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In 2007 all ECB publications feature a motif taken from the €20 banknote.

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CONTENTS

Abstract

This paper analyses the role of asset prices in comparison to other factors, in particular exchange rates, as a driver of the US trade balance. It employs a Bayesian structural VAR model that requires imposing only a minimum of economically meaningful sign restrictions. We find that equity market shocks and housing price shocks have been major determinants of the US current account in the past, accounting for up to 32% of the movements of the US trade balance at a horizon of 20 quarters. By contrast, shocks to the real exchange rate have been much less relevant, explaining less than 7% and exerting a more temporary effect on the US trade balance. Our findings suggest that sizeable exchange rate movements may not necessarily be a key element of an adjustment of today's large current account imbalances, and that in particular relative global asset price changes could be a more potent source of adjustment.

Keywords: current account; global imbalances; exchange rates; Bayesian VAR; sign restrictions.

JEL Classification: F32; F40; C30.

Non-technical summary

The debate about the origins and causes of today's large global current account imbalances continues to be highly contoversial. One camp of this debate points at the US as the culprit, and in particular at its low private and public savings, while others argue that it is a "saving glut" in Asia and among oil-exporting countries that has been the key driver of rising current account dispersions. Moreover, a number of scholars and policy makers have given central stage to the need for a large exchange rate depreciation to increase US exports and, hence, reduce its trade and current account deficits. In general, however, there is no clear-cut empirical evidence so far showing that exchange rates have been a major driver of current account positions for advanced economies. By contrast, other scholars have pointed at the relevance of asset prices for the current account determination and adjustment through wealth effects. The underlying logic is that a rise in asset prices (in particular if it is expected to be permanent) increases expected income of households and, therefore, increases their consumption.

Indeed, one striking feature of the global economy over the past 15 years has been the pronounced cycles and booms in asset prices. US equity prices have risen by almost 500% and non-US stock prices by more than 200% between 1990 and 2000, before losing about one third of their value in the subsequent two years. Prices of real estate and housing have also increased substantially over the past decade, while rising unevenly across the globe with a relatively stronger increase in the US, in particular after the stock market collapse in mid-2000. These developments in equity and real estate prices are likely to have impacted significantly on consumption-savings decisions of US households, and hence on the US current account position. Empirically, however, such connection has not been formally investigated, and therefore we know little about the potential for asset prices to induce current account reversals.

The objective of the present paper is to quantify the importance of asset prices versus exchange rates for the current account, and to identify the channels through which they operate from a US-specific perspective. Our empirical methodology is based on a Bayesian structural vector autoregressive (VAR) model, stemming from the work of Uhlig (2005) and Mountford and Uhlig (2005). Our approach requires imposing only a small number of sign restrictions that have an economically meaningful and rather uncontroversial interpretation, while avoiding some of the identification problems present in more traditional structural VAR models.

The empirical findings indicate that equity market shocks and housing price shocks have been major determinants of US current account developments in the past, accounting for up to 32% of the movements of the US trade balance at a horizon of 20 quarters. By contrast, shocks to the real exchange rate have been a less relevant driver, explaining about 7% and exerting a more temporary effect on the US trade balance. The impulse responses show that a 10% rise in US equity prices relative to the rest of the world lowers the US trade balance by around 0.9% of US GDP, while housing price shocks exhibit a slightly larger elasticity.

What do the findings of the paper imply for the future adjustment process of global imbalances? The analysis has been backward-looking and there is obviously no certainty that economic relationships of the past will hold in the future. The results of the paper suggest that while a large US dollar depreciation *could* be a key driver of the adjustment process, *it doesn't necessarily have to be*. Instead, a sizable part of the adjustment could stem from an unwinding of relative asset price developments. This may not necessarily require a drop in US asset prices, but could also entail a rise in asset prices outside the United States, in particular among EMEs. Improving financial market development and financial intermediation may help trigger such a revision, ultimately inducing lower savings and/or higher investment in EMEs, and thus becoming part of the adjustment process.

1 Introduction

The emergence of large global current account imbalances over the past decade has triggered a controversial academic as well as policy debate about their causes and likely adjustment. The controversy stems in part from the two-sided nature of these imbalances, reflected in a large current account deficit in the US - reaching close to 7% of US GDP in 2006 - and correspondingly high surpluses mainly among Asian economies and oil-exporting countries. One camp of this debate points at the US as a driver of these imbalances, and in particular at its low private and public savings (e.g. Krugman 2006). Yet the large US external deficit may not solely reflect policy distortions but is at least partly due to the rise in US productivity (e.g. Corsetti et al. 2006, Bems et al. 2007, Bussiere et al. 2005), expectations of a rising share of the US in world output which may rationalize even a substantial part of the size of today's US current account deficit (Engel and Rogers 2006), and a reduction in income volatility and uncertainty (Fogli and Perri 2006).

Another camp has been focusing on the role of surplus countries and points at the "saving glut" in Asia and oil-exporting countries (e.g. Bernanke 2005). In particular, Caballero et al. (2006) and Ju and Wei (2006) argue that the lack of financial assets and incompleteness of asset markets in emerging market economies (EMEs) is key for understanding the direction of capital flows from poor to rich countries and its composition, the ample liquidity in global capital markets and low interest rates.[1] A third strand of the literature has been concentrating on likely adjustment mechanisms. Some theoretical work argues that required exchange rate changes, in particular a depreciation of the US dollar, to reduce trade imbalances may potentially be large (e.g. Obstfeld and Rogoff 2005, Blanchard et al. 2005), while others point out that such implications are not necessarily borne out by all models and that, under some scenarios, required exchange rate changes may be smaller (Engel and Rogers 2006, Cavallo and Tille 2006).

An important question is the role of asset prices as a driver of global current account positions. One striking feature of the global economy over the past 15 years has been the pronounced cycles and booms in asset prices.[2] A key feature, and one that is central to the analysis of the paper, is that the rise in asset prices over the past 15 years has been much more pronounced in the US than in other advanced economies and many EMEs. These developments in global and relative

[1] Related studies point at the rapidly increasing degree of global financial integration and ensuing valuation effects on gross international asset positions (Gourinchas and Rey 2005, Lane and Milesi-Ferretti 2005), while others underline the role of pre-cautionary motives due to uncertainty and demographics as a rationale for the high saving rates in several EMEs (e.g. Gruber and Kamin 2007, Chinn and Ito 2007).

