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

Tim Oliver Berg

Goethe University Frankfurt, Institute for Monetary and Financial Stability (IMFS)

19. May 2010

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

Tim Oliver Berg*
Goethe University Frankfurt and IMFS

May 19, 2010

Abstract

This paper explores the relation between stock prices and the current account for 17 OECD countries in 1980-2007. I use a panel vector autoregression (VAR) to compare the effects of stock price shocks to those originating from monetary policy and exchange rates. While monetary policy shocks have little effects, shocks to stock prices and exchange rates have sizeable effects. A 10% contraction in stock prices improves the current account by 0.3% after two years. Hence I find a channel, in addition to the traditional exchange rate channel, through which external balance for an OECD country with a current account imbalance can be restored.

Keywords: current account fluctuations, stock prices, panel VAR

JEL-Codes: C33, E44, F32

*I am grateful to Stefan Gerlach and Maik Wolters for helpful comments. Contact information: Tim Oliver Berg, Institute for Monetary and Financial Stability (IMFS), Goethe University Frankfurt, Grüneburgplatz 1 (Box H12), 60629 Frankfurt am Main, Germany, Tel.: +49 (0) 69 798 34503, Fax: +49 (0) 69 798 34502, email: tberg@wiwi.uni-frankfurt.de

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, it is even more important to understand the sources of these imbalances and the likely adjustment mechanisms. Particularly, the role of stock prices is of interest and is thus the central issue of this paper. The existing literature on the link between stock 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 and compare the effects of stock price shocks to those originating from monetary policy and exchange rates.

Since the U.S. current account imbalance is so large, many authors focus on the U.S. in their analysis. While some point to low private savings in the U.S. as a main driver of this imbalance (see, e.g., Krugman, 2007), others investigate the role of public savings (see, e.g., Erceg et al., 2005; Corsetti and Müller, 2006, among others). From a simple accounting perspective, budget and current account deficits move in the same direction. Thus, the swing of the U.S. fiscal position from surplus to deficit during the Bush era may have accelerated the deterioration of the U.S. 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. Moreover, Kim and Roubini (2008) find even evidence of a 'twin divergence', i.e., when fiscal accounts worsen, the current account improves and vice versa.

Another camp identifies productivity shocks as a main determinant of the current account (see, e.g., Bussière et al., 2005; Corsetti et al., 2006; Bems et al., 2007, among others). Country-specific productivity shocks raise relative consumption as well as the price of nontradables and deteriorate the trade balance. Corsetti et al. (2006) find evidence that this effect is particularly persistent for the United States. A third strand focuses on the role exchange rates play in restoring external balance for countries with large external deficits (see, e.g., Obstfeld and Rogoff, 1995; Blanchard et al., 2005, among others). A common result of this literature is that a large and steady depreciation of the exchange rate is needed to rebalance the current account (see, e.g., Krugman, 2007).

Despite the vast literature on the sources of current account fluctuations, it is striking that only few authors discuss the contribution of stock 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 U.S. 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 reserves was huge among Asian countries. Since the U.S. financial market is the largest and most liquid in the world, a dominant fraction of these reserves were invested in U.S. dollar denominated assets, particularly in U.S. government bonds. Furthermore, the surge in oil prices created large surpluses among the oil-exporting countries that were in turn reinvested in U.S. bonds and equity. In addition, the lack of well functioning capital markets in the emerging world spurred the demand for U.S. assets. As Bernanke (2005) puts it, a 'saving glut' in Asia and among oil-exporting countries is a potential driver of the U.S. current account deficit.

Consequently, I 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 stock 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 stock 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.

Of course, there is no clear structural interpretation of a stock price shock. Building on the assumption that stock prices are forward-looking and thus reflect people's expectations, a large body of the literature interprets shocks to them as shifts in expectations, and so do I. For example, people expect productivity to rise in the future or the share of a country's output in the world to increase (see Engel and Rogers, 2006). Alternatively, one may also think of stock price shocks in the form of rational bubbles (see Kraay and Ventura, 2005).

Fratzscher et al. (2007) find that shocks to stock prices have large and persistent effects on the U.S. trade balance. Using a Bayesian VAR, they measure the impact of a 10% increase in stock 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 G-7 economies and obtain again evidence of a significant impact of stock price movements on the trade balance. However, the response of the trade balance to stock price shocks varies substantially across countries suggesting that a strong response is probably unique to the United States.

