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CROSS-COUNTRY EVIDENCE ON THE RELATION BETWEEN EQUITY PRICES AND THE CURRENT ACCOUNT

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Cross-country evidence on the relation between equity prices and the current account

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Abstract

This paper explores the relationship between equity prices and the current account for 17 industrialized countries in the period 1980-2007. Based on a panel vector autoregression, I compare the effects of equity price shocks to those originating from monetary policy and exchange rates. While monetary policy shocks have a limited impact, shocks to equity prices have sizeable effects. The results suggest that equity prices impact on the current account through their effects on real activity and exchange rates. Furthermore, shocks to exchange rates play a key role as well.

Keywords: current account fluctuations, equity prices, panel vector autoregression

JEL-Codes: C33, E44, F32

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1 Introduction

The determinants of current account fluctuations have been discussed extensively in the academic literature in recent years. One reason is that the dispersion in current account positions has never been so large as today. This triggered worries that an unwinding of global imbalances could cause a severe global financial crisis. In the wake of the current financial crisis 2007-2009, it is even more important to understand the sources of these imbalances and the likely adjustment mechanisms. Particularly, the role of equity prices is of interest and is thus the central issue of this paper. The existing literature on the link between equity prices and the current account is small and concentrates on individual countries. In contrast, I extend the analysis to a broad set of OECD countries[1] and compare the effects of equity price shocks to those originating from monetary policy and exchange rates.

Since the US contributed substantially to the emergence of global imbalances, many authors focus on the US in their analysis. While some point to low private savings in the US as a main driver of these imbalances (Krugman (2007)), others investigate the role of public savings (Erceg et al. (2005), Corsetti and Müller (2006)). From a simple accounting perspective, budget and current account deficits move in the same direction. Thus, the swing of the US fiscal position from surplus to deficit during the Bush era may have accelerated the deterioration of the US current account. However, the two aforementioned papers find little impact of fiscal shocks on the current account and reject what is known as the "twin deficit" hypothesis.[2] Another camp identifies productivity shocks as a main determinant of the current account (Bussière et al. (2005), Corsetti et al. (2006), Bems et al. (2007)). Country-specific productivity shocks raise relative consumption, deteriorate net exports, raise the relative price of nontradables and deteriorate the trade balance (and thus the current account). Corsetti et al. (2006) find evidence that this effect is particularly persistent for the US. A third strand focuses on the role exchange rates play in restoring external balance for countries with large external deficits (Obstfeld and Rogoff (1995), Blanchard et al. (2005)). A common result of this literature is that a large and steady depreciation of the exchange rate is needed to rebalance the current account (Krugman (2007)).

Despite the vast literature on the sources of current account fluctuations, it is striking that only few authors discuss the contribution of equity price shocks to the emergence of global imbalances. Some notable and recent exceptions are Fratzscher et al. (2007), Barnett and Straub (2008) and Fratzscher and Straub (2009). The motivation is the following. While the US reports remarkable current account deficits, many countries, particularly from emerging Asia and the Middle East, run current account surpluses of similar magnitude. Having recovered from the 1997-1998 Asian crisis, the demand for foreign exchange

[1] The countries included are: Australia, Austria, Belgium, Canada, France, Germany, Italy, Japan, Korea, the Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, the United Kingdom and the United States.

[2] Kim and Roubini (2008) find even evidence for a "twin divergence", i.e., when fiscal accounts worsen, the current account improves and vice versa.

reserves was huge among Asian countries. Since the US financial market is the largest and most liquid market in the world, a large fraction of these reserves were invested in US dollar denominated assets, particularly in US government bonds. Furthermore, the surge in oil prices created large surpluses among the oil-exporting countries that were in turn reinvested in US bonds and equity. In addition, the lack of well functioning capital markets in the emerging world spurred the demand for US assets.[3]

Consequently, one would expect that the (relative) attractiveness of a country's financial market is an important determinant of international capital flows. If a country experiences a favorable equity price shock more funds are allocated to the country, the exchange rate is likely to appreciate and the current account worsens. Furthermore, the increase in equity prices may impact on real activity through wealth effects on consumption and balance sheet effects on investment. Both raise the demand for imports and deteriorate the current account.

What is meant by an equity price shock is the following. Because equity prices are forward-looking, an equity price shock is interpreted as a shift in expectations about future economic conditions. For example, market participants expect productivity to rise in the future or the share of the country's output in the world to increase (Engel and Rogers (2006)). One may also think of an equity price shock in the form of a rational bubble (Kraay and Ventura (2005)).

Fratzscher et al. (2007) find that shocks to equity prices have large and persistent effects on the US trade balance. Using a Bayesian structural VAR, they measure the impact of a 10% increase in equity prices to be 0.9% over 10-15 quarters and find this effect to be larger than that of the exchange rate. In a more recent study Fratzscher and Straub (2009) extend the analysis to the G7 economies and obtain again evidence of a significant impact of equity price movements on the trade balance. However, the response of the trade balance to the equity price shock varies substantially across countries suggesting that a strong response is probably unique to the US.

This paper contributes to the existing literature in the following way. Using a panel vector autoregression (panel VAR), I identify the impact of monetary policy, equity price and exchange rate shocks on the current account. The panel set-up allows me to filter out country-specific effects and to study the average effects of the three shocks. The results suggest that both equity and exchange rate shocks have a significant impact, while monetary policy shocks have virtually no effects.

The rest of the paper is organized as follows. Section 2 explains the empirical model. Section 3 describes the data. An impulse response analysis and a forecast error variance decomposition are presented in section 4. Robustness checks are discussed in section 5 and section 6 concludes.

[3] This idea is put forward by Bernanke (2005). He argues that a "saving glut" in Asia and among oil-exporting countries is a main driver of the US current account deficit.

2 Methodology

The analysis is based on a panel VAR model of the form:

$$Y_{it} = B_i(L) Y_{i,t-1} + C_i(L) D_t + u_{it} \tag{1}$$

where $i = 1, 2, ..., N$; $t = 1, 2, ..., T$; Y_{it} is a $G \times 1$ vector of endogenous variables for each country i, B_i are $G \times G$ matrices in the lag operator L, D_t is a $K \times 1$ vector which may include deterministic variables (e.g., a constant, a time trend or a dummy) or common exogenous variables (e.g., oil prices), C_i are $G \times K$ matrices in the lag operator L and u_{it} is a $G \times 1$ vector of random disturbances with mean zero and country-specific variance σ_i^2.

