008	1.000	1.002	0.977	1.000	1.027	1.042	1.104	1.076	1.054	0.985
	(0.21)	(0.66)	(1)	(0.96)	(0.65)	(1)	(0.38)	(0.38)	(0.27)	(0.14)	(0.15)	(0.84)
9	1.042	1.019	0.879	1.002	0.954	0.986	1.064	1.080	1.210	1.077	1.059	0.963
	(0.24)	(0.53)	(0.09)	(0.97)	(0.45)	(0.86)	(0.1)	(0.1)	(0.03)	(0.07)	(0.2)	(0.63)
12	1.043	1.002	0.950	0.975	0.945	0.938	1.061	1.072	1.181	1.101	1.087	0.992
	(0.34)	(0.97)	(0.49)	(0.69)	(0.41)	(0.42)	(0.2)	(0.18)	(0.07)	(0.09)	(0.18)	(0.92)
18	1.029	0.970	0.872	0.933	0.926	1.077	1.055	1.088	1.443	1.147	1.157	0.961
	(0.56)	(0.52)	(0.06)	(0.33)	(0.3)	(0.37)	(0.31)	(0.14)	(0)	(0.08)	(0.14)	(0.64)
24	0.993	0.958	0.943	0.916	0.896	0.993	1.055	1.081	1.467	1.225	1.239	1.227
	(0.89)	(0.45)	(0.42)	(0.23)	(0.16)	(0.93)	(0.37)	(0.21)	(0)	(0.11)	(0.1)	(0.02)
No. of Stat. < 1	1	4	7	6	8	3	0	0	0	0	0	4
(of which sig.)	0	0	2	0	0	0	0	0	0	0	0	0
No. of Stat. >= 1	7	4	1	2	0	5	8	8	8	8	8	4
(of which sig.)	0	0	0	0	0	1	1	2	7	3	2	3

Table - 4. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in First-order Difference, Full Sample)

Forecast horizon	US			Japan			Germany			UK		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.983	0.923	1.018	0.980	0.947	1.032	1.020	1.046	1.127	1.084	1.002	1.000
	(0.64)	(0.01)	(0.51)	(0.63)	(0.17)	(0.25)	(0.1)	(0.09)	(0.13)	(0.37)	(0.97)	(1)
2	0.994	0.965	0.980	0.995	0.989	1.007	1.039	1.035	1.107	1.065	1.016	0.987
	(0.86)	(0.22)	(0.56)	(0.89)	(0.78)	(0.83)	(0.17)	(0.12)	(0.21)	(0.35)	(0.64)	(0.64)
3	0.986	0.970	0.941	0.999	0.987	1.000	1.044	1.035	1.162	1.053	1.022	1.000
	(0.63)	(0.27)	(0.08)	(0.97)	(0.62)	(1)	(0.23)	(0.18)	(0.06)	(0.32)	(0.48)	(1)
6	1.008	0.990	0.986	1.021	1.028	0.977	1.010	1.015	1.164	1.045	1.041	1.087
	(0.64)	(0.63)	(0.72)	(0.17)	(0.14)	(0.6)	(0.44)	(0.32)	(0.08)	(0.18)	(0.28)	(0.14)
9	1.016	1.002	1.022	1.017	1.014	1.046	1.015	1.025	1.321	1.047	1.047	1.129
	(0.34)	(0.92)	(0.59)	(0.17)	(0.32)	(0.32)	(0.23)	(0.07)	(0)	(0.16)	(0.18)	(0.06)
12	1.020	1.017	1.053	1.018	1.020	1.110	1.030	1.028	1.230	1.059	1.076	1.245
	(0.25)	(0.41)	(0.21)	(0.15)	(0.17)	(0.05)	(0.06)	(0.04)	(0.02)	(0.11)	(0.04)	(0)
18	1.029	1.037	1.096	1.031	1.021	1.183	1.046	1.043	1.353	1.076	1.131	1.270
	(0.14)	(0.06)	(0.1)	(0.04)	(0.2)	(0.01)	(0.05)	(0.02)	(0)	(0.07)	(0.01)	(0.01)
24	1.031	1.038	1.218	1.041	1.024	1.211	1.058	1.043	1.484	1.094	1.143	1.240
	(0.19)	(0.09)	(0)	(0.04)	(0.15)	(0)	(0.03)	(0.07)	(0)	(0.06)	(0.02)	(0.01)
No. of Stat. < 1	3	4	3	3	3	1	0	0	0	0	0	1
(of which sig.)	0	1	1	0	0	0	0	0	0	0	0	0
No. of Stat. >= 1	5	4	5	5	5	7	8	8	8	8	8	7
(of which sig.)	0	2	2	2	0	3	4	5	6	2	3	4

Forecast horizon	Canada			Norway			Mexico			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.013	0.979	0.969	0.944	0.957	1.100	1.061	1.104	1.514	1.121	1.061	1.071
	(0.46)	(0.35)	(0.4)	(0.94)	(0.96)	(1.1)	(0.06)	(0)	(0)	(0.28)	(0.46)	(0.11)
2	1.016	0.994	0.935	0.968	0.967	1.090	1.049	1.113	1.306	1.091	1.042	1.102
	(0.3)	(0.79)	(0.11)	(0.97)	(0.97)	(1.09)	(0.01)	(0)	(0)	(0.27)	(0.35)	(0.07)
3	1.010	0.989	0.925	0.983	0.997	1.109	1.059	1.103	1.319	1.094	1.064	1.127
	(0.54)	(0.57)	(0.15)	(0.98)	(1)	(1.11)	(0.01)	(0)	(0.01)	(0.22)	(0.12)	(0.06)
6	1.006	0.999	0.924	1.017	1.021	1.122	1.010	1.034	1.095	1.087	1.052	1.008
	(0.69)	(0.97)	(0.21)	(1.02)	(1.02)	(1.12)	(0.48)	(0.18)	(0.36)	(0.2)	(0.05)	(0.92)
9	1.001	0.999	0.753	1.019	1.018	1.171	1.033	1.061	1.309	1.073	1.033	1.008
	(0.95)	(0.98)	(0)	(1.02)	(1.02)	(1.17)	(0.02)	(0.02)	(0.01)	(0.14)	(0.13)	(0.92)
12	0.998	0.990	0.950	1.026	1.038	1.305	1.031	1.055	1.269	1.089	1.050	1.033
	(0.93)	(0.82)	(0.5)	(1.03)	(1.04)	(1.3)	(0.08)	(0.03)	(0.02)	(0.13)	(0.15)	(0.68)
18	1.005	0.978	0.890	1.041	1.067	1.538	1.040	1.069	1.538	1.111	1.062	1.000
	(0.88)	(0.71)	(0.12)	(1.04)	(1.07)	(1.54)	(0.12)	(0.03)	(0)	(0.13)	(0.26)	(1)
24	0.992	0.957	0.875	1.059	1.088	1.875	1.051	1.084	1.776	1.069	1.080	1.274
	(0.89)	(0.56)	(0.07)	(1.06)	(1.09)	(1.88)	(0.13)	(0.02)	(0)	(0.24)	(0.25)	(0)
No. of Stat. < 1	2	8	8	3	3	0	0	0	0	0	0	0
(of which sig.)	0	0	2	0	0	0	0	0	0	0	0	0
No. of Stat. >= 1	6	0	0	5	5	8	8	8	8	8	8	8
(of which sig.)	0	0	0	0	0	0	5	7	7	0	1	3

Table - 5. Forecast Evaluation of the Oil-Price Exchange Rate Model

(Forecasts on Oil Prices, VAR in Levels, Full Sample)