My paper contributes to the existing literature in the following way. Using a panel vector autoregression, I investigate the impact of monetary policy, stock 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 stock price and exchange rate shocks have a significant impact, while monetary policy shocks have little effects. Hence I find a channel, in addition to the traditional exchange rate channel, through which external balance for an OECD country with current account deficits can be restored. An extended period of falling stock prices is likely to reduce real activity through wealth and balance sheet effects as well as the demand for imports and thus improves the current account.

The rest of the paper is organized as follows. Section 2 outlines the Panel VAR model and the identification strategy. An impulse response analysis and a forecast error variance decomposition are presented in Section 3. Moreover, I provide robustness checks. Finally, Section 4 concludes.

2 Methodology

2.1 Panel VAR model

I use a panel VAR of the form:

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

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 includes 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, a nominal short-term interest rate, a nominal long-term interest rate, nominal stock prices, a real effective exchange rate, and a current account to GDP ratio. The estimation period is 1980Q1-2007Q4 and I provide a description of the data in Appendix A. The variables are expressed in logs, except the interest rate variables and the current account to GDP ratio, which are in percent. Since the current account is measured with respect to the 'rest of the world', I find it appropriate to incorporate all other endogenous variables in relative terms. I proceed 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 I attach to country j for country i in period t is:

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

where $imp_{i,j,t}$ is the amount of goods and services (in millions of U.S. 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 follows:

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

Using time-varying rather than fixed weights allows me to control for changing patterns in global trade. I proceed like this 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 stock prices (s_{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. Finally, I obtain relative variables by substracting foreign from domestic variables. Hence, the vector of endogenous variables becomes

$$Y_{it} = \begin{bmatrix} y_{it} - y_{it}^{*} & p_{it} - p_{it}^{*} & r_{it}^{s} - {r^{s}}_{it}^{*} & r_{it}^{l} - {r^{l}}_{it}^{*} & s_{it} - s_{it}^{*} & REER_{it} & ca_{it} \end{bmatrix}'. \tag{4}$$

The construction of foreign variables is comparable to the procedure of Pesaran et al. (2004) or Dees et al. (2007) in a Global VAR context. Moreover, Fratzscher et al. (2007) follow a similar strategy for the U.S. and specify the variables relative to the rest of the world. However, they use weights based on global GDP shares rather than trade weights. Alternatively, I could include domestic and foreign variables separately. Given the number of variables, however, this procedure is computationally hardly feasible.

One purpose of my paper is to evaluate the effects of monetary policy shocks on the current account and therefore I include all relevant channels through which monetary policy impacts on the economy. 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 stock prices and exchange rates. I include nominal stock prices since I expect that movements in them have contributed to the development of global imbalances in the last three decades. Finally, I add the real effective exchange rate to capture the external competitiveness of the country under study.

The vector of common exogenous variables, D_t, includes the U.S. dollar price of oil, p_t^{oil}, 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 helps to overcome this problem since both are correlated with inflation expectations. Second, some of the countries in the panel are net oil exporters (notably Canada, Norway and the UK) and are influenced by movements in the price of oil. Third, I do not control for cross-section dependence in the panel and expect that including an observed common factor reduces ineffiences that 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 Swamy (1970) and Pesaran and Smith (1995), I assume that the B_i and C_i matrices vary across countries according to the following random coefficient model:

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

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

As long as the time series dimension T is sufficiently large to run individual time series regressions, I can estimate the panel VAR 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 VARs 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 (see Nickell, 1981). 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, I prefer the mean group estimator and estimate the coefficient matrices as follows:

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

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}.

Furthermore, I obtain all relevant statistics, such as impulse responses or a forecast error variance decomposition, accordingly, i.e., by averaging the respective numbers over all countries.

2.2 Identification

A common way of analyzing the dynamics of a panel VAR is to calculate impulse responses. I assume that the reduced form errors (u_{it}) are linked to the structural innovations (ϵ_{it}) in the following way:

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

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 me to specify the instantaneous causal ordering of the variables. In what follows, I assume that the variables in the system are ordered as in Y_{it}.

Monetary policy shocks raise the relative short-term interest rate ($r^s_{it} - r^{s*}_{it} > 0$) but 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, I allow the financial market variables (long-term interest rates, stock prices and exchange rates) to respond immediately to changes in short-term interest rates. Similar identification schemes are often used in the analysis of monetary policy transmission in an open economy context (see, e.g., Eichenbaum and Evans, 1995; Grilli and Roubini, 1996, among others).