I include seven endogenous variables for each country: real GDP, consumer prices, nominal short-term interest rates, nominal long-term interest rates, nominal equity prices, the real effective exchange rate and a current account to GDP ratio. All variables are expressed in logs, except the interest rate variables and the current account to GDP ratio. Since the current account is measured with respect to the "rest of the world" it seems appropriate to incorporate all other endogenous variables in relative terms. This is achieved in the following way. First, I construct bilateral trade weights for each country with all other countries in the panel and each period. Particulary, the weight that is attached to country j for country i in period t is:

$$\omega_{i,j,t} = \frac{imp_{i,j,t} + exp_{i,j,t}}{\sum_{j=1}^{N} (imp_{i,j,t} + exp_{i,j,t})} \tag{2}$$

where $imp_{i,j,t}$ is the amount of goods and services (in millions of US dollars) that is imported by country i from country j in period t, $exp_{i,j,t}$ is the amount of goods and services that is exported by country i to country$j$ in period t and $\sum_{j=1}^{N} (imp_{i,j,t} + exp_{i,j,t})$ is the total sum of imports and exports of country i with all other countries in period t. Obviously, $\omega_{i,j,t} = 0$ for $i = j$. Thus, $\omega_{i,j,t}$ captures the importance of country j for country i with respect to trade. Second, I calculate foreign variables for each country i as:

$$x_{it}^{*} = \sum_{j=1}^{N} \omega_{i,j,t} x_{jt} \tag{3}$$

Using time-varying rather than fixed weights allows me to control for changing patterns in global trade. I proceed in this way for (log) real GDP y_{it}, (log) consumer prices p_{it}, nominal short-term interest rates r_{it}^{s}, nominal long-term interest rates r_{it}^{l} and (log) nominal equity prices q_{it}. But not for the (log) real effective exchange rate $REER_{it}$ and the current account to GDP ratio ca_{it} since both are already measured relative to major trading partners.[4] Finally, I obtain relative variables by substracting foreign from domestic variables.[5] [6] Hence, the vector of endogenous variables becomes

$$Y_{it} = \left[y_{it} - y^*_{it} \quad p_{it} - p^*_{it} \quad r^s_{it} - {r^s}^*_{it} \quad r^l_{it} - {r^l}^*_{it} \quad q_{it} - q^*_{it} \quad REER_{it} \quad ca_{it} \right]' \tag{4}$$

One purpose of the paper is to evaluate the effects of monetary policy shocks on the current account and therefore all relevant channels through which monetary policy impacts on the economy are included. Monetary policy impacts on short and long-term interest rates and thus on the term structure. Furthermore, monetary policy is transmited to the economy through equity prices and exchange rates. Since movements in equity prices have contributed to the development of global imbalances in the last two decades or so, I include nominal equity prices. Finally, the real effective exchange rate is added to the model in order to have a measure of the external competitiveness of the country under study.

The vector of common exogenous variables D_t includes the US dollar price of oil p^{oil}_t and a constant for each country. The oil price is considered for several reasons. First, it is a well known shortcoming of VAR analyses that inflation expectations cannot be taken into account explicitly. Including oil or commodity prices may help to overcome this problem since both are expected to be correlated with inflation expectations. Second, some of the countries in the panel are net oil exporters (notably Canada, Norway and the UK) and are significantly influenced by movements in the price of oil. Third, I will not control for cross-section dependence in the panel and including an observed common factor is expected to reduce ineffiences that may arise in this context.

Preliminary estimation of individual VAR models suggests that a lag order of four for the endogenous variables is optimal, using lag order selection criteria like AIC, SBC or likelihood ratio tests and is thus set to four for all countries. Furthermore, the oil price enters contemporaneously and with one lag.

Following Pesaran and Smith (1995), I assume that the B_i and C_i matrices vary across countries according to the following random coefficient model[7]:

$$B_{pi} = B_p + \eta_{1,p,i}, \qquad C_{qi} = C_q + \eta_{2,q,i} \tag{5}$$

where B_p and C_q are $G \times G$ and $G \times K$ constant matrices, $\eta_{1,p,i}$ and $\eta_{2,q,i}$ are $G \times G$ and $G \times K$ random matrices and p and q are the respective lag orders. Furthermore, $\eta_{1,p,i}$ and $\eta_{2,q,i}$ are assumed to

[4] Thus, the procedure is similar to the one employed by Pesaran et al. (2004) and Dees et al. (2007) in a Global VAR context.

[5] Fratzscher et al. (2007) follow a similar approach for the US and specify the variables relative to the rest of the world. However, they use weights based on global GDP shares rather than trade weights.

[6] An alternative strategy would be to include domestic and foreign variables separately. But this is computationally not feasible.

[7] The random coefficient model is introduced by Swamy (1970).

be distributed independently of u_{it} with zero mean and constant covariance matrices Ω_{1p} and Ω_{2q} - i.e. $vec\left(\eta_{1,p,i}\right) \sim iid\left(0, \Omega_{1p}\right)$ and $vec\left(\eta_{2,q,i}\right) \sim iid\left(0, \Omega_{2q}\right)$.

As long as the time series dimension T is sufficiently large to run individual time series regressions, the panel VAR model can be estimated in several ways: first, by stacking the data and using standard pooled estimators such as the random or fixed effects estimator; second, by estimating individual VAR models for each country seperately and averaging the estimated coefficients across countries. The second approach is proposed by Pesaran and Smith (1995) and is known as the mean group estimator. Provided the panel is not only large with respect to time, but also homogeneous (i.e. $\eta_{1,p,i} = \eta_{2,q,i} = 0$ for all i), all estimators yield consistent and unbiased estimates of the coefficients for N being large as well. But if the coefficients differ across countries (i.e. $\eta_{1,p,i} \neq \eta_{2,q,i} \neq 0$ for some i), the random and fixed effects estimators give inconsistent and potentially misleading estimates of the coefficients.[8] The mean group estimator, however, is consistent even in the presence of parameter heterogeneity for N and T being large. Since the cross-sectional and the time series dimension are both sufficiently large ($N = 17$ and $T = 112$) and some degree of parameter heterogeneity across countries seems likely, the mean group estimator is preferred and the coefficient matrices are estimated as:

$$\hat{B}_p = \frac{1}{N}\sum_{i=1}^{N}\hat{B}_{pi}, \qquad \hat{C}_q = \frac{1}{N}\sum_{i=1}^{N}\hat{C}_{qi} \tag{6}$$

for $p = 1, 2, ..., p^{max}$ and $q = 0, 1, ..., q^{max}$. Pesaran and Smith (1995) show that the mean group estimator converges relatively fast and that $\hat{B}_p$ and $\hat{C}_q$ are appropriate measures of the average effects of $Y_{i,t-p}$ and D_{t-q} on Y_{it}.