Forecast horizon	US			Japan			Germany			UK		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.986	1.015	1.036	0.987	1.018	1.116	1.011	1.038	1.079	1.497	1.453	0.924
	(0.7)	(0.54)	(0.58)	(0.77)	(0.53)	(0.14)	(0.66)	(0.05)	(0.31)	(0.01)	(0)	(0.23)
2	0.999	1.054	0.843	0.984	1.046	0.979	1.018	1.058	0.929	1.004	1.016	0.860
	(0.97)	(0.03)	(0.01)	(0.72)	(0.12)	(0.8)	(0.55)	(0)	(0.32)	(0.66)	(0.23)	(0.03)
3	1.011	1.054	0.753	0.992	1.041	0.906	1.022	1.071	0.962	1.180	1.209	0.797
	(0.65)	(0.04)	(0)	(0.81)	(0.2)	(0.22)	(0.43)	(0.01)	(0.59)	(0.04)	(0)	(0)
6	1.060	1.127	0.731	1.020	1.093	0.906	1.044	1.100	0.877	1.007	1.054	0.733
	(0.01)	(0)	(0)	(0.26)	(0.05)	(0.28)	(0.15)	(0.01)	(0.08)	(0.75)	(0.04)	(0)
9	1.088	1.219	0.750	1.025	1.108	0.889	1.060	1.104	0.954	1.046	1.114	0.791
	(0)	(0)	(0)	(0.25)	(0.14)	(0.13)	(0.19)	(0.06)	(0.52)	(0.3)	(0.01)	(0)
12	1.108	1.273	0.720	1.030	1.132	0.834	1.084	1.112	0.903	1.033	1.066	0.714
	(0)	(0)	(0)	(0.31)	(0.13)	(0.03)	(0.11)	(0.06)	(0.16)	(0.37)	(0.19)	(0)
18	1.202	1.448	0.670	1.079	1.188	0.872	1.179	1.298	0.979	1.115	1.222	0.658
	(0)	(0)	(0)	(0.15)	(0.1)	(0.11)	(0)	(0)	(0.78)	(0.01)	(0.01)	(0)
24	1.324	1.574	0.757	1.125	1.216	0.863	1.299	1.469	1.128	1.215	1.389	0.713
	(0)	(0)	(0)	(0.13)	(0.12)	(0.07)	(0)	(0)	(0.14)	(0)	(0)	(0)
No. of Stat. < 1	2	0	7	3	0	7	0	0	6	0	0	8
(of which sig.)	0	0	7	0	0	2	0	0	1	0	0	7
No. of Stat. >= 1	6	8	1	5	8	1	8	8	2	8	8	0
(of which sig.)	5	7	0	0	2	0	2	8	0	4	6	0

Forecast horizon	Canada			Norway			Mexico			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.995	1.034	1.006	0.999	1.029	1.299	1.000	1.025	1.179	1.000	1.011	0.965
	(0.89)	(0.15)	(0.92)	(0.98)	(0.18)	(0)	(0.98)	(0.29)	(0.02)	(0.99)	(0.51)	(0.56)
2	1.006	1.058	0.851	1.001	1.042	1.225	1.006	1.067	1.097	1.006	1.015	0.853
	(0.88)	(0.01)	(0.02)	(0.98)	(0.04)	(0.01)	(0.78)	(0.06)	(0.2)	(0.71)	(0.47)	(0.01)
3	1.007	1.045	0.883	1.000	1.044	1.319	1.011	1.074	1.063	0.987	0.994	0.821
	(0.82)	(0.02)	(0.1)	(0.99)	(0.1)	(0)	(0.61)	(0.1)	(0.41)	(0.48)	(0.83)	(0)
6	1.029	1.081	0.905	1.013	1.081	1.340	1.023	1.086	1.087	0.945	0.974	0.768
	(0.24)	(0)	(0.21)	(0.69)	(0.06)	(0)	(0.52)	(0.2)	(0.32)	(0.14)	(0.54)	(0)
9	1.026	1.090	0.789	1.003	1.085	1.371	1.031	1.128	1.116	0.915	0.954	0.763
	(0.5)	(0.02)	(0.01)	(0.95)	(0.2)	(0)	(0.61)	(0.19)	(0.17)	(0.23)	(0.54)	(0)
12	1.038	1.095	1.072	0.999	1.084	1.398	1.033	1.136	1.087	0.962	0.973	0.768
	(0.4)	(0.03)	(0.41)	(0.99)	(0.3)	(0)	(0.65)	(0.22)	(0.3)	(0.58)	(0.77)	(0)
18	1.141	1.235	1.016	1.051	1.208	1.373	1.107	1.156	1.094	1.113	1.140	0.787
	(0.01)	(0)	(0.86)	(0.5)	(0.09)	(0)	(0.2)	(0.24)	(0.27)	(0.31)	(0.28)	(0)
24	1.215	1.350	1.118	1.094	1.276	1.195	1.139	1.187	1.031	1.174	1.255	1.031
	(0)	(0)	(0.2)	(0.37)	(0.06)	(0.03)	(0.2)	(0.18)	(0.68)	(0.24)	(0.14)	(0.71)
No. of Stat. < 1	1	0	4	2	0	0	1	0	0	5	4	7
(of which sig.)	0	0	3	0	0	0	0	0	0	0	0	6
No. of Stat. >= 1	7	8	4	6	8	8	7	8	8	3	4	1
(of which sig.)	2	7	0	0	5	8	0	2	1	0	0	0

Table - 6. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Full Sample, with *Ex post* Oil Prices)

Forecast horizon	US			Japan			Germany			UK		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.990	0.947	1.146	0.952	0.941	1.072	1.041	1.089	1.198	1.086	1.004	1.054
	(0.79)	(0.14)	(0)	(0.35)	(0.17)	(0.2)	(0.07)	(0)	(0.02)	(0.4)	(0.92)	(0.16)
2	0.994	0.973	1.064	0.953	0.979	0.972	1.072	1.102	1.160	1.064	1.022	1.027
	(0.91)	(0.49)	(0.32)	(0.51)	(0.72)	(0.69)	(0.11)	(0.02)	(0.08)	(0.41)	(0.53)	(0.53)
3	0.981	0.980	1.059	0.945	0.986	0.912	1.095	1.094	1.144	1.050	1.023	0.973
	(0.78)	(0.75)	(0.44)	(0.47)	(0.82)	(0.23)	(0.11)	(0.1)	(0.09)	(0.39)	(0.44)	(0.59)
6	0.996	1.060	1.225	0.930	0.999	0.874	1.086	1.090	0.985	1.021	1.015	0.851
	(0.97)	(0.47)	(0.03)	(0.36)	(0.99)	(0.11)	(0.11)	(0.27)	(0.86)	(0.48)	(0.67)	(0.01)
9	1.047	1.104	1.282	0.924	0.978	0.925	1.067	1.063	1.000	1.006	1.011	0.940
	(0.67)	(0.33)	(0.01)	(0.34)	(0.82)	(0.33)	(0.24)	(0.56)	(1)	(0.84)	(0.8)	(0.38)
12	1.102	1.168	1.299	0.893	0.959	0.834	1.065	1.064	1.030	0.995	1.003	0.932
	(0.34)	(0.14)	(0.01)	(0.21)	(0.69)	(0.02)	(0.28)	(0.57)	(0.72)	(0.9)	(0.94)	(0.35)
18	1.213	1.286	1.273	0.819	0.886	0.763	1.088	1.045	0.993	0.972	1.002	0.794
	(0.06)	(0.03)	(0.02)	(0.05)	(0.26)	(0)	(0.27)	(0.73)	(0.93)	(0.56)	(0.98)	(0)
24	1.272	1.416	1.381	0.764	0.800	0.776	1.145	1.034	1.110	0.979	1.076	0.872
	(0.03)	(0)	(0)	(0.01)	(0.03)	(0)	(0.18)	(0.82)	(0.2)	(0.72)	(0.36)	(0.09)
No. of Stat. < 1	4	3	0	8	8	7	0	0	2	3	0	6
(of which sig.)	0	0	0	2	1	3	0	0	0	0	0	3
No. of Stat. >= 1	4	5	8	0	0	1	8	8	6	5	8	2
(of which sig.)	2	2	6	0	0	0	1	3	3	0	0	0