[2] We distinguish the exchange rate from other asset prices throughout the paper in order to stress that it affects the current account through fundamentally different channels than, for example, equity prices.

asset prices are likely to have impacted significantly global current account positions. Empirically, however, such a connection has not been formally investigated, and we know little about the potential for asset prices to induce current account movements.[3] In principle, asset prices are relevant for current account determination and adjustment through wealth effects. The underlying logic is that a rise in equity prices or housing prices (in particular if it is expected to be permanent) increases expected income of households and thus consumption, while also making it easier for firms to finance investment opportunities, thus inducing a deterioration in a country's trade balance.[4]

The objective of the present paper is to quantify the importance of asset prices versus exchange rates for the current account, and to identify the channels through which they operate from a US-specific perspective. Our empirical methodology is based on a Bayesian structural vector autoregressive (VAR) model, stemming from the work of Uhlig (2005) and Mountford and Uhlig (2005). Our approach requires imposing only a small number of sign restrictions that have an economically meaningful and rather uncontroversial interpretation, while avoiding some of the identification problems present in more traditional structural VAR models. We derive our identifying restrictions from the existing empirical literature rather than from a structural model, partly because structural models of the current account have been shown to be invalid empirically and partly because structural models with a broad set of asset prices are hard to formalize.[5] Importantly, our methodology using sign restrictions on the impulse responses of different types of shocks allows us to distinguish the effects of asset prices from those of other factors. The results are robust to a battery of VAR specifications that include not only variables commonly used in an open-economy setting – such as the real exchange rate, the trade balance, relative consumption, relative prices and relative interest rates (see Eichenbaum and Evans 1995) – but also asset prices.

[3] Seminal studies that analyze the theoretical connection of asset prices and current account developments include Ventura (2001), Caballero et al. (2006) and Kraay and Ventura (2005).

[4] Various segments of the academic literature analyze individual elements that are relevant for understanding the channels of this link. One strand investigates the effects of changes in wealth on consumption, finding marginal propensities to consume of between 0.06 and 0.12 with respect to changes in housing wealth, and somewhat smaller effects with regard to other forms of wealth (Betraut 2002, Case et al. 2005, Palumbo et al. 2002). A different literature has looked at the sensitivity of imports to changes in domestic demand, showing that there is a unit elasticity in the long-run (e.g. Clarida 1994), though recent work emphasizes important differences in these elasticities between changes in investment and changes in consumption (Erceg et al. 2006).

[5] Tests of the intertemporal model of the current account, which postulates that a country's current account position should be equal to the present discounted value of future changes in net output, are frequently based on the procedure developed by Campbell (1987) to test for the restrictions implied by a present value model of asset prices in a VAR framework. However, Sheffrin and Woo (1990) find that the simple intertemporal model of the current account cannot be rejected only for Belgium and Denmark but is invalid for other countries. Bergin and Sheffrin (2000) augment the present-value model to allow for stochastic interest rates and exchange rates, but also the evidence using such models is rather weak, with several papers suggesting a rejection of the model, at least in its canonical form (Nason and Rogers 2006).

The empirical findings indicate that equity market shocks and housing price shocks have been major determinants of US current account developments in the past, accounting for up to 32% of the movements of the US trade balance at a horizon of 20 quarters. By contrast, shocks to the real exchange rate have been a less relevant driver, explaining about 7% and exerting a more temporary effect on the US trade balance. The impulse responses show that a 10% rise in US equity prices relative to the rest of the world lowers the US trade balance by around 0.9% of US GDP, while housing price shocks exhibit a slightly larger elasticity. The effects of both asset price shocks build up gradually over time, with the impulse responses reaching their peaks after around 10 quarters. By contrast, a US real exchange rate shock of 10% affects the US trade balance by up to 0.6% at a horizon of 8 quarters. On impact, a real exchange rate depreciation induces a slight worsening of the trade balance, consistent with a J-curve effect and a standard Mundell-Fleming-Dornbusch model, before improving gradually and becoming positive in its effect on the trade balance. However, real exchange rate shocks exhibit less persistent effects than asset price shocks, and no effect of the exchange rate on the trade balance can be detected beyond a horizon of 16 quarters.

What do these empirical findings imply, and how do they fit into the findings and theories of the existing literature? We stress that we do not interpret the effects of asset price shocks that we find here necessarily as an alternative, but rather as an explanation that is complementary to those of the literature outlined above. For instance, our empirical findings suggest that also productivity shocks and monetary policy shocks have been highly relevant in the past, together accounting for up to 25% of the variations in the US trade balance. However, the importance of asset prices is robust to the inclusion of alternative shocks, underlining that they have indeed been a major determinant of current account movements.

The findings of the present paper support and are consistent with two specific hypotheses in the literature. In particular, they are consistent with the argument by Engel and Rogers (2006) that the large US current account deficit reflects expectations of rising US share of world output, as such expectations should directly be reflected in higher US asset prices relative to the rest of the world. Our findings are also consistent with the argument by Kraay and Ventura (2005) that the sharp increase in asset prices over the past decade constitutes to some extent a rational bubble, which may persist for a considerable period of time. The present paper is consistent with these two hypotheses, although it cannot provide an answer to which of them is more plausible. At the same time, we stress that our findings are not necessarily inconsistent with the conceptual work by Blanchard et al. (2005) and Obstfeld and Rogoff (2005), who show that if only an exchange rate channel were available for an adjustment, the required depreciation of the effective US dollar exchange rate would have to be very large, possibly in excess of 50 to 60 percent. However, our results suggest that a sizeable exchange rate adjustment may not be necessary for current account imbalances to adjust. In fact, movements in asset (equity and housing) prices have been a much more relevant driver of the US trade balance in the past; thus relative asset price changes in the future – either through a drop in US asset prices, a (stronger) rise in foreign asset prices, or both – may be a potent channel for a future adjustment.

The remainder of the paper is organized as follows. Section 2 presents our empirical methodology based on a structural, Bayesian VAR framework in the context of the pure sign restrictions approach. Section 2 also discusses in detail our identification assumptions. Section 3 describes the data. The benchmark results are presented in Section 4, while we report a battery of robustness tests in Section 5. Section 6 summarizes the results, outlines some policy implications and concludes. An Appendix describes an alternative procedure to estimate the Bayesian VAR in the context of the 'penalty function' approach, which we adopt for some robustness checks.

2 The Bayesian VAR Model and Identification

We are interested in analyzing the impact of an exchange rate shock, an equity market shock and a housing price shock on the trade balance of the US in the framework of a VAR model. We follow the approach based on sign restrictions proposed by Uhlig (2005). In addition, since we are interested in accounting for the international transmission mechanism, we consider a VAR model in an open economy framework. We achieve this by incorporating US variables measured with respect to "the rest of the world", proxied by the other G7 countries. This is an appealing feature given that our main variable of interest, the trade balance, is measured with respect to the rest of the world.[6]

2.1 VAR model

Consider the reduced form VAR

$$Y_t = B(L)Y_{t-1} + u_t, \tag{1}$$

where Y_t is an $n \times 1$ vector of time series; $B(L)$ is a matrix polynomial in the lag operator L; u_t is an $n \times 1$ vector of residuals, with variance-covariance matrix $E[u_t u_t{}'] = \Sigma$; and $t = 1, \ldots, T$. An intercept and a time trend may also be allowed for in the VAR model.