3 Data

The data are either from the OECD Main Economic Indicators (MEI) data base or from IMF's International Financial Statistics (IFS). Data on real GDP are obtained from the IMF with the exception of Canada and Italy where data from the OECD are used. For New Zealand real GDP data for the early 1980s are not available on a quarterly basis. Therefore, I interpolate annual real GDP with the Chow and Lin (1971) procedure, using industrial production as an indicator series, and link this series to the quarterly OECD series starting in 1982Q2.

Data on consumer prices are from the OECD. Consumer prices and real GDP are deseasonalized using the X-11 filter. The US dollar price of Brent crude oil is taken from the OECD. Short-term interest

[8] This problem arises because incorrectly ignoring coefficient heterogeneity induces serial correlation in the error terms and leads to inconsistent coefficient estimates in models with lagged dependent variables. It does not disappear even if $T \rightarrow \infty$. Thus, this inconsistency is different from that suffered by the fixed effects estimator in small T panels as $N \rightarrow \infty$ (see Nickell (1981)) and the solutions proposed in the literature do not solve this problem.

rates are 3-month rates and, where available, I use Treasury bill rates from the IMF. For Australia, Austria, Germany, Japan, Korea, the Netherlands, Spain and Sweden, Treasury bill rates are not available and I use money market rates instead. Furthermore, in case of New Zealand and Norway interbank rates from the OECD are used. For the euro area economies I replace domestic short-term interest rates by the 3-month EURIBOR rate after 1998. Data on long-term interest rates are taken from the IMF for Austria, Italy, Japan, Korea, Norway and Sweden; for all other countries from the OECD. In each case the long-term interest rate is the yield on a 10-year government bond. Equity prices are from the IMF, except for Switzerland and the United Kingdom where the data come from the OECD. For all countries a broad share price index is used. The real effective exchange rate is a trade weighted index, adjusted for relative consumer prices and comes from the IMF. In case of Korea the index is from the OECD.

In order to obtain current account to GDP ratios, I divide the nominal current account by nominal GDP of the same period. Current account data come from the IMF, except for Germany and Switzerland where the data are from the OECD. For Norway missing observations for 1992Q1-1993Q4 are replaced with data from Statistics Norway. Nominal GDP is from the IMF and in case of New Zealand is again interpolated from annual to quarterly frequency for the early 1980s. Since the current account is denominated in US dollar, it is converted to domestic currency using bilateral US dollar market exchange rates from the OECD, with the exception of Korea where data are from the IMF.

Finally, the bilateral trade flows that are used to construct trade weights are from the OECD. Unfortunately, there are missing values in trade flows between Belgium, Korea and New Zealand prior to 1988. I deal with this problem by setting trade flows between these countries equal to zero for all years up to 1988. Since trade between the three countries was limited until the late-1990s, it is unlikely that this contaminates the trade weights.

4 Empirical results

4.1 Unit root test results

Table 1 shows the results of augmented Dickey-Fuller (ADF) unit root tests for the endogenous variables in level specification. The ADF regressions contain a constant and a time trend. The lag order for the first differences is set equal to five.[9] Similar test results for the endogenous variables in first differences are reported in Table 2. In this case the ADF regressions include a constant only and the lag length is four. There is strong evidence that nearly all of the variables in the panel are integrated of order one. In fact, for most of the countries the null of a unit root in the level cannot be rejected at a 5% significance level for any variable. In contrast, the test statistics for the endogenous variables in first

[9]The results are insensitive to variations in the lag length.

differences are, with only a few exceptions, highly significant. Thus, I assume that the endogenous variables are $I(1)$.[10] The same applies to the oil price.[11] Thus, it would be a valid strategy to estimate the panel VAR in first differences. However, differencing the variables would destroy cointegrating relations in the model. Therefore, I estimate the panel VAR in levels, taking any cointegrating relationships implicitly into account. Indeed, Johansen cointegration tests indicate that there is evidence of at least one cointegrating vector, implying that the individual country models can be estimated in levels.

4.2 Impulse responses

4.2.1 Identification

The panel VAR model is estimated over the period 1980Q1-2007Q4. A common way of analyzing the dynamics of the system is to calculate impulse response functions. It is assumed that the reduced form errors u_{it} are linked to the structural innovations ϵ_{it} in the following way:

$$u_{it} = A_i \epsilon_{it} \tag{7}$$

In order to achieve identification, I impose the restriction that the A_i matrices are lower triangular. Such a recursive identification scheme is frequently employed in the literature and leaves it to the researcher to specify the instantaneous causal ordering of the variables. In what follows, I assume that the variables in the system are ordered as in (eq. 4) and will refer to this as benchmark identification.

Monetary policy shocks, defined as innovation in the relative short-term interest rate ($r^s_{it} - r^{s*}_{it} > 0$), do not have any contemporaneous impact on either real GDP or consumer prices. Both variables respond with a lag of one quarter to changes in monetary policy. However, the financial market variables (long-term interest rates, equity prices and the exchange rate) are allowed to respond immediately to changes in the monetary policy instrument. This identification scheme is often used in the analysis of monetary policy transmission in an open economy context.[12] Equity price shocks are defined as innovation in the relative equity price measure ($q_{it} - q^*_{it} > 0$). Again, real GDP and consumer prices respond with a lag. Furthermore, it seems likely that monetary policy takes changes in equity prices into account since they influence output and prices. However, one would not expect that monetary policy reacts instantaneously to changes in equity prices but only if they rise or fall for a longer period of time. The same argument

[10] Levin et al. (2002) and Im et al. (2003) panel unit root tests support the idea that the series are $I(1)$.

[11] The test results for p^{oil} are -0.48 (level) and -5.53 (first difference), respectively.

[12] Eichenbaum and Evans (1995) and Grilli and Roubini (1996) use similar recursive identification schemes. Faust and Rogers (2003) and Scholl and Uhlig (2005) identify monetary policy shocks in an open economy framework by imposing shape or sign restrictions on the impulse response functions.

applies to the exchange rate. Hence, both variables are ordered after real GDP, consumer prices and short-term interest rates. Within the block of financial market variables an appropriate ordering is, however, unclear. But it turns out that the impulse response functions are robust to alternative ordering schemes. Therefore, I order the financial market variables as: first, long-term interest rates; second, equity prices; and third, the real effective exchange rate. Exchange rate shocks are defined as innovation in the real effective exchange rate ($REER_{it} > 0$).