Forecast horizon	Canada			Norway			Mexico			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.016	0.983	1.019	0.930	0.946	1.051	1.029	1.146	1.299	1.117	1.102	1.122
	(0.36)	(0.43)	(0.63)	(0.2)	(0.37)	(0.37)	(0.55)	(0.02)	(0)	(0.27)	(0.19)	(0.06)
2	1.021	1.011	1.007	0.924	0.954	0.975	1.058	1.147	1.105	1.091	1.092	1.079
	(0.14)	(0.59)	(0.88)	(0.29)	(0.53)	(0.66)	(0.37)	(0.06)	(0.16)	(0.24)	(0.1)	(0.26)
3	1.013	1.005	0.965	0.905	0.930	0.939	1.109	1.165	1.109	1.097	1.128	1.036
	(0.42)	(0.75)	(0.47)	(0.27)	(0.42)	(0.3)	(0.21)	(0.07)	(0.18)	(0.19)	(0.06)	(0.62)
6	1.028	1.015	1.064	0.899	0.912	0.821	1.124	1.149	1.113	1.089	1.186	0.949
	(0.16)	(0.44)	(0.35)	(0.32)	(0.42)	(0)	(0.19)	(0.13)	(0.2)	(0.09)	(0.04)	(0.5)
9	1.045	1.068	1.094	0.906	0.890	0.800	1.133	1.122	1.125	1.088	1.187	0.963
	(0.07)	(0.02)	(0.29)	(0.32)	(0.34)	(0)	(0.08)	(0.22)	(0.15)	(0.07)	(0.1)	(0.66)
12	1.056	1.082	1.264	0.902	0.835	0.749	1.109	1.104	1.015	1.160	1.290	0.947
	(0.06)	(0.03)	(0)	(0.22)	(0.13)	(0)	(0.11)	(0.33)	(0.86)	(0.02)	(0.04)	(0.51)
18	1.084	1.094	1.240	0.880	0.800	0.773	1.090	1.076	0.993	1.290	1.448	0.976
	(0.04)	(0.05)	(0.01)	(0.13)	(0.07)	(0)	(0.23)	(0.49)	(0.93)	(0)	(0.02)	(0.78)
24	1.110	1.113	1.385	0.845	0.758	0.750	1.086	1.064	0.931	1.515	1.659	1.436
	(0.03)	(0.04)	(0)	(0.06)	(0.04)	(0)	(0.34)	(0.61)	(0.32)	(0)	(0.01)	(0)
No. of Stat. < 1	0	1	1	8	8	7	0	0	2	0	0	4
(of which sig.)	0	0	0	1	2	5	0	0	0	0	0	0
No. of Stat. >= 1	8	7	7	0	0	1	8	8	6	8	8	4
(of which sig.)	4	4	3	0	0	0	1	3	1	5	7	2

Table - 7. Forecast Evaluation of the Nonlinear Oil-Price Exchange Rate Model

(VAR in Levels, Full Sample, with *Ex post* Oil Prices)

Forecast horizon	US			Japan			Germany			UK		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.043	0.997	1.109	1.052	1.029	1.032	1.105	1.231	1.269	1.113	1.051	1.054
	(0.35)	(0.95)	(0.03)	(0.51)	(0.6)	(0.62)	(0)	(0)	(0)	(0.22)	(0.34)	(0.31)
2	1.071	1.016	0.987	1.087	1.078	0.934	1.198	1.297	1.218	1.157	1.116	1.000
	(0.27)	(0.73)	(0.82)	(0.37)	(0.18)	(0.35)	(0.01)	(0)	(0.02)	(0.17)	(0.12)	(1)
3	1.065	1.015	0.935	1.090	1.117	0.931	1.303	1.392	1.269	1.199	1.142	0.967
	(0.31)	(0.77)	(0.28)	(0.29)	(0.11)	(0.38)	(0.01)	(0)	(0)	(0.19)	(0.09)	(0.59)
6	1.029	1.073	0.986	1.062	1.131	0.912	7.600	2.559	1.239	1.162	1.157	0.907
	(0.71)	(0.31)	(0.85)	(0.36)	(0.1)	(0.31)	(0.28)	(0.12)	(0.02)	(0.18)	(0.11)	(0.17)
9	1.149	1.139	1.022	1.096	1.147	0.951	2.9E+26	2.6E+25	1.180	1.015	1.070	0.979
	(0.11)	(0.13)	(0.78)	(0.26)	(0.14)	(0.55)	(0.3)	(0.3)	(0.05)	(0.86)	(0.52)	(0.78)
12	1.266	1.206	1.053	1.091	1.159	0.868	5.2E+14	4.7E+13	1.198	0.943	1.008	0.851
	(0.08)	(0.07)	(0.53)	(0.37)	(0.17)	(0.09)	(0.31)	(0.31)	(0.04)	(0.65)	(0.95)	(0.03)
18	1.419	1.337	1.041	1.8E+03	1.4E+02	0.827	6.2E+19	5.4E+18	1.078	0.888	1.016	0.789
	(0.08)	(0.04)	(0.65)	(0.3)	(0.3)	(0.02)	(0.31)	(0.31)	(0.33)	(0.49)	(0.92)	(0)
24	1.383	1.398	1.207	0.884	0.905	0.776	9.0E+22	7.6E+21	1.380	0.903	1.062	0.822
	(0.06)	(0.03)	(0.03)	(0.28)	(0.41)	(0)	(0.3)	(0.3)	(0)	(0.54)	(0.7)	(0.01)
No. of Stat. < 1	0	1	3	1	1	7	0	0	0	3	0	6
(of which sig.)	0	0	0	0	0	3	0	0	0	0	0	3
No. of Stat. >= 1	8	7	5	7	7	1	8	8	8	5	8	2
(of which sig.)	3	3	2	0	1	0	3	3	7	0	1	0

Forecast horizon	Canada			Norway			Mexico			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.174	1.069	1.033	1.089	1.084	1.122	1.245	1.338	1.260	1.219	1.189	1.146
	(0.21)	(0.36)	(0.45)	(0.27)	(0.17)	(0.05)	(0.04)	(0)	(0)	(0.27)	(0.02)	(0.03)
2	1.130	1.090	1.044	1.167	1.150	1.033	1.278	1.379	1.234	1.431	1.236	1.020
	(0.07)	(0.04)	(0.42)	(0.12)	(0.05)	(0.61)	(0.04)	(0)	(0.01)	(0.27)	(0.03)	(0.77)
3	1.138	1.087	1.046	1.593	1.218	1.013	1.326	1.413	1.205	1.839	1.321	0.934
	(0.08)	(0.01)	(0.46)	(0.21)	(0.08)	(0.84)	(0.02)	(0)	(0.02)	(0.28)	(0.05)	(0.33)
6	1.168	1.155	1.231	1.2E+09	1.2E+08	0.979	1.333	1.335	1.053	2.0E+06	2.2E+05	0.843
	(0.04)	(0)	(0.01)	(0.3)	(0.3)	(0.78)	(0.04)	(0)	(0.51)	(0.3)	(0.3)	(0.03)
9	1.166	1.223	1.303	1.165	1.198	1.036	1.310	1.298	1.075	1.250	1.304	0.861
	(0.01)	(0)	(0.01)	(0.08)	(0.1)	(0.65)	(0.02)	(0.01)	(0.35)	(0.1)	(0.08)	(0.06)
12	1.179	1.270	1.836	1.163	1.175	1.015	1.246	1.237	1.054	1.365	1.395	0.840
	(0)	(0)	(0)	(0.09)	(0.19)	(0.86)	(0.01)	(0.01)	(0.51)	(0.1)	(0.08)	(0.03)
18	1.181	1.273	1.817	1.184	1.186	1.044	1.251	1.139	1.014	1.605	1.635	0.953
	(0)	(0)	(0)	(0.11)	(0.21)	(0.59)	(0.09)	(0.26)	(0.84)	(0.06)	(0.03)	(0.56)
24	1.226	1.291	2.254	1.195	1.184	1.080	1.308	1.151	1.055	1.692	1.681	1.107
	(0)	(0)	(0)	(0.1)	(0.19)	(0.35)	(0.11)	(0.27)	(0.48)	(0.06)	(0.03)	(0.2)
No. of Stat. < 1	0	0	0	0	0	1	0	0	0	0	0	5
(of which sig.)	0	0	0	0	0	0	0	0	0	0	0	3
No. of Stat. >= 1	8	8	8	8	8	7	8	8	8	8	8	3
(of which sig.)	7	7	5	3	3	1	7	6	3	4	7	1