[6] Ideally one may want to specify the benchmark model with US variables relative to those of a broader proxy for the rest of the world, e.g. including large emerging markets such as China. Data limitations for such countries do not allow a full specification for all relevant variables. However, we report results which address these limitations to some extent in the robustness section below.

Identification of the VAR in equation (1) is based on imposing enough restrictions to decompose u_t and obtain economically meaningful structural innovations. Let e_t be an $n \times 1$ vector of structural innovations, assumed to be independent, so that $E[e_t e_t'] = I_n$. We need to find a matrix A such that $Ae_t = u_t$. The j-th column of A, a_j, is the impulse vector and depicts the contemporaneous impact of the j-th structural shock of one standard deviation in size on the n endogenous variables in the system. The only restriction on A so far is

$$\Sigma = E[u_t u_t'] = AE[e_t e_t']A' = AA'. \tag{2}$$

We need at least $n \times (n-1)/2$ restrictions on A to achieve identification. One conventional method is to orthogonalize the reduced form disturbances by the Cholesky decomposition. This method assumes a recursive structure on A so that A is restricted to be lower triangular.

2.2 Pure sign restriction approach

Uhlig (2005) and Mountford and Uhlig (2005) achieve identification of the above VAR model imposing sign restrictions on the impulse responses of a set of variables. Uhlig (2005, Proposition A.1) shows that any impulse vector $a \in \mathbb{R}^n$ can be recovered if there is an n-dimensional vector q of unit length such that $a = \widetilde{A}q$, where $\widetilde{A}\widetilde{A}' = \Sigma$, and $\widetilde{A}$ is the lower triangular Cholesky factor of Σ.

Let us start from the case where we wish to identify one structural shock, as in our benchmark VAR results in Section 5.1 below. After estimating the coefficients of the $B(L)$ matrix using ordinary least squares (OLS), the impulse responses of n variables up to S horizons can be calculated for a given structural impulse vector a_j as follows

$$r_s = [I - B(L)]^{-1} a_j, \tag{3}$$

where r_s is the matrix of impulse responses at horizon s. Sign restrictions can be imposed on $m \leqq n$ variables over the horizon $0, ..., S$ so that the impulse vector a_j identifies the shock of interest. The estimation of the impulse responses is obtained by simulation. Given the estimated reduced form VAR, we draw q vectors from a uniform distribution in $\mathbb{R}^n$, divide it by its length, obtain a candidate draw for a_j and calculate its impulse responses, while discarding any q where the sign restrictions are violated.

More precisely, as shown in Uhlig (2005), the estimation and inference is carried out as follows. A prior is formed for the reduced-form VAR. In this case, using as a prior a Normal-Wishart in $(B(L), \Sigma)$ implies that the posterior is the Normal-Wishart for $(B(L), \Sigma)$ times the indicator function on $\widetilde{A}q$.[7] To draw from this posterior we take a joint draw from the posterior of the Normal-Wishart for $(B(L), \Sigma)$ as well as a draw from the unit sphere to obtain candidate q vectors. The draw from the posterior is used to calculate the Cholesky decomposition as in equation (2).[8] Using each q draw, we compute the associated a_j vectors and calculate the impulse responses as described in equation (3). If all of the impulse responses satisfy the sign restrictions, the joint draw on $(B(L),\ \Sigma, a)$ is kept. Each q draw for which the sign restrictions are not satisfied is discarded. This procedure is repeated 500 times and error bands are calculated based on the draws that are kept.

Let us now turn to the more general case where we wish to identify multiple shocks. In our empirical work, we identify up to three structural shocks. In this case, we can characterize an impulse matrix $\left[a^{(1)}, a^{(2)}, a^{(3)}\right]$ of rank 3. This can be accomplished by imposing economically motivated sign restrictions on the impulse responses in addition to restrictions that ensure orthogonality of these structural shocks, since by construction the covariance between the structural shocks $e_t^{(1)}, e_t^{(2)}$ and $e_t^{(3)}$ corresponding to $a^{(1)}, a^{(2)}$ and $a^{(3)}$ is zero.

To see this, start from noting that $\left[a^{(1)}, \ldots, a^{(m)}\right] = \widetilde{A}Q$, with the $m \times n$ matrix $Q = \left[q^{(1)}, \ldots, q^{(m)}\right]$ of orthonormal rows $q^{(j)}$, i.e. $QQ' = I_m$. Mountford and Uhlig (2005, Appendix A) show that the impulse responses for the impulse vector a can be written as a linear combination of the impulse responses to the Cholesky decomposition of Σ in the following way. Define r_{jis} as the impulse response of the j-th variable at horizon s to the i-th column of $\widetilde{A}$, and the n-dimensional column vector $r_{is} = [r_{1i}, \ldots, r_{ni}]$. The n-dimensional impulse response r_{as} at horizon s to the impulse vector $a^{(k)}$, where $k \in \{1, \ldots, m\}$, is given by

$$r_{as} = \Sigma_{i=1}^{n} q_i r_{is} \tag{4}$$

where q_i is the i-th entry of $q = q_k$.

To identify an impulse matrix $\left[a^{(1)}, a^{(2)}, a^{(3)}\right]$, identify $a^{(1)}$, $a^{(2)}$ and $a^{(3)}$ using the relevant sign restrictions $a^{(1)} = \widetilde{A}q^{(1)}$, $a^{(2)} = \widetilde{A}q^{(2)}$ and $a^{(3)} = \widetilde{A}q^{(3)}$, and jointly impose orthogonality conditions in the form $q'q^{(1)} = 0$ and $q'q^{(2)} = 0$. In practice, we take a joint draw from the posterior of the Normal-Wishart for $(B(L), \Sigma)$ and obtain candidate q vectors. If all of the impulse responses satisfy the above restrictions, the joint draw is kept. Each q draw for which the sign restrictions

[7] Essentially the indicator function discriminates the draws where the sign restrictions are satisfied and where they are not. Also, Uhlig (2005) points out that different priors might affect the VAR results. This experiment is, however, beyond the scope of this paper.

[8] Note that this identification scheme does not use the Cholesky decomposition for the purpose of identifying shocks but only as a useful computational tool. Any other factorization would deliver the same results (for a formal proof, see Mountford and Uhlig, 1995).

are not satisfied is discarded. This procedure is repeated 5000 times and error bands are calculated based on the draws that are kept.