Finally, the current account to GDP ratio is ordered last. This imposes the restriction that the current account responds immediately to changes in other variables, but these react with a lag to a change in the current account. This seems plausible since the current account is nothing else than the accumulation of foreign assets or debt (at least if one abstracts from valuation effects) and one would not expect that macroeconomic variables react to changes in the stock of net foreign assets within the period.

4.2.2 Error bands

Before I discuss the impulse responses, I explain how the error bands for the impulse reponses are obtained. Since the underlying time series are all integrated of order one and thus have stochastic trends, bootstrapping procedures such as residual based methods (see Lütkepohl (2000) or Benkwitz et al. (2001)) or the standard block-bootstrap (see Künsch (1989)) are not directly applicable since they demand stationarity. Paparoditis and Politis (2002) propose a modification to the standard block-bootstrap that accounts for the changing stochastic structure of the time series. The basic idea of the local block-bootstrap is to only resample blocks that are close to each other, i.e a block that starts at time t can only be replaced with blocks whose starting point is close to t. However, even if the stochastic structure is changing smoothly the realization of a local block-bootstrap pseudo replication typically exhibits strong discontinuities where the independent bootstrap blocks join. In order to avoid this problem, Paparoditis and Politis (2001) suggest to force the sample path to be continuous. This can be done by shifting the blocks up and down in such a way that the bootstrap series starts off at the same point as the original series and that the bootstrap sample path is continuous. Comparing the bootstrap series with the original series leads to the conclusion that the bootstrap series may well be generated by the same probability mechanism as the original series. Thus, such a continuous-path block-bootstrap algorithm is successful in imitating important features of the original series. In what follows, I will use a version of the continuous-path block-bootstrap procedure that takes the $I(1)$ property of the time series explicitly into account. The idea is simple and intuitive and is proposed by Politis (2003). The algorithm is outlined below.

Suppose a time series z_t is non-stationary; $t = 1, 2, ..., T$; and z_0 is available. Then

- calculate the series of stationary first differences Δz_t, where $\Delta z_t = z_t - z_{t-1}$
- perform a block-bootstrap of the first differences Δz_t, i.e. randomly draw blocks of size b with replacement from $\Delta z_1, \Delta z_2, ..., \Delta z_T$, yielding $\Delta z_1^*, \Delta z_2^*, ..., \Delta z_T^*$
- construct a bootstrap pseudo-series for z_t by "integrating" the Δz_t^* - i.e. - by letting $z_t^* = z_0 + \sum_{i=1}^{t} \Delta z_i^*$
- use the bootstrap pseudo-series z_t^* to re-estimate the coefficients of the VAR (or panel VAR) model
- calculate the bootstrap impulse response functions
- repeat the previous four steps a large number of times

This non-parametric bootstrap imposes only a minimum set of restrictions on the data, but requires the user to specify which block size b to use. The block-bootstrap literature recommends to take b small with respect to the sample size T. But since the original time series is expected to be (weakly) dependent, b should also reflect the degree of dependence. I choose to set $b = 4$ because this size fullfills both criteria reasonable well.[13] Letting the block size vary between 2 and 12 produces similar error bands.

Each bootstrap replication is initialized with the first observation z_0 of the respective original time series. The whole procedure is repeated 1,000 times and on the basis of the empirical distribution of the impulse response functions confidence intervals are calculated as

$$CI = \left[\hat{\phi} + 1.645 \times \left(var\left(\hat{\phi}^*\right)\right)^{\frac{1}{2}}, \quad \hat{\phi} - 1.645 \times \left(var\left(\hat{\phi}^*\right)\right)^{\frac{1}{2}}\right] \tag{8}$$

where $\hat{\phi}$ are the impulse responses based on the original data and $\hat{\phi}^*$ are the bootstrap counterparts. Thus, the impulse responses show the original response when the VAR coefficients are fixed at their respective OLS point estimates and a 90% confidence band.

4.2.3 Monetary policy shocks

Figure 1 shows the responses of real GDP, consumer prices, short-term interest rates, long-term interest rates, equity prices, the exchange rate and the current account to a one standard error monetary policy shock, corresponding to an increase in the relative short-term interest rate of about 50 basis points. The effect on the short-term interest rate settles at around zero after two and a half years. Long-term interest rates rise immediately, however, the initial impact is only one third of that of the short-term interest rate.

[13] The autocorrelation functions settle around zero after 3-6 quarters.

Long-term interest rates fall thereafter and the response is zero after two and a half years as well. Real GDP contracts significantly following the monetary policy shock and reaches its trough after two years, before it recovers. Consumer prices rise on impact, displaying a “price puzzle”, but start to fall after around two years. Equity prices fall sharply in response to the monetary policy tightening, but recover quickly. The trough is reached after four quarters. Furthermore, the response of the exchange rate exhibits a puzzle as well. The domestic currency depreciates on impact and it takes nearly one year until the effect turns positive.[14] But since consumer prices are used to construct the exchange rate, and consumer prices show a “price puzzle”, it is not suprising that the “price puzzle” is evident in the response of the exchange rate, too.

Finally, the response of the current account is ambiguous. It is slightly negative on impact, but quickly changes sign and is above the initial level after seven quarters. After about three years it settles at around zero. Moreover, the response is never significantly different from zero. Consequently, it seems implausible that loose monetary policies contribute to current account deficits. While an expansionary monetary policy shock raises domestic demand and deteriorates net exports, it also depreciates the domestic currency and improves net exports. The results of the impulse response analysis suggest that the overall effect on net exports, or more exactly, the current account, is about zero.

4.2.4 Equity price shocks

Figure 2 shows the responses to a one standard error shock that raises relative equity prices by more than 4% initially. The rise in equity prices is followed by a significant and long lasting increase in both real GDP and consumer prices.[15] In response to the increase in real activity and rising prices the monetary policy authority is tightening. In addition, long-term interest rates react positively as well. The effect on the exchange rate is, however, unclear. While the point estimate suggests that the domestic currency appreciates, the uncertainty surrounding the impulse response is quite high. Finally, the current account reacts immediately and reaches a trough after eight quarters. Thereafter, the current account improves and external balance is restored after around five years. The maximum impact of the 4% (a 10 %) increase in equity prices on the current account is -0.12% (-0.3%). Hence, the results are in line with Fratzscher and Straub (2009) who report responses of the trade balance to an equity price shock (of size 10%) between -1.02 (for Germany) and 0.28 (for the UK) after eight quarters.