Stock price shocks are associated with an increase in relative stock prices ($s_{it} - s^*_{it} > 0$). Again, real GDP and consumer prices respond with a lag. Furthermore, it seems likely that monetary policy takes changes in stock prices into account since they potentially influence real GDP and consumer prices. However, I do not expect that monetary policy reacts instantaneously to changes in stock prices but only if they rise or fall for a longer period of time. The same argument applies to the exchange rate. Hence, both variables are ordered after real GDP, consumer prices and the short-term interest rate.

Within the block of financial market variables an appropriate ordering is, however, unclear. But it turns out that the impulse responses are robust to alternative ordering schemes. Therefore, I order the financial market variables as follows: first, long-term interest rates; second, stock

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

	$y - y^*$	$p - p^*$	$r^s - r^{s*}$	$r^l - r^{l*}$	$s - s^*$	$REER$	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
U.S.	-2.22	$-3.75^{\dagger}$	$-3.49^{\dagger}$	-2.65	-1.67	-2.61	-1.90

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

prices; and third, the real effective exchange rate. Furthermore, exchange rate shocks raise the real effective exchange rate ($REER_{it} > 0$).

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

3 The results

3.1 Unit root tests

Before presenting the main results in the next sections, I explore the integrating properties of the variables in the panel VAR. I have to decide whether estimating the model in levels or first differences, which depends on the order of integration of the variables. Table 1 shows the results of augmented Dickey-Fuller (ADF) unit root tests for the endogenous variables in level

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

	$y-y^*$	$p-p^*$	r^s-r^{s*}	r^l-r^{l*}	$s-s^*$	$REER$	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	$-3.41^\dagger$	$-3.43^\dagger$	$-6.04^\dagger$	$-5.91^\dagger$	$-5.52^\dagger$	$-5.22^\dagger$	$-5.86^\dagger$
U.S.	$-4.38^\dagger$	$-3.02^\dagger$	$-3.04^\dagger$	$-4.92^\dagger$	$-4.58^\dagger$	$-3.40^\dagger$	$-3.78^\dagger$

Notes: ADF tests include a constant only. A † denotes significance at the 5 percent level.

specification. The ADF regressions contain a constant and a time trend. I set the lag order for the first differences equal to five. I report similar test results for the endogenous variables in first differences in Table 2. In this case the ADF regressions include a constant only and the lag length is four. The results are insensitive to variations in the lag length.

Overall, 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 percent significance level for any variable. In contrast, the test statistics for the endogenous variables in first differences are, with only a few exceptions, highly significant. Consequently, I conclude that the endogenous variables are $I(1)$. I draw the same conclusion for the oil price. In this case the test results are -0.48 (level) and -5.53 (first difference), respectively. Thus, it would be a valid strategy to estimate the panel VAR in first differences. However, differencing the variables destroys cointegrating relationships 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.

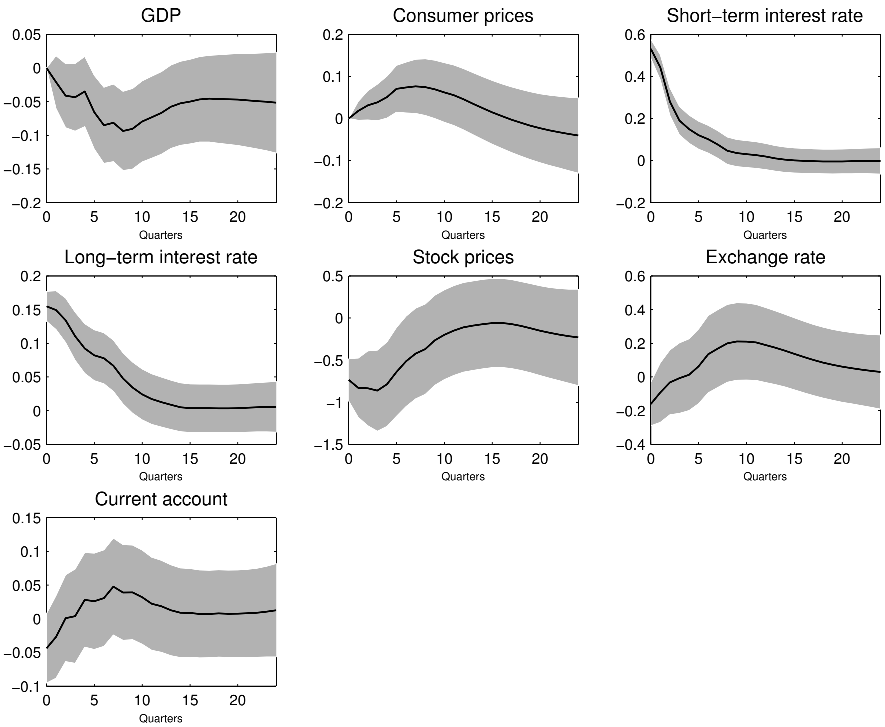

Figure 1: Monetary policy shocks. Notes: I show the responses when the VAR coefficients are fixed at their OLS point estimates, together with a 90 percent confidence interval. Entries are percent.