Note that Uhlig (2005) and Mountford and Uhlig (2005) also provide another identification method based on a 'penalty function approach'. We briefly describe this method in Appendix A below and implement it for some of our robustness checks.

From a methodological perspective, the pure-sign restriction approach has several advantages. In particular, the results are independent of the chosen decomposition of Σ. This means that a different ordering of the variables and consequent selection of a different Cholesky decomposition does not alter the results. In addition, this method involves simultaneous estimation of the reduced-from VAR and the impulse vector. The idea is that the draws of the VAR parameters from the unrestricted posterior that do not satisfy the sign restrictions receive a zero prior weight.

2.3 Related methods

Other seminal contributions on the sign restriction approach include Faust (1998), Canova and Pina (1999), Canova and de Nicoló (2002). Faust (1998) imposes sign restrictions only on impact. In contrast, Canova and Pina (1999) and Canova and de Nicoló (2002) impose sign restrictions on the cross-correlation function of impulse responses of the VAR variables. Uhlig's (2005) approach differs from Faust (1998) in that restrictions are imposed for several periods. In comparison to Canova and de Nicoló (2002), the identification in Uhlig (2005) is based on impulse responses and not on cross-correlations.

By contrast, conventional VAR identification techniques based on the Cholesky decomposition have often been questioned on various grounds; for example, because they yield counter-intuitive impulse response functions of key endogenous variables which are not easy to rationalize on the basis of conventional economic theory (see Sims and Zha, 2006; Christiano, Eichenbaum and Evans, 1999; Kim and Roubini, 2000), and because the results are often highly sensitive to changes in the ordering of the variables in the VAR (e.g. Sarno and Thornton, 2004). Other approaches to identification include the Blanchard-Quah decomposition (Blanchard and Quah, 1989), which relies on zero long-run restrictions. This procedure identifies permanent and temporary shocks which are usually interpreted as supply and demand shocks, respectively.[9]

[9] Artis and Ehrmann (2006) identify exchange rate and monetary policy shocks in order to analyze the role of exchange rates as a shock absorber. Lee and Chinn (2006) use an identification strategy based on a combination of zero short- and long-run restrictions to bridge the gap between exchange rates and current account studies. They show that permanent shocks have a long-term impact on the real exchange rate but a small impact on the current account. See also the related literature on testing the 'twin deficit hypothesis' in a VAR framework (e.g. Corsetti and Müller, 2006).

It is important to emphasize that our decision to use a Bayesian VAR with sign restrictions does not represent a general criticism of work on identified VARs. Indeed, several authors have proposed, often for reasons similar to the ones which lead us to use a Bayesian VAR in the context of this paper, identification schemes in classical statistical inference without relying on recursive ordering (e.g. see Leeper *et al.*, 1996; Faust and Leeper, 1997; Bernanke and Mihov, 1998). Our chosen methodology is related to and builds on this literature.

2.4 The empirical model

Using quarterly data over the sample period 1974-2005, consider the VAR model:

$$Y_t = \left[\begin{array}{cccccccc} c - c^* & \pi - \pi^* & i - i^* & REER & EQ - EQ^* & H - H^* & HEW & TB \end{array} \right]', \tag{5}$$

where $c - c^*$ is the log of the difference between private consumption of the US versus other G7 countries; $\pi - \pi^*$ is the difference between inflation in the US and inflation in the other G7 countries; $i - i^*$ is the difference between the short-term interest rate in the US and the short-term interest rate in the other G7 countries; $REER$ is the log of the real effective exchange rate (expressed as the foreign price of the domestic currency); $EQ - EQ^*$ is the difference between the market capitalization divided by GDP for the US and the other G7 countries; $H - H^*$ is the difference between the rate of change of real US housing prices and the rate of change of real housing prices in the other G7 countries;[10] HEW denotes home equity withdrawals in the US; and TB is the US trade balance divided by GDP. A detailed description of these variables (and the weights used to construct the relative variables) is provided in the next section.[11] Note that we use consumption, rather than GDP, as we are interested in the transmission channels, especially wealth effects, of shocks to asset prices and exchange rates. Results using GDP instead are, however, qualitatively the same.

Our modelling strategy involves starting from a simple 5-variable VAR ("benchmark" VAR) that includes the exchange rate but initially excludes equity, housing prices and home equity withdrawals. We then examine larger VARs where the two asset price variables, plus home equity withdrawals, are added to the model specification ("augmented" VAR). We use 2 lags for each VAR. The lag

[10] It would be ideal to have a relative measure for housing wealth, similar to that for equity wealth. However, data limitations on this variable did not allow us to construct such an index, as it is only available at an annual basis for the G7 countries, generally for a shorter sample period than the one investigated here. Hence, we employ solely housing prices. However, a comparison of the findings for equity wealth and equity prices, given in the robustness section below, shows that using a wealth measure rather than a price measure does not alter the empirical findings much for the equity market.

[11] In essence, the specification of the domestic real and inflation variables is rather standard in the literature, following Eichenbaum and Evans (1995), Kim and Roubini (2000) and Uhlig (2005).

selection is based on the Akaike (AIC) and Schwartz (SIC) information criteria. Moreover, we set the time horizon for which the restrictions hold after a shock to $S = 2$ quarters for the baseline model.[12]

The crucial issue is the identification of the three shocks of interest, i.e. the exchange rate shock and the shocks to equity markets and housing prices. There are two conceptual challenges. First, the sign restrictions imposed to identify these shocks should be economically meaningful. To this end, we impose sign restrictions that have received substantial support in previous empirical work. Second, the sign restrictions must uniquely identify these three shocks, and not other types of shocks that are included or excluded in our model specification. Table 1 summarizes the short-run sign restrictions imposed. The restrictions imposed to identify a depreciation of the real effective exchange rate are that the real effective exchange rate decreases (i.e. depreciates), the short-term interest rate differential between the US and the other G7 countries increases, the inflation differential between the US and the other G7 countries rises, and relative domestic consumption falls. The rationale for these restrictions stems from the perspective of a monetary policy reaction function: a depreciation should raise import prices and domestic inflation, thus requiring an increase in domestic short-term interest rates and thereby lowering consumption.

Table 1. Identification of shocks through sign restrictions

Shock:	$c - c^*$	$\pi - \pi^*$	$i - i^*$	$REER$	$EQ - EQ^*$	$H - H^*$	HEW	TB
Depreciation	−	+	+	−				
Equity	+		+		+		−	
Housing	+	+	+			+	+	

To identify a positive equity market shock, we impose that relative equity prices, the interest rate differential, and relative consumption all increase. The first of these restrictions is obvious; the second is perhaps less clear-cut and is largely inspired by compelling evidence in the literature. For instance, Rigobon and Sack (2003) – using an identification method based on the underlying heteroskedasticity of the data – show that short-term interest rates rise significantly in response to higher equity prices. Moreover, domestic consumption should rise in response to a positive equity shocks, reflecting the wealth channel discussed above.