[14] The exchange rate is defined in such a way that an increase means an appreciation.

[15] Thus, an equity price shock is distinct from a technology shock with respect to the behavior of prices. While an equity price shock induces a positive correlation between output and prices, a favorable technology shock leads to higher output but lower prices. See e.g. Adolfson et al. (2005) or Galí and Monacelli (2005).

4.2.5 Exchange rate shocks

Figure 3 shows the responses to a one standard error innovation in the exchange rate. The exchange rate appreciates by 1.8% on impact, falls thereafter and finally settles around zero after 12 quarters. The appreciation is associated with a loss of external competitiveness and net exports are likely to fall. Thus, real GDP contracts significantly following the exchange rate shock. Furthermore, since the appreciation lowers import prices, consumer prices fall. Consumer prices reach a through after around eight quarters. The monetary policy authority reacts to the fall in real GDP and consumer prices by lowering short-term interest rates and long-term interest rates match the behavior of short-term interest rates nearly one-to-one. In addition, equity prices fall immediately and are well below their initial level after five years. Finally, the current account falls sharply in response to the appreciation. It reaches a trough right in the first quarter and then improves. However, the response is negative for the next five years. The effect of the exchange rate shock on the current account is strong, significant and long lasting. A 10% increase in the exchange rate depresses the current account by 0.4%, more than the impact of an equity price shock of similar magnitude.

4.3 Forecast error variance decomposition

The forecast error variance decomposition shows the proportion of the unanticipated changes of a variable that can be attributed to own innovations and to innovations to other variables in the system. Table 3 shows the variance decomposition of the current account. The contribution of the structural innovations is reported up to 24 quarters following the shock. For instance, about 77% of the 4-step ahead forecast error variance of the current account is due to own innovations. This number decreases considerably over time and is 38% after six years. Moreover, innovations in prices and long-term interest rates contribute less than 8% over all forecast horizons. About 13% of the forecast error variance of the current account is accounted for by innovations in real GDP. For any forecast horizon, monetary policy shocks contribute less than 8%. This is in line with the results of the impulse response analysis. Monetary policy shocks are probably not a main source of fluctuations in the current account. This is in contrast to the findings of Barnett and Straub (2008) who identify the US federal funds rate as a main source of the variability in the US current account. They estimate the contribution of monetary policy shocks to the forecast error variance to be 62% at low forecast horizons and 41% at a seven year forecast horizon. Furthermore, Fratzscher et al. (2007) find also evidence that monetary policy exerts influence. However, their numbers are considerably smaller and comparable to those stemming from my panel VAR.

The results are different for innovations in equity prices and the exchange rate. For long-term forecasts, 16% and 12% of the forecast error variance is accounted for by equity price and exchange rate innovations, respectively. Thus, both variables contribute substantially to the forecast error variance of the current account and their joint contribution is nearly as large as the contribution of all other variables together (not taking own innovations into account). Fratzscher et al. (2007) instead report a much smaller impact of the exchange rate on the US trade balance. Only a tiny fraction of the variability can be attributed to exchange rate shocks at long-term forecast horizons. Exchange rate movements appear to be less important for the US than for other countries. This does not come as a surprise since the US is a large and rather closed economy. However, most countries in my panel are small, open and thus sensitive to exchange rate movements. But the results reconcile with the notion that equity prices explain a considerable part of current account fluctuations. Though the effect is smaller than typically found for the US, it is nevertheless remarkable.

5 Robustness

It is useful to evaluate how sensitive the results are to variations in identification. To do so, I estimate the 7-variable panel VAR model for all 5,040 possible Cholesky orderings. The procedure is agnostic with respect to the appropriate ordering of the variables and thus conservative in measuring identification uncertainty. Figure 4 shows the responses of the current account to a monetary policy, an equity price and an exchange rate shock. The top of the shaded area represents the maximum response for each quarter and the lower end corresponds to the minimum. The shape of the respones is the same as when using the benchmark identification, while the uncertainty surrounding the point estimates is moderate, suggesting that the results are independent of the restrictions imposed on the covariance matrix.[16]

Figure 5 delivers the joint distribution of the peak and its altitude for the current account. Following a monetary policy shock, the current account improves by 0.05% after 7-9 quarters. Thus, monetary policy shocks appear to impact only moderately on the current account. In contrast, equity price shocks have sizeable effects. Following an equity price shock, the current account worsens by more than 0.1% after 9-11 quarters. The distribution is sharply peaked, which leads to the conclusion that this results holds regardless of the identification scheme employed. Apparently, things are different in case of an exchange rate shock. There is considerable mass on an early and strong as well as on a late and somewhat milder deterioration. This is the result of the w-shaped response of the current account to an exchange rate shock. Depending on whether one allows the current account to respond instantaneously or not, the peak deterioration is either 0.1% after 1-2 quarters or 0.08% after 7-11 quarters.

Thus, I conclude that the responses of the current account to both monetary policy and equity price shocks are robust to different identification schemes. With respect to the exchange rate shock, I find that the location of the peak deterioration is sensitive to changes in identification, but not the size of the peak.

[16]In fact, the covariance matrix is nearly diagonal and thus different identification schemes lead to similar results.

6 Conclusion

In this paper, I examine the role of shocks to monetary policy, equity prices and exchange rates in explaining current account fluctuations. While a considerable fraction of the existing literature focuses on individual countries, I extend the analysis to a set of 17 industrialized economies. Based on a panel VAR model using data on real GDP, consumer prices, short and long-term interest rates, equity prices, exchange rates and the current account, I find a small role for monetary policy shocks. This finding does not square with the empirical evidence for the US, but can be attributed to the behavior of the exchange rate which mitigates the effects of monetary policy shocks, particularly for small open economies. However, equity price shocks are presumably a main driver of current account fluctuations. They impact on the current account through their effects on real activity and exchange rates. While their impact on exchange rates is small and insignificant, they have considerable effects on real activity. But since the relationship between nominal equity prices and real activity is discussed controversially in the literature, I am careful in interpreting this result. From my perspective it would be interesting to investigate the transmission from nominal equity prices to real consumption and investment in more detail and I leave this question unanswered for future research. To conclude, even though the influence of equity price shocks is remarkable, the results suggest that they are probably not the most important determinant of current account fluctuations. Exchange rates play a key role as well.