3.2 Dynamic responses to monetary policy shocks

Figure 1 shows the responses of real GDP, consumer prices, short-term interest rates, long-term interest rates, stock prices, the exchange rate and the current account to one standard error monetary policy shocks, corresponding to an increase in the short-term interest rate of about 50 basis points. I report the responses when the VAR coefficients are fixed at their ordinary least squares (OLS) point estimates, together with a 90 percent confidence interval. I construct error bands using a non-parametric bootstrap that I describe in Appendix B. The figure shows the responses at each horizon between 0 and 28 quarters after the shock.

As you can see, 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. Long-term interest rates fall thereafter and the response is zero after two and a half years. Real GDP contracts significantly following monetary policy shocks 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. Including oil prices does not help to overcome the ’price puzzle’ in my context, presumably the result of the sample period chosen or the fact that I include them as exogenous, not endogenous, variable. Furthermore, stock prices fall sharply in response to a monetary policy tightening, but recover quickly. The trough is reached after four quarters. Furthermore, the response of the exchange rate, which is defined in such a way that an increase means an appreciation, exhibits a puzzle as well. The domestic currency depreciates on impact and it takes nearly one year until the effect turns positive. 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 as well. Overall, these findings are compatible with those of a large body of the monetary VAR literature.

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.

3.3 Dynamic responses to stock price shocks

Figure 2 shows the responses to one standard error shocks that raise relative stock prices by more than 4% initially. As we can see, the rise in stock prices is followed by a significant and long lasting increase in both real GDP and consumer prices, suggesting the presence of wealth and balance sheet effects on consumption and investment, respectively. Moreover, I can distinguish between stock price shocks and technology disturbances. While the former induce a positive correlation between real GDP and consumer prices, the latter are typically associated with a negative correlation. This distinction between stock price and technology shocks is important since I expect that technology innovations are a potential source for movements in stock prices.

Furthermore, in response to the increase in real GDP and rising consumer prices, the monetary policy authority is tightening. Short-term interest rates display a hump-shaped pattern, consistent with the idea that monetary policy follows a Taylor-type feedback rule when setting short-term interest rates. 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 high. Finally, the current account worsens immediately (though not significantly) and reaches a trough after

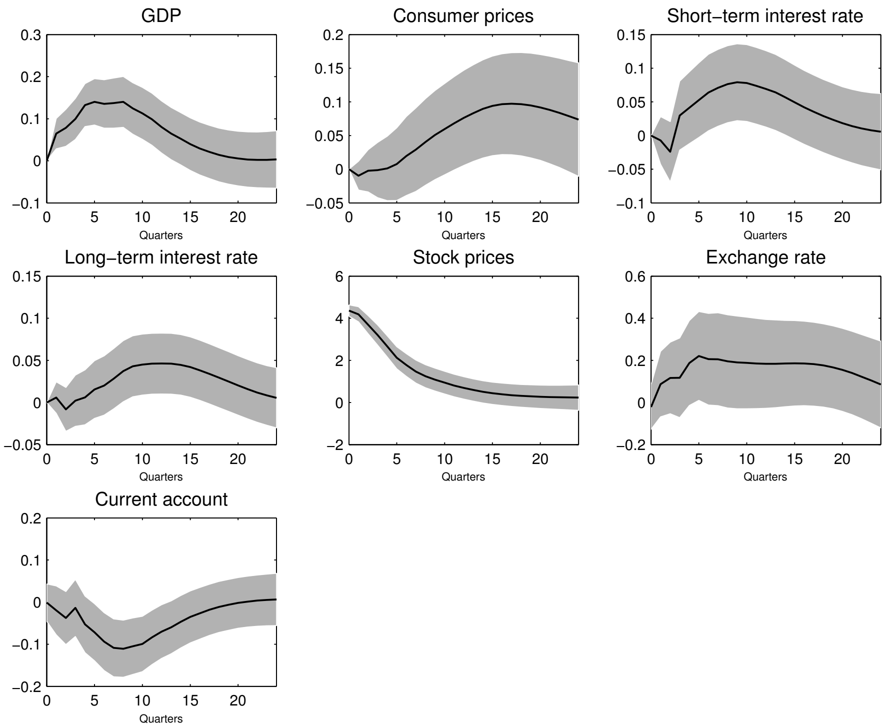

Figure 2: Stock price shocks. Notes: See Figure 1.