Similarly, a positive relative housing price shock is identified by restricting the relative housing

[12] The SIC generally suggests 1 or 2 lags, whereas the AIC usually selects 2 to 4 lags. When a model with more than 2 lags is estimated, the coefficient estimates are similar though the degree of uncertainty surrounding parameters estimates increases. Qualitatively, however, the impulse responses calculated when using 4 lags are identical to the ones generated from the VAR with 2 lags. In addition, we also experimented for different values of K, obtaining very similar results (not reported but available upon request).

price index, the inflation differential, relative consumption and relative interest rates not to decrease. The impulse responses for variables on which sign restrictions are not imposed were left unrestricted. In particular, the response of the trade balance is unrestricted, which is the main focus of our analysis.

How can we ensure that these sign restrictions uniquely identify shocks to the real exchange rate, equity markets and housing prices? A first caveat is that these shocks may at least in part reflect other shocks; for instance an increase in equity prices may be due to a positive productivity shock or any other shock. However, we argue that other shocks differ fundamentally from equity shocks. In the case of a productivity shocks equity prices may also increase, but contrary to equity market shocks, a productivity shock should lower domestic inflation and domestic interest rates, rather than raise them. We will return to a detailed robustness test, including these shocks into the model, in Section 5.

Moreover, it is important to ensure that shocks to the real exchange rate, equity markets and housing prices are distinct from one another. Although our methodology ensures orthogonality of these three shocks in the empirical estimation, a critical issue is to what extent housing price shocks and equity shocks coincide as both exert similar effects, e.g. on domestic interest rates and consumption. We ensure that equity shocks and housing shocks are strictly different by introducing home equity withdrawals (HEW) into the VAR. There is strong evidence that a rise in housing prices increases home equity withdrawals, but equity shocks do not have such an effect (e.g. Case et al. 2005). The economic argument for imposing a negative sign restriction on HEW for equity shocks may be less clear, but it constitutes our preferred specification as it ensures unique identification of all three shocks in the system. Nevertheless, it should be stressed that qualitatively and quantitatively the results presented below are robust to whether or not we impose this restriction. We will return to identification issues in detail in Section 5, where we present a battery of robustness checks and extensions.

3 Data

We use quarterly data over the period 1974-2005. The "rest of the world" series are calculated as weighted averages of the G7 countries without the US, and are identified by an asterisk in our notation.

The output series are taken from the *International Financial Statistics* (IFS) of the International Monetary Fund (IMF), and they are expressed as seasonally adjusted and in local currency. We convert the series from local currency to US dollars using the average market exchange rate for each quarter. Figure 1 (Panel A) shows consumption in the US (c) and the other G7 countries (c^*). Inflation and interest rates are also from the IFS, while the real effective exchange rate is taken from the US Federal Reserve Board Statistics. Inflation series were calculated based on the consumer price index (CPI). The short-term interest rates are 3-month money market rates. The evolution of inflation (π) and interest rates (i) in the US and the other G7 countries (π^* and i^*, respectively, calculated using an arithmetic mean) may be seen in Panels B and C, respectively, of Figure 1. Inflation in the US and the other G7 countries move together for the whole period, with the relative inflation series fluctuating between 1% and 4% until the mid-1980s and since then moving in a narrower range between 0% and 2%. Interest rates reveal a clear downward trend since the beginning of the 1980s. However, interest rates in the US since the 1990s have generally been lower than in the other G7 countries. The highest differential of around 7 basis points occurred in 1993 around the time of the crisis of the Exchange Rate Mechanism (ERM) in Europe. Panel D of Figure 1 shows the evolution of the real effective exchange rate ($REER$). The US dollar experienced a strong real appreciation from the early 1980s until the mid 1980s, then depreciated until 1995 before appreciating again. Note that the REER in our VAR specification is measured in logs.

The US equity measure, EQ, is the stock market capitalization, sourced from *Datastream*, divided by US GDP. EQ^* is the sum of the market capitalization divided by the sum of the GDPs of the other G7 countries. Panel E of Figure 1 shows EQ and EQ^*. The divergence between the two series is substantial, with the market capitalization in the US having increased about fourfold in the period 1974-2000. The strongest increase in the stock market valuation took place in the 1990s. The relative market capitalization roughly tripled between the early 1990s and 2000.

In our robustness analysis, we replace the equity measure based on market capitalization with an index based on equity prices. For all countries, we used a major stock price index, and construct alternative proxies for EQ and EQ^*. In this case, EQ^* is calculated as the weighted average of the price indices of the G7 countries (without the US) weighted by their respective market capitalizations. This equity price measure is shown in Panel F of Figure 1. The divergence between EQ and EQ^* using these series is even more pronounced than the relative equity wealth measure. With the dot-com bubble, the US index soared in the second half of the 1990s and peaked in 2000. US equity prices increased nearly five times more than the average of the other G7 markets. Overall, the stock market boom took place in a period characterized by relatively low inflation, low interest rates and higher output growth in the US than in the other G7 countries.

The US real housing price index is obtained from the Bank of International Settlements (BIS). This index is constructed dividing the nominal housing price index by the personal consumption deflator. The index for the other G7 countries also stems from the BIS and is calculated using time-varying GDP shares at PPP weights. Panel G of Figure 1 plots the US housing price index and the corresponding average housing price index for the other G7 countries. The US housing price index reveals a pronounced cycle from the mid 1970s until the late 1990s between 100 and 120. The strongest increase in housing prices occurred over the past six years, when the index increased by about 40%. The housing price index of the other G7 economies was generally lower than the US index except for a short period from the early to mid 1990s. Since then the difference between the two series has become more pronounced. Note that in our estimation we use the rate of change of the housing price series to calculate $H - H^*$.

Moreover, home equity withdrawal HEW stems from the US Federal Reserve Board Statistics and is defined as the net change in home mortgage debt minus gross residential investment. Since gross residential investment includes additions and alterations to existing homes, this measure captures equity extracted beyond that used to make these improvements.

Finally, the US trade balance, TB series was obtained from IFS (seasonally adjusted) and is expressed as a ratio of GDP. Panel I of Figure 1 shows that since the early 1990s the US has experienced a steady widening in the trade balance deficit, reaching about 6.6% of GDP in 2005 and about 7.0% of GDP in 2006.