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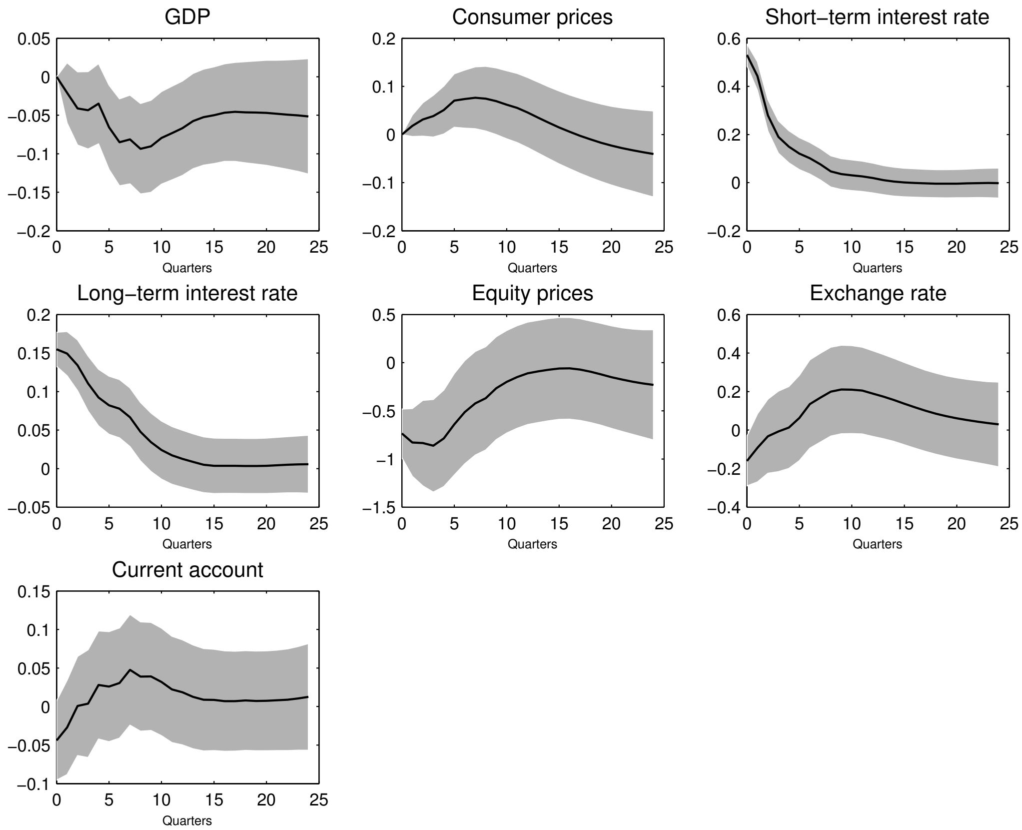

Figure 1: Monetary policy shock. Tables show responses to one standard error innovation in the relative short-term interest rate. Solid lines are responses when VAR coefficients are fixed at their OLS point estimates. Shaded areas are 90% confidence bands obtained by a non-parametric bootstrap with 1,000 replications.

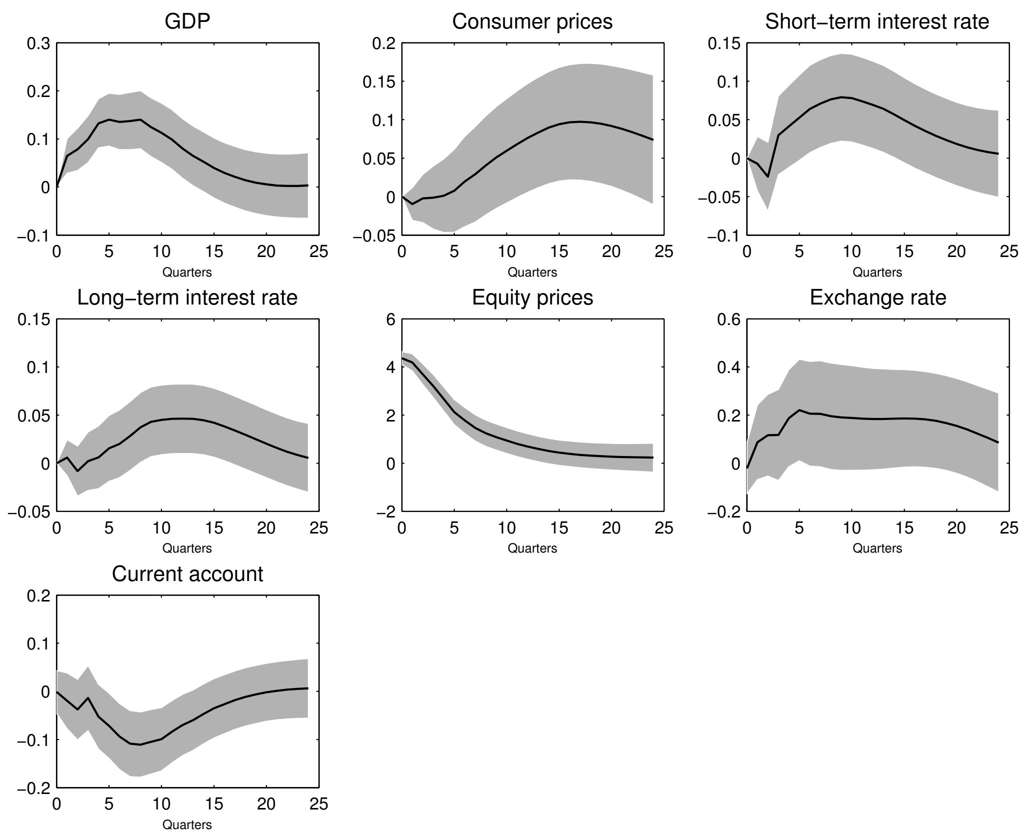

Figure 2: Equity price shock. Tables show responses to one standard error innovation in the relative equity price measure. Solid lines are responses when VAR coefficients are fixed at their OLS point estimates. Shaded areas are 90% confidence bands obtained by a non-parametric bootstrap with 1,000 replications.