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 stock prices on the current account is -0.12% (-0.3%). Hence, the impact of stock price shocks on the current account is not only statistically, but economically, significant, given that stock price movements of 10% (or more) are the norm rather than the exception. Moreover, the results are compatible with these in Fratzscher and Straub (2009) who report responses of the trade balance to stock price shocks (of size 10%) between -1.02% (for Germany) and 0.28% (for the UK) after eight quarters.

3.4 Dynamic responses to exchange rate shocks

I show the responses to one standard error innovations in the exchange rate in Figure 3. 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 exchange rate shocks. Furthermore, the appreciation lowers import prices and as a consequence consumer prices fall. Consumer prices reach a through after around eight quarters. Monetary policy authorities respond to the fall in real GDP and consumer prices by lowering short-term interest

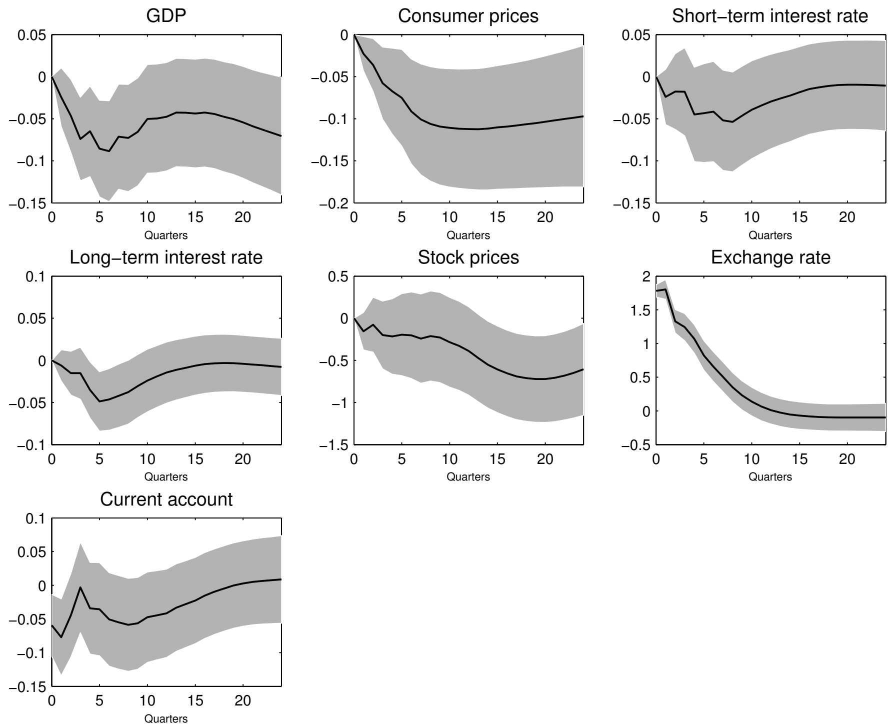

Figure 3: Exchange rate shocks. Notes: See Figure 1.

rates and long-term interest rates match the behavior of short-term interest rates nearly one-to-one. In addition, stock prices fall immediately and are well below their initial level after five years, reflecting the contraction in real GDP.

The current account falls sharply in response to the appreciation. It reaches a trough right in the first quarter after the shock and then improves. However, the response is negative for the next five years. The effect of exchange rate shocks 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 stock price shocks of similar magnitude.

3.5 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. I fix the VAR coefficients at their OLS point estimates and identify monetary policy, stock price and exchange rate shocks in the same recursive way as before. Moreover, I report the contribution of the structural innovations up to 24 quarters following the shock.

Table 3: Forecast error variance decomposition of the current account variable

Horizon	$y-y^*$	$p-p^*$	r^s-r^{s*}	r^l-r^{l*}	$s-s^*$	$REER$	ca
4	6	2	5	3	3	4	77
8	6	3	6	5	7	7	65
12	7	5	7	6	10	9	57
16	9	6	7	6	12	11	50
20	11	7	7	6	14	11	43
24	13	8	7	6	16	12	38

Notes: Entries are percent. I fix the VAR coefficients at their OLS point estimates.