Table - 8. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Full Sample, Balassa-Samuelson Effect in Restricted Form)

Forecast horizon	US			Japan			Germany			UK		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.988	0.949	1.116	0.980	0.968	1.052	1.034	1.096	1.189	1.000	1.005	1.319
	(0.75)	(0.13)	(0.01)	(0.65)	(0.41)	(0.28)	(0.01)	(0)	(0.03)	(0.97)	(0.35)	(0)
2	1.006	0.994	1.087	0.998	1.013	1.014	1.053	1.090	1.133	1.001	1.006	1.213
	(0.88)	(0.85)	(0.19)	(0.97)	(0.74)	(0.83)	(0.08)	(0.02)	(0.13)	(0.92)	(0.45)	(0.02)
3	1.008	1.019	1.059	1.007	1.034	0.993	1.061	1.089	1.110	1.000	1.005	1.208
	(0.84)	(0.63)	(0.43)	(0.83)	(0.4)	(0.92)	(0.11)	(0.09)	(0.19)	(0.98)	(0.59)	(0.03)
6	1.063	1.100	1.236	1.033	1.081	1.000	1.030	1.047	0.877	1.000	1.004	1.096
	(0.28)	(0.11)	(0.03)	(0.46)	(0.18)	(1)	(0.56)	(0.57)	(0.1)	(0.99)	(0.74)	(0.3)
9	1.091	1.129	1.259	1.015	1.063	1.015	1.028	1.046	0.929	1.004	1.010	1.111
	(0.22)	(0.13)	(0.01)	(0.8)	(0.47)	(0.86)	(0.67)	(0.68)	(0.34)	(0.77)	(0.51)	(0.23)
12	1.118	1.177	1.264	0.982	1.038	0.916	1.024	1.048	0.921	1.003	1.002	1.062
	(0.17)	(0.06)	(0.02)	(0.81)	(0.7)	(0.29)	(0.76)	(0.68)	(0.3)	(0.84)	(0.9)	(0.48)
18	1.165	1.214	1.212	0.915	0.968	0.843	1.035	1.025	0.890	0.993	0.986	0.948
	(0.12)	(0.07)	(0.06)	(0.31)	(0.76)	(0.03)	(0.71)	(0.86)	(0.12)	(0.73)	(0.55)	(0.53)
24	1.179	1.236	1.301	0.887	0.912	0.892	1.077	1.006	0.993	0.998	0.999	1.049
	(0.16)	(0.09)	(0.01)	(0.21)	(0.41)	(0.15)	(0.51)	(0.97)	(0.93)	(0.93)	(0.96)	(0.6)
No. of Stat. < 1	1	2	0	5	3	4	0	0	5	3	2	1
(of which sig.)	0	0	0	0	0	1	0	0	1	0	0	0
No. of Stat. >= 1	7	6	8	3	5	4	8	8	3	5	6	7
(of which sig.)	0	3	6	0	0	0	2	3	1	0	0	3

Forecast horizon	Canada			Norway			Mexico			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.014	0.981	1.033	0.958	0.975	1.115	0.968	1.054	1.146	1.122	1.075	1.146
	(0.48)	(0.43)	(0.46)	(0.31)	(0.64)	(0.05)	(0.64)	(0.39)	(0.02)	(0.27)	(0.36)	(0.01)
2	1.017	1.007	1.007	0.982	1.004	1.000	1.008	1.132	1.206	1.042	1.044	1.041
	(0.35)	(0.78)	(0.89)	(0.66)	(0.95)	(1)	(0.91)	(0.09)	(0.01)	(0.21)	(0.2)	(0.52)
3	1.013	1.003	1.071	0.995	0.998	0.981	1.033	1.127	1.186	1.065	1.060	0.979
	(0.54)	(0.9)	(0.25)	(0.89)	(0.97)	(0.75)	(0.6)	(0.1)	(0.03)	(0.17)	(0.13)	(0.75)
6	1.029	1.027	1.127	1.026	1.002	0.926	1.099	1.163	1.211	1.034	1.050	0.921
	(0.28)	(0.32)	(0.11)	(0.5)	(0.98)	(0.28)	(0.08)	(0.09)	(0.05)	(0.16)	(0.26)	(0.31)
9	1.045	1.061	1.045	1.031	0.991	0.941	1.127	1.188	1.125	1.029	1.056	0.942
	(0.16)	(0.05)	(0.62)	(0.56)	(0.93)	(0.39)	(0.06)	(0.09)	(0.18)	(0.33)	(0.38)	(0.47)
12	1.055	1.066	1.165	1.009	0.967	0.932	1.106	1.162	1.070	1.041	1.088	0.920
	(0.12)	(0.12)	(0.08)	(0.87)	(0.74)	(0.35)	(0.15)	(0.16)	(0.45)	(0.29)	(0.24)	(0.33)
18	1.074	1.082	1.217	0.974	0.936	0.966	1.060	1.082	1.176	1.066	1.156	0.976
	(0.06)	(0.11)	(0.03)	(0.7)	(0.56)	(0.65)	(0.48)	(0.52)	(0.07)	(0.13)	(0.08)	(0.79)
24	1.089	1.102	1.371	0.955	0.901	0.971	1.025	1.027	1.055	1.097	1.192	1.337
	(0.07)	(0.09)	(0)	(0.56)	(0.41)	(0.71)	(0.79)	(0.84)	(0.53)	(0.17)	(0.09)	(0)
No. of Stat. < 1	0	1	0	5	6	6	1	0	0	0	0	5
(of which sig.)	0	0	0	0	0	0	0	0	0	0	0	0
No. of Stat. >= 1	8	7	8	3	2	2	7	8	8	8	8	3
(of which sig.)	2	2	3	0	0	1	2	4	5	0	2	2

Table - 9. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Full Sample, Balassa-Samuelson Effect in Unrestricted Form)

Forecast horizon	US			Japan			Germany			UK		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.002	0.954	1.153	0.979	0.972	1.065	1.032	1.063	1.180	1.085	1.018	1.113
	(0.96)	(0.17)	(0)	(0.66)	(0.5)	(0.17)	(0.07)	(0.03)	(0.04)	(0.34)	(0.61)	(0.02)
2	1.025	1.002	1.071	1.008	1.027	1.060	1.049	1.098	1.115	1.068	1.040	1.027
	(0.45)	(0.94)	(0.27)	(0.83)	(0.46)	(0.38)	(0.08)	(0.02)	(0.19)	(0.27)	(0.15)	(0.63)
3	1.035	1.035	1.092	1.020	1.045	1.000	1.060	1.113	1.258	1.057	1.067	1.051
	(0.35)	(0.38)	(0.24)	(0.61)	(0.26)	(1)	(0.09)	(0.06)	(0.01)	(0.21)	(0.02)	(0.46)
6	1.084	1.129	1.225	1.042	1.084	1.016	1.082	1.099	0.993	1.054	1.080	1.015
	(0.15)	(0.04)	(0.03)	(0.42)	(0.17)	(0.85)	(0.14)	(0.31)	(0.93)	(0.07)	(0.03)	(0.85)
9	1.116	1.174	1.259	1.015	1.054	1.015	1.096	1.082	0.986	1.046	1.074	1.007
	(0.13)	(0.05)	(0.01)	(0.82)	(0.54)	(0.86)	(0.18)	(0.52)	(0.87)	(0.25)	(0.14)	(0.93)
12	1.147	1.229	1.252	0.972	1.022	0.923	1.096	1.092	0.986	1.042	1.048	0.979
	(0.09)	(0.02)	(0.02)	(0.71)	(0.83)	(0.33)	(0.26)	(0.5)	(0.87)	(0.39)	(0.38)	(0.79)
18	1.205	1.276	1.273	0.895	0.940	0.838	1.133	1.116	0.972	1.019	1.040	0.920
	(0.05)	(0.02)	(0.02)	(0.24)	(0.59)	(0.03)	(0.25)	(0.46)	(0.72)	(0.75)	(0.64)	(0.32)
24	1.222	1.303	1.327	0.870	0.892	0.880	1.217	1.136	1.068	1.014	1.070	1.000
	(0.08)	(0.03)	(0)	(0.17)	(0.34)	(0.11)	(0.14)	(0.44)	(0.42)	(0.84)	(0.5)	(1)
No. of Stat. < 1	0	1	0	4	3	3	0	0	4	0	0	2
(of which sig.)	0	0	0	0	0	1	0	0	0	0	0	0
No. of Stat. >= 1	8	7	8	4	5	5	8	8	4	8	8	6
(of which sig.)	3	5	6	0	0	0	3	3	2	1	2	1