4 Empirical Results

4.1 Benchmark VAR and augmented VAR

We now turn to the empirical findings, by presenting the benchmark results from implementing the Bayesian VAR described in Section 2. We begin from a VAR containing a subset of the variables in equation (5). Specifically, we start from a benchmark VAR specification without asset prices and then extend the model gradually to include equity prices and housing prices. This allows us to understand whether and how the inclusion of asset prices into the model changes the empirical findings. The benchmark 5-variable VAR comprises the real exchange rate, trade balance, relative output, relative prices and relative interest rates, i.e. $Y_t = [c - c^* \ \pi - \pi^* \ i - i^* \ REER \ TB]'$.[13]

[13] One may argue that a general to specific approach to econometric modelling may be preferable and hence we should start from the most general, augmented VAR. However, conscious of the fact that no VAR model can possibly avoid possible omitted variable problems, we prefer to start from a simple VAR and move upwards in terms of model size. As shown below, the dynamics depicted by the responses to exchange rate shocks in the simplest VAR without asset prices turn out to be robust when increasing the size of the VAR in subsequent estimations.

Figure 2 shows the effect of a real exchange rate shock using the Bayesian VAR approach with sign restrictions. In all cases, impulse responses are calculated by simulation using the methods described in Sims and Zha (2006). Following Uhlig (2005), each figure shows the median (solid line) as well as the 16th and 84th quantiles (dashed lines). In a normal distribution these quantiles would represent a one standard deviation band. It is common in the VAR literature to report the 2.3% and 97.7% quantiles, which would represent a two standard deviation band in a normal distribution. Given that inference is based on a small number of replications (500) due to computational reasons, resulting in higher uncertainty surrounding the estimates of the quantiles, we report the 16th and 84th quantiles, consistent with the practice by Uhlig (2005) and others in this context.

The impulse responses in Figure 2 suggest that a negative shock (a depreciation of the US dollar) worsens the US trade balance slightly upon impact, consistent with a J-curve effect. After this initial reaction, the trade balance improves from 2 quarters onwards and shows a statistically significant positive reaction between 4 and 20 quarters, while the real exchange rate gradually reverts back to its mean, consistent with some notion of (long-run) PPP. However, the magnitude of the effect is fairly small: a 1.2% real depreciation of the US dollar raises the US trade balance by about 0.08% of US GDP. This implies that even a relatively high real depreciation of the US dollar, for instance by 10%, would improve the US trade balance by a modest 0.6%.

Through what channels does the exchange rate influence the trade balance? The rise of domestic inflation and domestic interest rate relative to their foreign counterparts imply a small rise in the US real interest rate relative to the average rest-of-the-world real interest rate. Moreover, the interaction between interest rates and exchange rates resembles the dynamics implied by the *forward discount bias* routinely recorded in empirical work on exchange rates (Engel, 1996). To see this, note that the negative real exchange rate shock induces an upward movement in interest rates, which according to uncovered interest rate parity (UIP) should imply expectations of a subsequent currency depreciation. However, the US dollar appreciates after the initial depreciation, consistent with the presence of a forward discount bias. One interpretation of the interest rate response is that it is in line with a monetary policy reaction function – in particular as our interest rates are short-term rates, controlled by the central bank. Hence, domestic interest rates are raised by the monetary authority after a currency depreciation due to higher inflationary pressures and are then subsequently lowered as price pressures and output growth subside.[14]

Figure 3 provides the impulse responses using the penalty-function Bayesian VAR approach

[14] We do not pursue further the investigation of welfare effects of different shocks, which are not the focus of this paper and has been addressed in a different context in the literature (e.g. Obstfeld and Rogoff 1995; Corsetti and Pesenti 2001).

described in Appendix A. The results are generally robust to using the penalty function approach, though a few differences emerge. In particular, the parameters are estimated more precisely for some of the variables, consistent with Uhlig (2005). Moreover, the J-curve effect of the exchange rate shock becomes clearer and statistically significant under the penalty function approach. In a few cases the elasticities are somewhat higher, though the differences are small.

Next, we add asset prices to the benchmark VAR specification. Figures 4.A and 4.B show the findings for a 6-variable VAR which includes first only relative equity market changes. Figure 4.A for the real exchange rate shock is virtually identical to the one for the 5-variable VAR. But the figure indicates that the impact of exchange rate shocks on the US trade balance is slightly smaller when including equity prices. Moreover, equity prices show no significant response to real exchange rate shocks. Figure 4.B reveals that an equity shock has a relatively larger, and a more persistent effect on the US trade balance than a real exchange rate shock. A positive relative equity shock of 1.5% (10%) lowers the US trade balance by 0.14% (0.9%) after 10 to 15 quarters. Although this elasticity may not be much higher than the one for real exchange rate shocks, relative equity market changes, in particular throughout the 1990s, have been substantially larger than those for real exchange rates. Moreover, US equity shocks raise relative real interest rates and relative consumption, and both effects are sizeable and persistent. This set of impulse response functions is consistent with the functioning of an equity market shock through wealth effects: a rise in equity prices, in particular if it is expected to be persistent, increases the expected income of households and thus consumption, as well as investment and output due to higher demand, thus overall worsening the trade balance.

In the final model specification, we add relative housing prices and home equity withdrawals to the VAR specification, which is now the "augmented" 8-variable VAR of equation (5). Figures 5.A to 5.C show the impulse response functions for real exchange rate shocks, relative equity shocks and housing price shocks, respectively. The impulse responses for the real exchange rate shock in this 8-variable VAR is similar though slightly smaller than in the VAR without asset prices. The main difference is the faster reversion of the US trade balance to its initial level after a real exchange rate shock (Figure 5.A), again underlining the limited as well as more temporary influence of the real exchange rate on the US trade balance.

The effects of an equity market shock (Figure 5.B) are also similar in the 8-variable VAR as compared to the 6-variable VAR without housing prices. Interestingly, a relative equity market shock does not have a significant effect on the real exchange rate of the US dollar, but it leads to higher US housing prices after 7-8 quarters. This positive effect of equity shocks on housing prices again underlines the functioning of the wealth channel: higher equity prices imply higher wealth and raise both consumption of non-durables and housing prices. Importantly, the negative sign restriction on home equity withdrawals imposed on positive equity shocks does not affect the empirical results as removing this restriction (not shown here for brevity reasons) leaves the impulse response of the variables in the VAR unchanged.

Finally, a positive US housing market shock (Figure 5.C) leads to a significant and persistent deterioration of the US trade balance, with a 1.0% rise in US housing prices lowering the US trade balance by 0.10% after 15 quarters and a gradual reversion thereafter. US housing price development appear to affect the other variables in the VAR in a similar fashion as US equity market shocks. Another important similarity is that a positive housing price shock raises relative equity prices somewhat. This is again indicative of ensuing wealth effects of the housing market increase not only to raise consumption and output but also the demand for equities, thus exerting an upward pressure on equity prices. Moreover, positive housing price shocks raise real interest rates while leaving the real exchange rate largely unchanged.