For instance, 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 consumer prices and long-term interest rates contribute less than 8% over all forecast horizons. 13% of the forecast error variance of the current account is accounted for by innovations in real GDP. Givent that I do not attach any structural interpretation to these shocks, the numbers are difficult to interpret.

For any forecast horizon, monetary policy shocks contribute less than 8%. This is compatible with the results of the impulse response analysis. Monetary policy shocks are thus 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 U.S. Federal funds rate as a main source of the variability in the U.S. 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 stock prices and the exchange rate. For long-term forecasts, 16% and 12% of the forecast error variance is accounted for by stock 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 U.S. trade balance. Only a small fraction of the variability can be attributed to exchange rate shocks at long-term forecast horizons. Exchange rate movements appear less important for the U.S. than for other countries. This is not surprising since the U.S. is a large and rather closed econ-

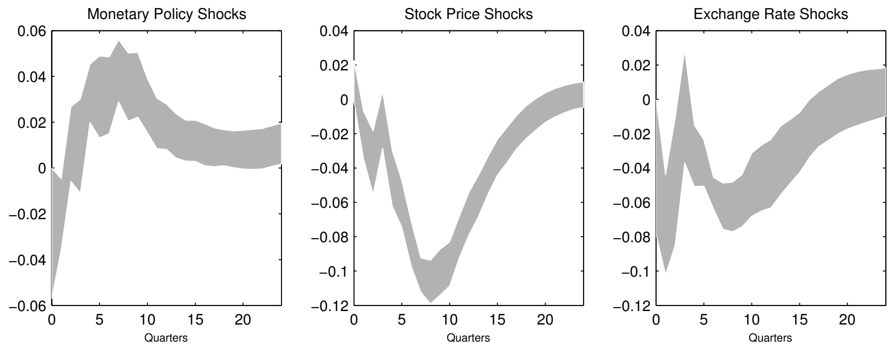

Figure 4: Current account. Distribution of impulse responses. Notes: I show the distribution (shaded area) of responses to monetary policy, stock price and exchange rate shocks based on 5,040 different recursive identification schemes. I fix the VAR coefficients at their OLS point estimates.

omy. However, most countries in my panel are small, open and thus sensitive to exchange rate movements. But the results reconcile with the notion that stock prices explain a considerable part of current account fluctuations. Though the effect is smaller than typically found for the U.S., it is nevertheless notable.

3.6 Robustness

As a robustness check, I evaluate how sensitive the results are to variations in identification. In particular, I estimate the 7-variable panel VAR and construct impulse respones using all 5,040 possible Cholesky orderings. Since I am interested in identification uncertainty, but not sampling uncertainty, I fix the VAR coefficients at their OLS point estimates. As a result of this exercise, I obtain a distribution of impulse responses for the current account variable. 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 monetary policy, stock price and exchange rate shocks, respectively. The top of the shaded area represents the maximum response for each quarter and the lower end corresponds to the minimum. As you can see, the shape of the respones is the same as when using the benchmark identification scheme, while the uncertainty surrounding the point estimates is moderate, suggesting that the results are independent of the restrictions imposed on the covariance matrix. In fact, the covariance matrix is nearly diagonal and thus different identification schemes inevitably lead to similar results.

Furthermore, Figure 5 delivers the joint distribution of the peak and its altitude for the current account. Following monetary policy shocks, the current account improves by 0.05% after 7-9 quarters, confirming that monetary policy shocks have at best a moderate impact on the

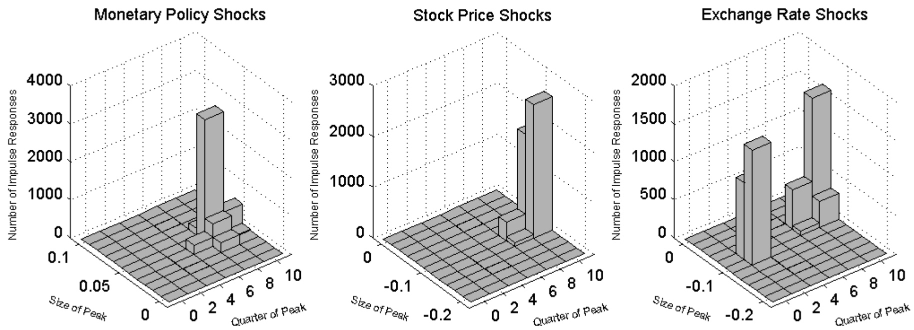

Figure 5: Current account. Distribution of size and location of peak deterioration or improvement. I show the distribution of size (in percent) and location (in quarters) of peak deterioration or improvement, conditional on monetary policy, stock price and exchange rate shocks and based on 5,040 different recursive identification schemes. I fix the VAR coefficients at their OLS point estimates.

current account. In contrast, stock price shocks have sizeable effects. Following stock price shocks, 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. As you can see, the results are different for exchange rate shocks. 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 exchange rate shocks. 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.