Forecast horizon	Canada			Norway			Mexico			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.011	0.986	1.013	0.984	1.041	1.162	1.033	1.156	1.260	1.001	1.012	1.146
	(0.54)	(0.55)	(0.76)	(0.69)	(0.49)	(0.01)	(0.42)	(0.01)	(0)	(0.72)	(0.65)	(0.05)
2	1.020	1.009	1.021	1.013	1.059	1.053	1.099	1.245	1.339	0.995	1.013	1.071
	(0.32)	(0.73)	(0.67)	(0.75)	(0.41)	(0.43)	(0.05)	(0)	(0)	(0.75)	(0.7)	(0.35)
3	1.020	1.012	1.124	1.033	1.065	1.070	1.162	1.307	1.214	0.992	0.999	0.953
	(0.39)	(0.64)	(0.07)	(0.37)	(0.37)	(0.34)	(0.01)	(0)	(0.01)	(0.73)	(0.97)	(0.51)
6	1.043	1.042	1.177	1.098	1.090	1.053	1.291	1.404	1.113	1.012	1.039	0.884
	(0.22)	(0.2)	(0.05)	(0.05)	(0.34)	(0.51)	(0.01)	(0)	(0.21)	(0.65)	(0.4)	(0.13)
9	1.061	1.078	1.094	1.121	1.092	1.091	1.321	1.377	1.059	1.025	1.070	0.928
	(0.14)	(0.04)	(0.35)	(0.06)	(0.4)	(0.27)	(0.01)	(0.02)	(0.46)	(0.47)	(0.23)	(0.36)
12	1.072	1.084	1.288	1.094	1.062	1.054	1.267	1.313	1.070	1.016	1.060	0.869
	(0.09)	(0.08)	(0)	(0.16)	(0.58)	(0.53)	(0.02)	(0.05)	(0.41)	(0.73)	(0.34)	(0.08)
18	1.094	1.113	1.344	1.036	1.014	1.007	1.173	1.228	1.207	1.062	1.130	0.968
	(0.02)	(0.03)	(0)	(0.64)	(0.91)	(0.93)	(0.07)	(0.09)	(0.03)	(0.27)	(0.09)	(0.72)
24	1.118	1.137	1.462	1.006	0.965	0.993	1.100	1.171	1.164	1.221	1.240	1.337
	(0.01)	(0.02)	(0)	(0.94)	(0.78)	(0.93)	(0.31)	(0.19)	(0.09)	(0.09)	(0.03)	(0)
No. of Stat. < 1	0	1	0	1	1	1	0	0	0	2	1	5
(of which sig.)	0	0	0	0	0	0	0	0	0	0	0	1
No. of Stat. >= 1	8	7	8	7	7	7	8	8	8	6	7	3
(of which sig.)	3	4	5	2	0	1	6	7	5	1	2	2

Table - 10. Forecast Evaluation of the Terms-of-trade Exchange Rate Model

(VAR in Levels, Full Sample)

Forecast horizon	US			Japan			Germany		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.987	0.955	1.102	0.967	0.935	1.032	1.028	1.062	1.180
	(0.67)	(0.12)	(0.01)	(0.43)	(0.11)	(0.4)	(0.14)	(0.04)	(0.04)
2	1.000	0.988	1.056	0.979	0.967	0.979	1.031	1.066	1.058
	(1)	(0.68)	(0.29)	(0.58)	(0.49)	(0.68)	(0.19)	(0.1)	(0.48)
3	0.994	1.002	1.029	0.984	0.974	0.925	1.030	1.050	1.102
	(0.87)	(0.95)	(0.63)	(0.64)	(0.55)	(0.17)	(0.35)	(0.37)	(0.22)
6	0.983	1.000	1.015	1.002	1.031	0.969	1.032	1.034	0.877
	(0.76)	(1)	(0.84)	(0.96)	(0.59)	(0.68)	(0.55)	(0.7)	(0.1)
9	0.981	1.000	1.052	0.982	1.018	1.038	1.036	1.043	0.900
	(0.78)	(0.99)	(0.5)	(0.75)	(0.81)	(0.63)	(0.62)	(0.71)	(0.16)
12	0.985	0.988	1.007	0.962	0.996	0.985	1.028	1.040	0.897
	(0.83)	(0.88)	(0.93)	(0.59)	(0.96)	(0.85)	(0.74)	(0.75)	(0.16)
18	1.002	1.011	0.969	0.934	0.941	0.942	1.045	1.027	0.857
	(0.98)	(0.9)	(0.71)	(0.43)	(0.54)	(0.45)	(0.67)	(0.86)	(0.03)
24	1.017	1.067	1.098	0.912	0.906	0.943	1.093	1.022	1.000
	(0.86)	(0.48)	(0.29)	(0.33)	(0.36)	(0.45)	(0.47)	(0.89)	(1)
No. of Stat. < 1	5	4	1	7	6	6	0	0	4
(of which sig.)	0	0	0	0	0	0	0	0	2
No. of Stat. >= 1	3	4	7	1	2	2	8	8	4
(of which sig.)	0	0	1	0	0	0	0	2	1

Forecast horizon	UK			Canada			Turkey		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.081	1.001	1.000	1.016	0.983	1.019	1.115	1.074	1.162
	(0.39)	(0.97)	(1)	(0.4)	(0.41)	(0.6)	(0.29)	(0.37)	(0.01)
2	1.058	1.001	0.975	1.029	1.025	1.014	1.040	1.020	1.102
	(0.41)	(0.97)	(0.56)	(0.11)	(0.24)	(0.79)	(0.3)	(0.45)	(0.17)
3	1.043	1.023	0.967	1.031	1.028	0.971	1.052	1.044	1.052
	(0.41)	(0.42)	(0.56)	(0.17)	(0.21)	(0.64)	(0.28)	(0.22)	(0.5)
6	1.018	1.028	0.938	1.058	1.058	1.090	1.021	1.006	0.890
	(0.52)	(0.43)	(0.36)	(0.13)	(0.08)	(0.3)	(0.46)	(0.89)	(0.14)
9	1.005	1.033	1.053	1.091	1.119	1.018	0.995	1.011	0.843
	(0.86)	(0.44)	(0.51)	(0.09)	(0.01)	(0.86)	(0.86)	(0.86)	(0.03)
12	0.994	1.010	0.958	1.115	1.141	1.107	1.007	1.058	0.906
	(0.87)	(0.84)	(0.58)	(0.1)	(0.03)	(0.24)	(0.84)	(0.45)	(0.25)
18	0.965	1.004	0.858	1.141	1.159	1.057	1.015	1.113	0.830
	(0.47)	(0.95)	(0.05)	(0.09)	(0.05)	(0.52)	(0.8)	(0.27)	(0.03)
24	0.948	1.034	0.915	1.145	1.179	1.198	1.071	1.122	1.080
	(0.38)	(0.68)	(0.29)	(0.08)	(0.06)	(0.04)	(0.35)	(0.34)	(0.36)
No. of Stat. < 1	3	0	6	0	1	1	1	0	4
(of which sig.)	0	0	1	0	0	0	0	0	2
No. of Stat. >= 1	5	8	2	8	7	7	7	8	4
(of which sig.)	0	0	0	4	5	1	0	0	1

Table - 11.1. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Subsample, US)