4.2 Variance decomposition

As the final step of the core empirical analysis, we turn to a variance decomposition and, in particular, to the question of how much of the variation of the US trade balance over the sample is accounted for by asset price shocks as compared to real exchange rate developments. The results, given in Table 2, show a compelling finding: a substantial share of the variations in the US trade balance is explained by asset price shocks. Indeed, close to 32% of the trade balance is explained by equity market shocks and housing price shocks at a horizon of 20 quarters. By contrast, at most 7% of the US trade balance is accounted for by shocks to the US dollar real exchange rate.

Table 2. Variance decomposition for the US trade balance

	Shocks		
Steps	REER	Relative Equity	Relative Housing
4 quarters	2.9	4.1	4.4
8 quarters	3.4	9.1	7.0
12 quarters	4.5	14.9	9.7
16 quarters	5.7	18.5	10.6
20 quarters	6.9	19.8	12.1

In summary, the findings of the benchmark model in this section indicate that asset price changes have been an important driver of developments in the US trade balance over the past 30 years. By contrast, real exchange rate movements have been less important, exerting a relatively moderate as well as more temporary effect on the US trade balance.

5 Robustness and Extensions

Empirical results are often dependent on underling assumptions and variables definitions. A key advantage of the Bayesian VAR using sign restrictions is that it requires only a small number of identification restrictions, which are relatively uncontroversial from an economic theory perspective and can be tested in most instances by removing them individually. While we have discussed the identifying assumptions in detail above, we now turn to discussing various alternative variables definitions and also different VAR specifications.

5.1 Allowing for productivity and monetary policy shocks

A first important issue refers to the distinction between productivity shocks and asset price shocks. There is some evidence that productivity shocks have been an important determinant of current account positions (Bussiere et al. 2005; Corsetti et al. 2006). Besides, productivity increases may be important drivers of US asset prices. However, as pointed out by Kraay and Ventura (2005), the large asset price boom in the US in the 1990s and the decline in the early 2000s may hardly be attributed to productivity.

To shed some light on the difference between these two shocks, we incorporate a relative productivity measure in our VAR model and estimate the effects of a shock to this variable.[15] The short-run sign-restrictions imposed for $S = 2$ quarters to identify a positive productivity shock are that relative productivity and relative consumption increase and that the inflation differential and relative interest rate decrease. The first two restrictions should be obvious. The third restriction is motivated by the "mainstream" model of inflation dynamics developed in the 1970s (Gordon 1977) and also present in Staiger et al. (1997). The idea is to treat a productivity shock as a supply-side shock, so that increases in productivity should lower inflation. Dedola and Neri (2005) provide a thorough analysis of identifying productivity and technology shocks in a structural VAR. The main point to note is that the identification of productivity shocks separates them from equity shocks, as a positive equity shock is identified as raising interest rates and consumption.

Figure 6.A shows the effects of a positive productivity shock in a 7-variable VAR which is the augmented VAR discussed in Section 4.2 with the addition of the productivity variable, but without housing and home equity withdrawals in order to keep the model tractable. The main feature of the

[15] We use output per hour in the manufacturing sector as a proxy for productivity for each country. The data are from the BIS. The relative productivity of the other G7 countries is calculated using time-varying GDP shares at PPP weights (Canada was excluded due to the unavailability of a similar productivity measure). The relative productivity measure is the log of the difference between productivity in the US and productivity in the other G7 countries.

productivity shock is that it tends to be rather persistent. A 1.5% increase in relative productivity worsens the trade balance by 0.10% of US GDP with the effects being significant after 5 quarters. Productivity shocks also affect other variables in the system, for example inducing a persistent increase in consumption. Most interesting for our purpose is that equity markets reacts positively to a productivity shock, in line with our argument that equity increases in part reflect productivity developments. By contrast, the impact on the real exchange rate is small.

Turning to the analysis of the effects of an exchange rate depreciation shock and an equity shock in this VAR which conditions on productivity, the impulse response functions reported in Figures 6.B and 6.C exhibit very similar patterns to those obtained from the 8-variable VAR without productivity described in Section 4.2. Overall, these results indicate that the benchmark results are robust to the inclusion of productivity in the model.

Finally, we analyze the impact of monetary policy shocks. Our aim for extending the analysis to this shock is not only to ensure the robustness of the effects of asset prices as a driver of the US trade balance, but also because it is frequently mentioned in the debate on global imbalances as a relevant factor (e.g. Bems et al. 2007). In particular, it has been argued that the high saving rates in EMEs, in particular in countries such as China, have put downward pressure on US inflation and US interest rates along the whole yield curve (Warnock and Warnock 2006). To allow for this channel in our VAR framework, we introduce monetary policy shocks. The sign restrictions imposed are the same as those in the literature discussed above, such that a tightening shock of short-term interest rates lowers inflation and asset prices, and raises the US real effective exchange rate.

Table 3. Extended variance decomposition for the US trade balance.

	Shocks			
Steps	REER	Equity	Productivity	Mon.Pol.
4 quarters	1.7	7.3	4.6	3.2
8 quarters	2.4	10.7	8.7	6.3
12 quarters	3.6	16.4	9.2	7.6
16 quarters	3.9	20.6	11.1	8.2
20 quarters	4.1	23.5	14.9	10.4

Table 3 shows the variance decomposition for our extended VAR specification including also productivity shocks and monetary policy shocks. Two key results stand out. First, the role of asset prices as a driver of the US trade balance is confirmed. In fact, the share of the US trade balance explained by equity shocks becomes even slightly higher (23.5% after 20 quarters). Second, we also find support for the findings of the literature in that both productivity and monetary policy have been exerting substantial effects, explaining up to 15% and 10% after 20 quarters, respectively, of the US trade balance. Thus, overall, asset prices are confirmed as a main driver of the US trade balance while exchange rate movements appear to have been much less important.

5.2 Alternative definitions and identification

We estimate the Bayesian VAR using an alternative definition of several variables in our benchmark model. First, we replace the trade balance by the current account. In recent years, the difference between these two variables has been relatively small as the US income account was close to balance. However, the difference was much more sizeable in previous years, primarily due to the large positive net income stemming from higher returns on US assets compared to US liabilities. The problem with including income into our trade measure is that it captures very different elements (including changes in returns) from trade in goods and services; thus, our preferred measure is the narrow trade balance.