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

4 Conclusion

In this paper, I examine the role of shocks to monetary policy, stock 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, stock 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 U.S., but can be attributed to the behavior of the exchange rate which mitigates the effects of monetary policy shocks, in particular, for small open economies.

In contrast, shocks to stock prices and exchange rates have a significant impact on the current account. While a 10% increase in stock prices leads to a deterioration of the current account of 0.3%, an appreciation of the exchange rate of similar magnitude depresses the current account by 0.4%. The effect of stock price shocks on the current account builds up gradually over time and reaches its maximum after around 9-11 quarters. Depending on the identification scheme, exchange rate shocks exert their maximal influence either after 7-11 quarters or within two quarters after the shock. The latter response of the current account to exchange rate shocks is hence inconsistent with the prediction of the Mundell-Fleming-Dornbusch model. In this model, the current account improves on impact following an exchange rate appreciation, before falling over time. Such a 'J-curve' effect is frequently observed in single country VAR models but counterintuitive. Given that the panel set-up allows me to estimate the impulse responses more precisely as compared to single country VARs, it seems plausible that the findings of the existing VAR literature are the result of overparametrized and hence imprecisely estimated models. And finally, stock price and exchange rate shocks explain a notable fraction of the variation of the current account at medium and long-term forecast horizons as compared to others shocks, in particular, monetary policy shocks.

The analysis suggests that stock price and exchange rate shocks are about equally important in explaining current account fluctuations. Thus I find a channel, in addition to the traditional exchange rate channel, through which external balance for an OECD country with a current account imbalance can be restored. I demonstrate the economic relevance of this stock price channel with an example. According to the numbers reported above, a stock market underperformance of 100% improves the current account by about 3%. Such a large underperformance is not unusual if the country under study is in a crisis situation. Moreover, the adjustment needs not to happen immediately but may take several years. Take Japan as an example. While the Japanese stock market lost about half of its value during the 1990s, the U.S. market soared by more than 400% at the same time, providing an explanation for the current account surpluses in Japan and the deficits in the United States. Thus, even for countries with large current account deficits, stock price movements are a potential driver of the adjustment process.

A The data

The data are from the OECD Main Economic Indicators data base and IMF's International Financial Statistics. The estimation period is 1980Q1-2007Q4. I obtain data on real GDP from the IMF with the exception of Canada and Italy where I use data from the OECD. 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. If necessary, I deseasonalize consumer prices and real GDP using the X-11 filter. I take the U.S. dollar price of Brent crude oil from the OECD. Short-term interest 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 I use interbank rates from the OECD. 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 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. Stock prices are from the IMF, except for Switzerland and the United Kingdom where I use data from the OECD. For all countries, I use a broad stock price index. 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, I replace missing observations for 1992Q1-1993Q4 with data from Statistics Norway. Nominal GDP data are from the IMF and in case of New Zealand I again interpolate GDP from annual to quarterly frequency for the early 1980s. Since the current account is denominated in U.S. dollar, I convert it to domestic currency using bilateral U.S. dollar market exchange rates from the OECD, with the exception of Korea where I use data from the IMF.

Finally, the bilateral trade flows that I use 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.

B Error bands

To construct error bands for impulse responses, I use the continuous-path block-bootstrap of Politis (2003). The bootstrap takes the $I(1)$ property of the series into account and is implemented as follows. Suppose a series z_t is non-stationary, $t = 1, 2, ..., T$, and an initial observation z_0 is available. First, I calculate the series of stationary first differences Δz_t, where $\Delta z_t = z_t - z_{t-1}$. Second, I perform a block-bootstrap of the first differences Δz_t by randomly drawing blocks of size four with replacement from $\Delta z_1, \Delta z_2, ..., \Delta z_T$, yielding $\Delta z_1^*, \Delta z_2^*, ..., \Delta z_T^*$. Letting the block size vary between 2 and 12 produces similar error bands. Third, I construct a bootstrap series for z_t by 'integrating' the Δz_t^*, i.e. $z_t^* = z_0 + \sum_{i=1}^{t} \Delta z_i^*$. Fourth, I use the bootstrap series z_t^* to re-estimate the coefficients of the panel VAR. Finally, I calculate the bootstrap impulse responses. I repeat the steps 1,000 times and hence obtain a distribution of impulse responses. I calculate 90 percent confidence intervals as follows

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

where $\hat{\phi}$ are the impulse responses based on the original data and $\hat{\phi}^*$ are the bootstrap counterparts.