Forecast horizon	1988M1-92M12			1993M1-97M12			1998M1-2002M12			2003M1-09M7		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.932	0.882	1.075	0.892	0.854	1.024	1.000	0.957	1.171	1.049	1.031	0.939
	(0.17)	(0.04)	(0.37)	(0.03)	(0.01)	(0.77)	(1)	(0.56)	(0.13)	(0.4)	(0.54)	(0.47)
2	1.009	0.942	0.973	0.970	0.933	0.921	1.131	1.101	1.276	1.056	1.050	0.875
	(0.85)	(0.38)	(0.82)	(0.33)	(0.22)	(0.44)	(0.08)	(0.19)	(0.09)	(0.34)	(0.35)	(0.16)
3	1.066	1.015	0.970	1.030	1.001	0.838	1.247	1.256	1.462	1.048	0.978	0.768
	(0.3)	(0.84)	(0.84)	(0.48)	(0.98)	(0.18)	(0.01)	(0.01)	(0.02)	(0.41)	(0.7)	(0.01)
6	1.243	1.196	0.800	1.113	1.108	1.522	1.447	1.422	2.053	1.159	1.030	0.882
	(0.12)	(0.23)	(0.31)	(0.1)	(0.12)	(0.02)	(0)	(0.01)	(0)	(0.24)	(0.72)	(0.24)
9	1.366	1.313	1.154	1.150	1.125	1.444	1.465	1.526	1.842	1.254	1.073	1.021
	(0.18)	(0.24)	(0.5)	(0.19)	(0.28)	(0.02)	(0)	(0)	(0.01)	(0.27)	(0.59)	(0.85)
12	1.466	1.392	1.227	1.085	1.168	1.379	1.498	1.555	1.947	1.286	1.091	1.022
	(0.1)	(0.13)	(0.39)	(0.47)	(0.32)	(0.05)	(0)	(0)	(0.01)	(0.3)	(0.61)	(0.85)
18	1.604	1.680	1.238	1.132	1.162	1.500	1.447	1.619	2.235	1.052	0.940	0.980
	(0.06)	(0.02)	(0.38)	(0.36)	(0.41)	(0.02)	(0)	(0)	(0)	(0.76)	(0.68)	(0.85)
24	1.814	2.068	1.391	1.200	1.224	1.481	1.429	1.637	4.091	0.865	0.827	0.981
	(0)	(0)	(0.03)	(0.15)	(0.34)	(0.02)	(0)	(0)	(0)	(0.07)	(0.09)	(0.84)
No. of Stat. < 1	1	2	3	2	2	2	1	1	0	1	3	6
(of which sig.)	0	1	0	1	1	0	0	0	0	1	1	1
No. of Stat. >= 1	7	6	5	6	6	6	7	7	8	7	5	2
(of which sig.)	3	2	1	1	0	5	7	6	7	0	0	0

Table - 11.2. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Subsample, Japan)

Forecast horizon	1988M1-92M12			1993M1-97M12			1998M1-2002M12			2003M1-09M7		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.979	0.962	0.975	0.880	0.866	1.050	0.913	0.946	1.139	1.104	1.050	0.932
	(0.64)	(0.59)	(0.77)	(0.1)	(0.1)	(0.62)	(0.11)	(0.42)	(0.13)	(0.07)	(0.38)	(0.49)
2	1.030	0.989	0.868	0.938	0.959	1.265	0.924	0.946	0.946	1.165	1.168	1.000
	(0.64)	(0.89)	(0.2)	(0.34)	(0.59)	(0.08)	(0.22)	(0.29)	(0.6)	(0.02)	(0)	(1)
3	0.978	0.967	0.947	0.969	0.998	1.393	0.950	0.950	0.838	1.207	1.226	0.818
	(0.71)	(0.66)	(0.66)	(0.6)	(0.97)	(0.05)	(0.43)	(0.42)	(0.2)	(0.06)	(0.01)	(0.15)
6	0.872	0.864	0.947	1.012	1.088	1.231	1.011	1.032	0.886	1.371	1.381	0.791
	(0.05)	(0.04)	(0.69)	(0.85)	(0.3)	(0.31)	(0.89)	(0.76)	(0.47)	(0.13)	(0.03)	(0.13)
9	0.826	0.785	0.775	0.996	1.083	1.152	0.989	1.006	1.026	1.585	1.545	0.848
	(0.05)	(0.05)	(0.06)	(0.96)	(0.48)	(0.4)	(0.93)	(0.97)	(0.87)	(0.1)	(0.04)	(0.24)
12	0.752	0.732	0.878	0.980	1.114	1.156	0.931	0.899	1.000	1.748	1.655	0.864
	(0.04)	(0.02)	(0.23)	(0.84)	(0.41)	(0.4)	(0.62)	(0.53)	(1)	(0.09)	(0.03)	(0.31)
18	0.650	0.580	0.861	0.921	0.970	1.026	0.867	0.892	0.930	1.764	1.778	0.953
	(0.01)	(0)	(0.2)	(0.41)	(0.79)	(0.85)	(0.31)	(0.51)	(0.58)	(0.04)	(0)	(0.73)
24	0.544	0.517	0.829	0.856	0.848	0.959	0.840	0.819	1.024	1.793	1.919	1.500
	(0.02)	(0)	(0.08)	(0.07)	(0.05)	(0.69)	(0.15)	(0.19)	(0.85)	(0.01)	(0)	(0.01)
No. of Stat. < 1	7	8	8	7	5	1	7	6	4	0	0	6
(of which sig.)	5	5	2	2	2	0	0	0	0	0	0	0
No. of Stat. >= 1	1	0	0	1	3	7	1	2	4	8	8	2
(of which sig.)	0	0	0	0	0	2	0	0	0	7	7	1

Table - 11.3. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Subsample, Germany)

Forecast horizon	1988M1-92M12			1993M1-97M12			1998M1-2002M12			2003M1-09M7		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.023	1.122	1.321	1.027	1.029	1.393	1.069	1.058	1.179	1.021	1.035	1.079
	(0.3)	(0.09)	(0.12)	(0.36)	(0.42)	(0.06)	(0.02)	(0.01)	(0.37)	(0)	(0.02)	(0.64)
2	1.047	1.194	1.138	1.023	0.996	1.194	1.115	1.127	1.308	1.044	1.063	1.281
	(0.3)	(0.05)	(0.5)	(0.59)	(0.94)	(0.31)	(0.01)	(0.01)	(0.16)	(0)	(0)	(0.16)
3	1.055	1.186	1.464	1.019	0.991	1.121	1.167	1.156	1.850	1.077	1.102	1.345
	(0.33)	(0.09)	(0.02)	(0.73)	(0.88)	(0.5)	(0.01)	(0.03)	(0)	(0.01)	(0.01)	(0.11)
6	1.041	1.120	1.167	1.046	1.032	1.091	1.260	1.234	1.391	1.114	1.146	1.556
	(0.47)	(0.29)	(0.4)	(0.61)	(0.74)	(0.62)	(0.03)	(0.03)	(0.12)	(0.01)	(0.01)	(0.02)
9	1.035	1.080	1.067	0.992	0.965	1.108	1.327	1.291	1.545	1.116	1.187	1.846
	(0.49)	(0.51)	(0.74)	(0.93)	(0.71)	(0.48)	(0.03)	(0.05)	(0.03)	(0.13)	(0.05)	(0)
12	1.014	1.022	0.938	0.921	0.868	0.952	1.397	1.414	2.105	1.114	1.173	1.500
	(0.77)	(0.86)	(0.74)	(0.33)	(0.14)	(0.72)	(0.01)	(0)	(0)	(0.27)	(0.17)	(0.01)
18	1.030	0.965	1.000	0.796	0.731	0.800	1.431	1.545	2.800	1.043	1.063	1.167
	(0.56)	(0.72)	(1)	(0.02)	(0.01)	(0.04)	(0)	(0)	(0)	(0.8)	(0.7)	(0.24)
24	1.139	1.035	1.364	0.767	0.690	0.741	1.477	1.573	11.500	1.049	1.055	1.275
	(0.22)	(0.79)	(0.09)	(0)	(0)	(0)	(0)	(0)	(0)	(0.8)	(0.77)	(0.04)
No. of Stat. < 1	0	1	1	4	6	3	0	0	0	0	0	0
(of which sig.)	0	0	0	2	2	2	0	0	0	0	0	0
No. of Stat. >= 1	8	7	7	4	2	5	8	8	8	8	8	8
(of which sig.)	0	3	2	0	0	1	8	8	5	4	5	4