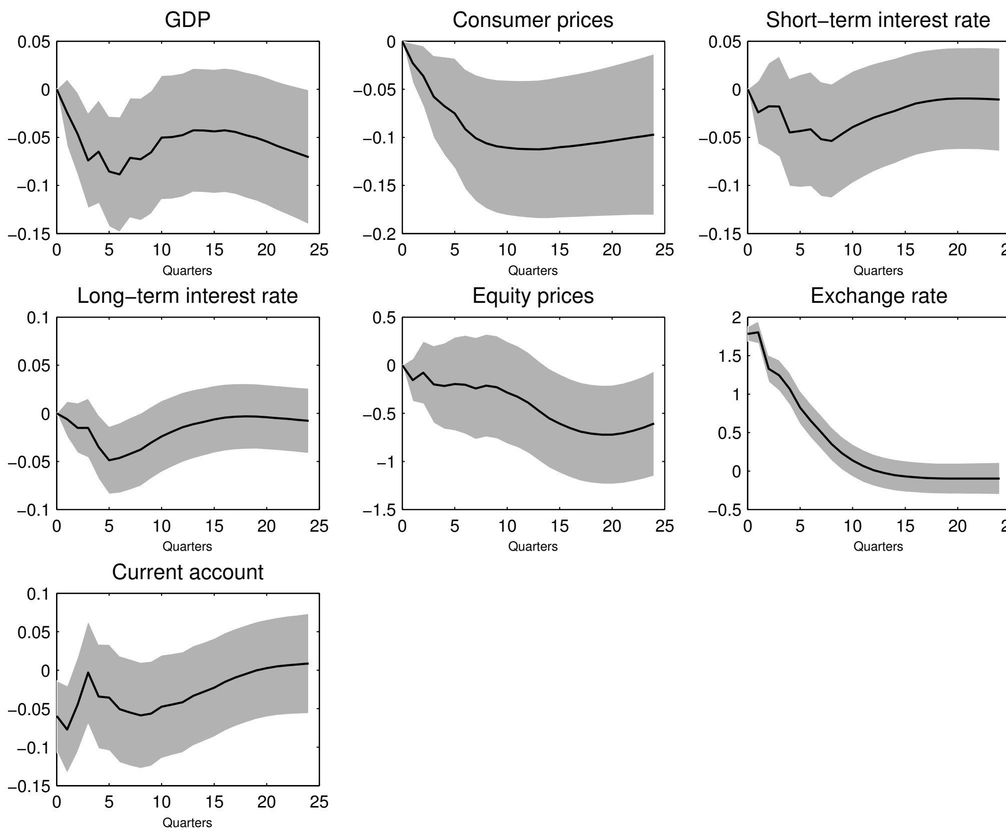

Figure 3: Exchange rate shock. Tables show responses to one standard error innovation in the real effective exchange rate. Solid lines are responses when VAR coefficients are fixed at their OLS point estimates. Shaded areas are 90% confidence bands obtained by a non-parametric bootstrap with 1,000 replications.

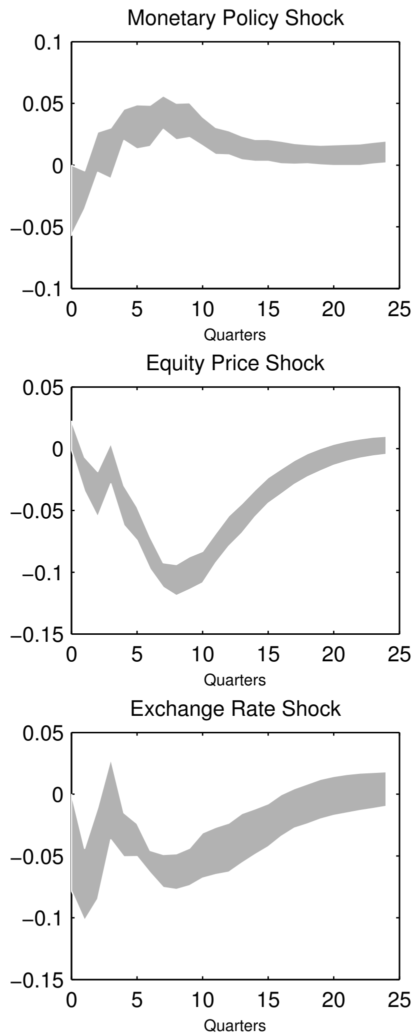

Figure 4: Current account. Distribution of impulse responses. Tables show distribution (shaded area) of responses to monetary policy, equity price and exchange rate shock based on 5,040 different recursive identification schemes.

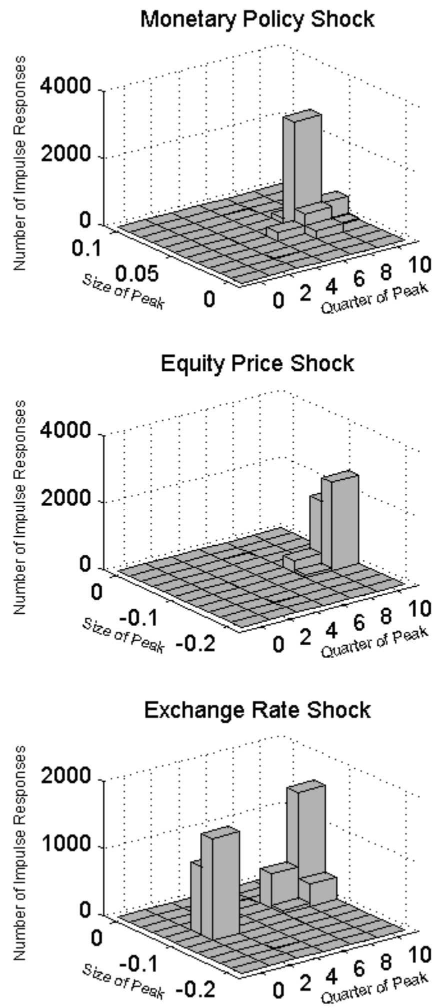

Figure 5: Current account. Distribution of size and location of peak deterioration or improvement. Tables show distribution of size (in %) and location (in quarters) of peak deterioration or improvement, conditional on monetary policy, equity price and exchange rate shock and based on 5,040 different recursive identification schemes.