References

Barnett, A. and Straub, R. (2008). What drives U.S. current account fluctuations? ECB Working Paper No. 959.

Bems, R., Dedola, L., and Smets, F. (2007). US imbalances: The role of technology and policy. *Journal of International Money and Finance*, 26:523–545.

Bernanke, B. S. (2005). The global saving glut and the U.S. current account deficit. Speech, Sandridge Lecture, Virginia Association of Economics, Richmond, Virginia, March 10.

Blanchard, O., Giavazzi, F., and Sa, F. (2005). The U.S. current account and the dollar. NBER Working Paper No. 11137.

Bussière, M., Fratzscher, M., and Müller, G. J. (2005). Productivity shocks, budget deficits and the current account. ECB Working Paper No. 509.

Chow, G. C. and Lin, A.-L. (1971). Best linear unbiased interpolation, distribution, and extrapolation of time series by related series. *The Review of Economics and Statistics*, 53(4):372–375.

Corsetti, G., Dedola, L., and Leduc, S. (2006). Productivity, external balance and exchange rates: Evidence on the transmission mechanism among G7 countries. NBER Working Paper, No. 12483.

Corsetti, G. and Müller, G. J. (2006). Budget deficits and current accounts: Openness and fiscal persistence. *Economic Policy*, 48:598–638.

Dees, S., di Mauro, F., Pesaran, M. H., and Smith, L. V. (2007). Exploring the international linkages of the Euro area: A Global VAR analysis. *Journal of Applied Econometrics*, 22(1):1–38.

Eichenbaum, M. and Evans, C. L. (1995). Some empirical evidence on the effects of shocks to monetary policy on exchange rates. *The Quarterly Journal of Economics*, 110(4):975–1009.

Engel, C. and Rogers, J. H. (2006). The U.S. current account deficit and the expected share of world output. *Journal of Monetary Economics*, 53:1063–1093.

Erceg, C. J., Guerrieri, L., and Gust, C. (2005). Expansionary fiscal shocks and the US trade deficit. *International Finance*, 8(3):363–397.

Fratzscher, M., Juvenal, L., and Sarno, L. (2007). Asset prices, exchange rates and the current account. ECB Working Paper No. 790.

Fratzscher, M. and Straub, R. (2009). Asset prices and current account fluctuations in G7 economies. ECB Working Paper No. 1014.

Grilli, V. and Roubini, N. (1996). Liquidity models in open economies: Theory and empirical evidence. *European Economic Review*, 40:847–859.

Kim, S. and Roubini, N. (2008). Twin deficit or twin divergence? Fiscal policy, current account, and real exchange rate in the U.S. *Journal of International Economics*, 74:362–383.

Kraay, A. and Ventura, J. (2005). The dot-com bubble, the Bush deficits and the U.S. current account. NBER Working Paper No. 11543.

Krugman, P. (2007). Will there be a dollar crisis? *Economic Policy*, 22(51):435–467.

Nickell, S. (1981). Biases in dynamic models with fixed effects. *Econometrica*, 49(6):1417–1426.

Obstfeld, M. and Rogoff, K. (1995). Exchange rate dynamics redux. *The Journal of Political Economy*, 103(3):624–660.

Pesaran, M. H., Schuermann, T., and Weiner, S. (2004). Modelling regional interdependencies using a global error-correcting macroeconometric model. *Journal of Business and Economics Statistics*, 22(2):129–162.

Pesaran, M. H. and Smith, R. P. (1995). Estimating long-run relationships from dynamic heterogeneous panels. *Journal of Econometrics*, 68:79–113.

Politis, D. N. (2003). The impact of bootstrap methods on time series analysis. *Statistical Science*, 18(2):219–230.

Swamy, P. A. V. B. (1970). Efficient inference in a random coefficient regression model. *Econometrica*, 38(2):311–323.