Table - 11.4. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Subsample, UK)

Forecast horizon	1988M1-92M12			1993M1-97M12			1998M1-2002M12			2003M1-09M7		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.917	0.949	1.114	1.002	0.974	1.051	1.034	1.032	1.313	1.111	1.059	0.895
	(0.38)	(0.38)	(0.04)	(0.98)	(0.79)	(0.62)	(0.5)	(0.57)	(0.02)	(0.28)	(0.29)	(0.52)
2	0.938	1.003	1.000	1.118	1.143	1.281	1.140	1.148	1.118	1.051	1.016	0.956
	(0.44)	(0.95)	(1)	(0.39)	(0.25)	(0.08)	(0.07)	(0.09)	(0.42)	(0.53)	(0.77)	(0.73)
3	0.942	0.985	1.138	1.196	1.247	1.161	1.198	1.199	1.029	1.008	0.992	0.956
	(0.41)	(0.85)	(0.21)	(0.25)	(0.13)	(0.36)	(0.07)	(0.18)	(0.85)	(0.91)	(0.87)	(0.72)
6	1.000	1.030	1.065	1.034	1.206	1.194	1.572	1.538	1.231	0.914	0.924	0.824
	(1)	(0.71)	(0.64)	(0.8)	(0.28)	(0.28)	(0.12)	(0.08)	(0.32)	(0.34)	(0.33)	(0.09)
9	1.017	1.057	1.097	0.965	1.119	1.258	1.744	1.763	1.385	0.889	0.922	0.843
	(0.78)	(0.58)	(0.54)	(0.79)	(0.56)	(0.1)	(0.1)	(0.04)	(0.1)	(0.3)	(0.43)	(0.15)
12	1.032	0.983	0.906	0.889	1.013	1.103	1.758	1.826	1.267	0.924	0.935	0.917
	(0.67)	(0.84)	(0.47)	(0.44)	(0.95)	(0.42)	(0.06)	(0.03)	(0.2)	(0.36)	(0.51)	(0.48)
18	1.029	0.969	1.111	0.793	0.898	0.975	1.379	1.721	1.194	0.961	0.987	0.940
	(0.83)	(0.83)	(0.44)	(0.22)	(0.6)	(0.84)	(0.06)	(0.03)	(0.32)	(0.46)	(0.88)	(0.62)
24	1.099	1.174	2.462	0.766	0.863	0.875	1.092	1.396	1.135	0.928	0.940	0.980
	(0.52)	(0.29)	(0)	(0.15)	(0.47)	(0.3)	(0.64)	(0.11)	(0.39)	(0.19)	(0.49)	(0.87)
No. of Stat. < 1	3	4	1	4	3	2	0	0	0	5	6	8
(of which sig.)	0	0	0	0	0	0	0	0	0	0	0	1
No. of Stat. >= 1	5	4	7	4	5	6	8	8	8	3	2	0
(of which sig.)	0	0	2	0	0	2	5	5	2	0	0	0

Table - 11.5. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Subsample, Canada)

Forecast horizon	1988M1-92M12			1993M1-97M12			1998M1-2002M12			2003M1-09M7		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.892	0.854	0.951	1.020	0.960	1.091	0.983	0.924	0.902	1.036	1.030	1.023
	(0.06)	(0.01)	(0.57)	(0.64)	(0.51)	(0.26)	(0.75)	(0.11)	(0.25)	(0.12)	(0.17)	(0.78)
2	0.900	0.881	0.976	1.051	1.037	1.036	0.998	0.977	0.946	1.041	1.048	1.176
	(0.11)	(0.08)	(0.8)	(0.17)	(0.51)	(0.74)	(0.95)	(0.64)	(0.57)	(0.03)	(0.01)	(0.11)
3	0.919	0.922	1.059	1.064	1.052	1.000	0.995	1.025	0.917	1.041	1.021	1.158
	(0.2)	(0.19)	(0.66)	(0.06)	(0.13)	(1)	(0.87)	(0.57)	(0.41)	(0.17)	(0.43)	(0.16)
6	0.959	0.931	0.944	1.216	1.140	1.138	1.016	0.981	1.031	1.076	1.050	1.294
	(0.52)	(0.33)	(0.66)	(0.03)	(0.04)	(0.38)	(0.55)	(0.68)	(0.82)	(0.12)	(0.08)	(0.07)
9	0.964	0.927	0.838	1.453	1.419	1.231	1.028	0.983	0.971	1.091	1.085	1.321
	(0.57)	(0.22)	(0.16)	(0.02)	(0.02)	(0.22)	(0.42)	(0.74)	(0.83)	(0.12)	(0.02)	(0.12)
12	0.938	0.925	1.030	1.501	1.428	1.393	1.033	1.003	1.172	1.083	1.076	1.323
	(0.22)	(0.18)	(0.82)	(0.03)	(0.04)	(0.03)	(0.4)	(0.95)	(0.3)	(0.1)	(0.1)	(0.1)
18	0.919	0.922	1.138	1.412	1.400	1.565	1.064	1.044	1.636	1.001	1.011	1.278
	(0.08)	(0.11)	(0.32)	(0.01)	(0.04)	(0.03)	(0.11)	(0.28)	(0.01)	(0.98)	(0.66)	(0.11)
24	0.911	0.896	1.269	1.321	1.355	1.773	1.100	1.102	2.929	1.036	1.055	2.038
	(0.03)	(0.04)	(0.05)	(0.02)	(0.03)	(0)	(0)	(0.02)	(0)	(0.24)	(0.07)	(0)
No. of Stat. < 1	8	8	4	0	1	0	3	4	4	0	0	0
(of which sig.)	3	3	0	0	0	0	0	0	0	0	0	0
No. of Stat. >= 1	0	0	4	8	7	8	5	4	4	8	8	8
(of which sig.)	0	0	1	6	5	3	1	1	2	2	5	3

Table - 11.6. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Subsample, Norway)

Forecast horizon	1988M1-92M12			1993M1-97M12			1998M1-2002M12			2003M1-09M7		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.933	1.041	1.162	0.956	0.982	1.061	1.021	1.029	1.355	0.867	0.904	0.898
	(0.67)	(0.73)	(0.16)	(0.41)	(0.79)	(0.62)	(0.74)	(0.61)	(0.02)	(0.27)	(0.29)	(0.13)
2	1.023	1.139	1.258	0.949	0.923	0.974	1.145	1.153	1.212	0.929	0.861	0.846
	(0.83)	(0.26)	(0.09)	(0.23)	(0.3)	(0.83)	(0.15)	(0.13)	(0.16)	(0.34)	(0.13)	(0.06)
3	1.045	1.212	1.357	0.915	0.885	1.025	1.226	1.172	1.273	1.000	0.945	0.760
	(0.69)	(0.14)	(0.03)	(0.16)	(0.12)	(0.84)	(0.11)	(0.2)	(0.08)	(0.99)	(0.25)	(0.01)
6	1.110	1.307	1.333	0.903	0.790	0.851	1.405	1.430	1.652	1.172	1.069	0.761
	(0.53)	(0.13)	(0.13)	(0.08)	(0.02)	(0.16)	(0.06)	(0.08)	(0.01)	(0.25)	(0.48)	(0.02)
9	1.255	1.270	1.682	0.850	0.726	0.731	1.490	1.511	1.857	1.275	1.123	0.813
	(0.27)	(0.3)	(0)	(0.03)	(0.01)	(0.01)	(0.06)	(0.06)	(0)	(0.25)	(0.49)	(0.05)
12	1.328	1.224	1.478	0.784	0.672	0.709	1.550	1.539	1.792	1.333	1.172	0.814
	(0.18)	(0.41)	(0.04)	(0)	(0)	(0)	(0.04)	(0.04)	(0)	(0.26)	(0.41)	(0.07)
18	1.448	1.279	1.417	0.753	0.650	0.696	1.511	1.507	3.214	1.184	1.101	0.915
	(0.1)	(0.32)	(0.07)	(0)	(0)	(0)	(0.04)	(0.04)	(0)	(0.33)	(0.64)	(0.37)
24	1.604	1.535	1.500	0.751	0.661	0.667	1.357	1.357	2.474	1.059	0.963	1.045
	(0.09)	(0.12)	(0.05)	(0)	(0)	(0)	(0.08)	(0.14)	(0)	(0.65)	(0.86)	(0.64)
No. of Stat. < 1	1	0	0	8	8	6	0	0	0	3	4	7
(of which sig.)	0	0	0	5	5	4	0	0	0	0	0	5
No. of Stat. >= 1	7	8	8	0	0	2	8	8	8	5	4	1
(of which sig.)	2	0	6	0	0	0	5	4	7	0	0	0