Table 1: Augmented Dickey-Fuller unit root tests for endogenous variables in levels

	$y-y^*$	$p-p^*$	r^s-r^{s*}	r^l-r^{l*}	$q-q^*$	$REER$	ca
Australia	-2.22	-2.72	-2.98	$-3.61^\dagger$	-3.42	-1.97	$-3.86^\dagger$
Austria	-1.70	-2.36	$-3.66^\dagger$	-2.97	-2.07	-1.74	-2.10
Belgium	-2.34	-3.04	-2.24	-2.55	-2.71	-2.63	-0.89
Canada	-1.71	$-4.00^\dagger$	-3.06	-3.27	-0.55	-0.92	-1.66
France	-2.59	$-4.30^\dagger$	-2.72	-1.59	-2.37	-2.54	-0.72
Germany	-1.46	-2.74	-2.80	-1.27	-2.69	-2.35	-1.07
Italy	-0.87	-3.12	-3.22	-3.00	-2.68	-2.32	-1.73
Japan	-1.81	-2.71	$-4.20^\dagger$	-2.87	-2.25	-1.36	-2.58
Korea	-1.52	-2.31	-2.11	-2.67	-1.83	-2.26	-3.26
Netherlands	-2.82	-2.03	-2.73	-3.01	-1.82	-1.89	-2.67
New Zealand	-2.10	-2.47	-2.21	-2.27	-2.28	-2.69	-2.38
Norway	-2.11	-2.51	-3.29	-2.69	-1.68	-2.23	-3.40
Spain	-1.41	$-4.43^\dagger$	$-4.71^\dagger$	-3.24	-2.76	-2.06	-1.84
Sweden	-0.86	-1.25	-2.83	-2.54	$-4.06^\dagger$	-2.84	-1.44
Switzerland	0.27	-2.07	-1.17	-1.83	-1.91	-2.42	$-3.49^\dagger$
UK	-2.42	-0.77	-2.60	-2.61	-1.74	-2.79	-2.50
US	-2.22	$-3.75^\dagger$	$-3.49^\dagger$	-2.65	-1.67	-2.61	-1.90

Notes: ADF tests include constant and trend. † denotes significance at a 5% level.

Table 2: Augmented Dickey-Fuller unit root tests for endogenous variables in first differences

	$y-y^*$	$p-p^*$	r^s-r^{s*}	r^l-r^{l*}	$q-q^*$	$REER$	ca
Australia	$-4.44^\dagger$	$-3.30^\dagger$	$-4.50^\dagger$	$-3.90^\dagger$	$-4.68^\dagger$	$-3.62^\dagger$	$-4.70^\dagger$
Austria	$-4.83^\dagger$	$-3.92^\dagger$	$-5.15^\dagger$	$-4.90^\dagger$	$-3.67^\dagger$	$-4.69^\dagger$	$-5.79^\dagger$
Belgium	$-5.75^\dagger$	$-4.65^\dagger$	$-6.86^\dagger$	$-5.08^\dagger$	$-6.30^\dagger$	$-3.80^\dagger$	$-6.43^\dagger$
Canada	$-4.47^\dagger$	-2.81	$-5.05^\dagger$	$-4.74^\dagger$	$-5.28^\dagger$	$-3.12^\dagger$	$-6.64^\dagger$
France	$-3.67^\dagger$	-1.97	$-5.18^\dagger$	$-4.60^\dagger$	$-5.82^\dagger$	$-4.15^\dagger$	$-5.28^\dagger$
Germany	$-3.55^\dagger$	-2.13	$-4.46^\dagger$	$-4.92^\dagger$	$-4.73^\dagger$	$-5.45^\dagger$	$-4.42^\dagger$
Italy	$-4.90^\dagger$	-2.12	$-4.65^\dagger$	$-4.77^\dagger$	$-7.06^\dagger$	$-4.62^\dagger$	$-4.96^\dagger$
Japan	$-3.78^\dagger$	$-3.93^\dagger$	$-5.86^\dagger$	$-6.31^\dagger$	$-4.08^\dagger$	$-4.02^\dagger$	$-4.62^\dagger$
Korea	$-4.48^\dagger$	$-5.31^\dagger$	$-5.45^\dagger$	$-5.83^\dagger$	$-4.95^\dagger$	$-4.55^\dagger$	$-4.68^\dagger$
Netherlands	$-3.62^\dagger$	$-2.02^\dagger$	$-4.88^\dagger$	$-3.71^\dagger$	$-3.50^\dagger$	$-4.56^\dagger$	$-5.98^\dagger$
New Zealand	$-5.52^\dagger$	-2.87	$-5.45^\dagger$	$-4.98^\dagger$	$-4.54^\dagger$	$-4.06^\dagger$	$-5.80^\dagger$
Norway	$-3.77^\dagger$	-2.71	$-4.90^\dagger$	$-4.12^\dagger$	$-5.10^\dagger$	$-5.96^\dagger$	$-4.76^\dagger$
Spain	$-4.37^\dagger$	-2.37	$-5.94^\dagger$	$-6.53^\dagger$	$-4.54^\dagger$	$-3.87^\dagger$	$-3.62^\dagger$
Sweden	$-3.71^\dagger$	-2.85	$-6.16^\dagger$	$-6.49^\dagger$	$-4.55^\dagger$	$-4.49^\dagger$	$-6.35^\dagger$
Switzerland	$-5.23^\dagger$	$-3.73^\dagger$	$-6.99^\dagger$	$-5.09^\dagger$	$-3.76^\dagger$	$-4.96^\dagger$	$-4.72^\dagger$
UK	$-3.41^\dagger$	$-3.43^\dagger$	$-6.04^\dagger$	$-5.91^\dagger$	$-5.52^\dagger$	$-5.22^\dagger$	$-5.86^\dagger$
US	$-4.38^\dagger$	$-3.02^\dagger$	$-3.04^\dagger$	$-4.92^\dagger$	$-4.58^\dagger$	$-3.40^\dagger$	$-3.78^\dagger$

Notes: ADF tests include constant only. † denotes significance at a 5% level.

Table 3: Forecast error variance decomposition of current account variable

Horizon	$y - y^*$	$p - p^*$	$r^s - r^{s*}$	$r^l - r^{l*}$	$q - q^*$	$REER$	ca
4	5.73	2.49	4.58	3.44	2.82	4.38	76.57
8	6.40	3.44	6.19	5.04	6.59	7.03	65.30
12	7.19	4.60	6.62	5.67	9.86	9.28	56.77
16	8.70	5.91	6.78	5.89	12.44	10.65	49.63
20	10.73	7.09	6.99	6.03	14.35	11.43	43.37
24	12.82	7.97	7.28	6.22	15.79	11.74	38.18

Notes: Contribution of structural innovations in %. Recursive identification.

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