Table - 11.7. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Subsample, Mexico)

Forecast horizon	1988M1-92M12			1993M1-97M12			1998M1-2002M12			2003M1-09M7		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	1.048	1.191	1.654	1.005	1.045	1.167	1.045	1.059	1.400	0.897	0.961	0.957
	(0.76)	(0.16)	(0)	(0.94)	(0.55)	(0.17)	(0.44)	(0.27)	(0.02)	(0.41)	(0.69)	(0.57)
2	1.172	1.315	1.696	1.001	1.053	1.233	1.165	1.168	1.333	0.954	0.925	0.846
	(0.03)	(0.01)	(0.01)	(0.98)	(0.49)	(0.13)	(0.08)	(0.07)	(0.07)	(0.53)	(0.38)	(0.06)
3	1.156	1.337	2.111	0.939	0.963	1.242	1.262	1.223	1.448	1.044	1.010	0.792
	(0.02)	(0)	(0)	(0.18)	(0.65)	(0.13)	(0.05)	(0.1)	(0.02)	(0.29)	(0.89)	(0.02)
6	1.047	1.150	1.524	0.921	0.868	1.081	1.528	1.563	1.727	1.266	1.204	0.761
	(0.63)	(0.19)	(0.05)	(0.06)	(0.08)	(0.59)	(0.01)	(0.03)	(0.01)	(0.17)	(0.19)	(0.02)
9	1.113	1.182	1.762	0.913	0.852	0.884	1.688	1.733	2.294	1.365	1.225	0.780
	(0.42)	(0.22)	(0.01)	(0.04)	(0.04)	(0.39)	(0.01)	(0.01)	(0)	(0.2)	(0.35)	(0.02)
12	1.156	1.129	1.545	0.872	0.813	0.830	1.774	1.803	2.688	1.483	1.324	0.761
	(0.32)	(0.41)	(0.04)	(0.01)	(0.02)	(0.17)	(0)	(0)	(0)	(0.19)	(0.26)	(0.02)
18	1.226	1.109	2.000	0.844	0.802	0.867	1.700	1.762	3.462	1.431	1.311	0.896
	(0.25)	(0.6)	(0)	(0)	(0.01)	(0.32)	(0)	(0)	(0)	(0.14)	(0.31)	(0.28)
24	1.295	1.170	1.500	0.831	0.822	0.909	1.491	1.520	3.615	1.357	1.307	1.000
	(0.24)	(0.53)	(0.05)	(0)	(0.02)	(0.51)	(0.01)	(0.03)	(0)	(0.11)	(0.24)	(1)
No. of Stat. < 1	0	0	0	6	6	4	0	0	0	2	2	7
(of which sig.)	0	0	0	5	5	0	0	0	0	0	0	5
No. of Stat. >= 1	8	8	8	2	2	4	8	8	8	6	6	1
(of which sig.)	2	2	8	0	0	0	7	7	8	0	0	0

Table - 11.8. Forecast Evaluation of the Oil-Price Exchange Rate Model

(VAR in Levels, Subsample, Turkey)

Forecast horizon	1988M1-92M12			1993M1-97M12			1998M1-2002M12			2003M1-09M7		
	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR	MSE	MAE	DIR
1	0.972	0.941	1.257	0.982	1.017	1.464	0.968	0.878	1.000	1.108	1.181	0.900
	(0.62)	(0.38)	(0.06)	(0.71)	(0.81)	(0.01)	(0.63)	(0.08)	(1)	(0.22)	(0.11)	(0.42)
2	1.007	0.985	1.156	0.999	1.130	1.250	0.983	0.922	0.952	1.029	1.038	0.977
	(0.93)	(0.85)	(0.36)	(0.99)	(0.05)	(0.18)	(0.79)	(0.33)	(0.6)	(0.27)	(0.22)	(0.81)
3	1.017	0.991	1.214	1.046	1.190	1.167	0.996	0.925	0.929	1.067	1.070	0.886
	(0.85)	(0.94)	(0.26)	(0.13)	(0.03)	(0.37)	(0.95)	(0.27)	(0.5)	(0.26)	(0.24)	(0.25)
6	1.022	1.003	1.032	1.123	1.180	1.345	0.968	0.948	0.769	1.039	1.021	0.809
	(0.88)	(0.98)	(0.86)	(0.13)	(0.2)	(0.09)	(0.73)	(0.53)	(0.07)	(0.42)	(0.77)	(0.08)
9	1.046	1.059	1.161	1.171	1.189	1.379	0.897	0.879	0.800	1.052	0.980	0.851
	(0.82)	(0.76)	(0.37)	(0.13)	(0.18)	(0.05)	(0.47)	(0.42)	(0.16)	(0.42)	(0.82)	(0.2)
12	0.966	0.948	1.000	1.246	1.287	1.481	0.932	0.975	0.848	1.088	1.011	0.978
	(0.87)	(0.82)	(1)	(0.06)	(0.02)	(0.02)	(0.65)	(0.89)	(0.32)	(0.44)	(0.92)	(0.86)
18	0.785	0.729	1.400	1.443	1.697	2.053	1.116	1.183	0.975	1.191	1.000	1.071
	(0.18)	(0.24)	(0.01)	(0)	(0)	(0)	(0.39)	(0.31)	(0.85)	(0.37)	(1)	(0.59)
24	0.691	0.613	1.192	1.648	1.872	43.000	1.419	1.546	1.367	0.942	0.884	1.317
	(0.12)	(0.07)	(0.13)	(0)	(0)	(0)	(0.02)	(0.06)	(0.03)	(0.59)	(0.45)	(0.02)
No. of Stat. < 1	4	6	0	2	0	0	6	6	6	1	3	6
(of which sig.)	0	1	0	0	0	0	0	1	1	0	0	1
No. of Stat. >= 1	4	2	8	6	8	8	2	2	2	7	5	2
(of which sig.)	0	0	2	3	5	6	1	1	1	0	0	1

Table - 12. Volatility of Oil Prices and Forecast of the Oil-Price Exchange Rate Model

		1988M1 -1992M12	1993M1 -1997M12	1998M1 -2002M12	2003M1 -2009M7
US	Volatity of Oil Prices	5.9	2.6	6.6	21.3
	Volatility rank	3	4	2	1
	Performance rank	3	1	4	2
Japan	Volatity of Oil Prices	637.2	365.0	679.3	2683.9
	Volatility rank	3	4	2	1
	Performance rank	1	2	3	4
Germany	Volatity of Oil Prices	8.6	4.4	16.1	26.6
	Volatility rank	3	4	2	1
	Performance rank	2	1	4	3
UK	Volatity of Oil Prices	3.1	1.7	4.9	10.0
	Volatility rank	3	4	2	1
	Performance rank	3	2	4	1
Canada	Volatity of Oil Prices	6.2	3.5	10.1	19.4
	Volatility rank	3	4	2	1
	Performance rank	1	4	2	3
Norway	Volatity of Oil Prices	31.8	16.2	62.0	105.2
	Volatility rank	3	4	2	1
	Performance rank	3	1	4	2
Mexico	Volatity of Oil Prices	85.7	64.2	61.8	211.1
	Volatility rank	2	3	4	1
	Performance rank	3	1	4	2
Turkey	Volatity of Oil Prices	10.0	7.0	15.7	17.6
	Volatility rank	3	4	2	1
	Performance rank	2	4	1	3