Figure 7 gives the impulse response of the US current account to the three types of shocks of interest. Overall, the baseline findings from the augmented VAR prove robust to using the current account instead of the trade balance. The only meaningful difference is that the magnitude of the current account response is slightly larger for all three shocks, confirming the effect of the exchange rate on income via returns and valuation changes (Gourinchas and Rey, 2005).

Second, we test for the sensitivity of the results by using nominal effective exchange rate shocks instead of those of the real exchange rate. As shown for the benchmark VAR above, relative prices react to real exchange rate shocks, which leaves open the question of how much of the real exchange rate shocks reflect nominal exchange rate changes and how much reflects relative price adjustments. Figure 8 shows the impulse responses of the US trade balance to nominal effective exchange rate shocks for the three benchmark VAR models estimated in Section 4. Again the results are not sensitive to using nominal or real exchange rates, with the elasticities only being marginally different.

Third, we also experimented with alternative measures of equity shocks. Our benchmark model defines equity market changes as changes in relative stock market capitalization. Changes in the stock market capitalization can in principle have two sources: changes in prices and through the listing or delisting of firms. Market capitalization is our preferred measure because we are primarily interested in the effects of an equity shock on the trade balance via wealth effects, and (financial) wealth is more closely proxied by market value than prices. However, it can be argued that there are many other forms of (less liquid) wealth that are not captured in our model, but which may exert a significant influence, in particular on consumption and thus on the trade balance. As we have no appropriate alternative measures of wealth internationally, we test for the robustness of our findings by using changes in equity prices as the measure of relative equity market shocks. The impulse responses obtained using this alternative proxy suggest that the results – not reported for space reasons – are robust. As expected, however, the magnitude of the response of the US trade balance is slightly smaller for this alternative definition of equity market shocks.

Finally, as an alternative identification we estimated the full 8-variable VAR using a recursive VAR where identification is achieved using the Cholesky decomposition under the ordering given in equation (5). In this case, some of the impulse responses in the model display counter-intuitive signs, but the impulse responses to and the relative importance of exchange rate, equity and housing price shocks implied by the model are consistent with our core results from the Bayesian VAR. Specifically, both the impulse response analysis and the variance decomposition indicate that exchange rate shocks are significantly less powerful in driving the variation in the US trade balance than shocks to equity prices or house prices. Again, results are not reported to conserve space but available upon request.

6 Conclusions

The debate on the causes of global current account imbalances is still wide open. This paper has focused on one specific question: How important are asset prices and exchange rates as drivers of the US trade balance? There has been important theoretical work stressing the relevance of both the asset price channel through wealth effects and of relative price changes implicit in exchange rate movements, but little systematic empirical work has been carried out to quantify the role of asset prices.

To address this question, the paper has employed a Bayesian VAR model, based on Uhlig (2005), which requires imposing only a minimum of sign restrictions that have meaningful and relatively uncontroversial structural interpretations. Our main finding is that asset prices are a substantially more important driver of the US trade balance than the exchange rate. In fact, 32% of the movements of the US trade balance after 20 quarters can be accounted for by asset price shocks, and only about 7% by the US dollar real exchange rate. These results are robust to various extensions and alternative specifications. For instance, while also US productivity shocks and monetary policy shocks are found to exert a significant effect on the US current account, asset prices remain the single most important driver of the US current account.

An important question is what structural interpretation to give to asset price shocks. We have stressed and have shown empirically that the identified asset price shocks in our model are clearly distinct from other factors that influence asset prices, such as shocks to productivity and monetary policy. Our findings support and are consistent with the argument by Engel and Rogers (2006) that the large US current account deficit may reflect expectations of a rising US share of world output, which should directly be reflected in US asset prices relative to those of the rest of the world. Moreover, our findings are consistent with Kraay and Ventura's (2005) argument that the increase in asset prices over the past decade partly constitutes a rational bubble, which may persist for a considerable period of time.

We have stressed several caveats and open issues. Our analysis focuses on the US role in global imbalances, although we define the variables in our VAR in relative terms, i.e. as US developments in comparison to those in the rest of the G7. Nevertheless, there are other factors that influence global current account positions and which we have not modelled, such as financial market imperfections in EMEs and fiscal policies, among others. The paper has concentrated more narrowly on the role of asset prices and exchange rates as drivers of trade imbalances. Many open questions remain, such as to what extent the findings also apply to smaller and more open economies in other parts of the world. Nevertheless, we argue that it is an important question to assess empirically the impact of asset price shocks in order to better understand the origins of today's global imbalances. This has been the primary objective and intended contribution of the paper.

From a policy perspective, a question that arises is what the findings of the paper imply for the future adjustment process of global imbalances. Our analysis has been backward-looking and there is obviously no certainty that economic relationships of the past will hold in the future. The results of the paper suggest, however, that while a large US dollar depreciation *could* be a key driver of the adjustment process, *it doesn't necessarily have to be.* Instead, a sizable part of the adjustment could stem from an unwinding of relative asset price developments, either via a moderation in US asset prices or a rise in asset prices outside the US.

A Appendix: The Penalty Function Approach

This appendix describes how to identify a shock of interest in the benchmark VAR by minimizing a penalty function which punishes violations from the sign restrictions more strongly than what it rewards to responses that satisfy the sign restrictions.

Using the notation in Section 3, assume that $B(L)$ and Σ are drawn from a Normal-Wishart prior (Assumption B.2 in Uhlig, 2005). Let the penalty function be

$$f(x) = \begin{cases} x & \text{if } x \leqslant 0 \\ 100 \times x & \text{if } x \geqslant 0 \end{cases} \tag{A.1}$$

This function rewards negative responses in a linear proportion and penalizes positive responses in linear proportion, though 100 times bigger. Let $r_{j,a}(k)$ be the impulse response of variable j and σ_j be the standard deviation of the first difference of the variable j. Let $\iota_j = 1$ if variable j should respond with a negative sign up to horizon S to satisfy the sign restriction that identifies the shock of interest; and let $\iota_j = -1$ if the *j-th* variable should respond with a positive sign.[16] Let the total penalty $\Psi(a)$ be

$$\Psi(a) = \sum_{j \in \{\text{"vbles with restr"}\}} \sum_{s=o}^{S} f\left(\iota_j \frac{r_{j,a}(k)}{\sigma_j}\right). \tag{A.2}$$

The impulse vector is obtained through the minimization of the total penalty $\Psi(a)$ for all the variables for which restrictions are imposed up to horizon S. Given that the true VAR is unknown, we should find an impulse vector from each draw of the posterior. This requires the minimization of the penalty function. To make inference from the posterior, we take $1,000$ draws from it using a Monte Carlo approach. This allows us to calculate the impulse responses and the error bands (Uhlig, 2005; Mountford and Uhlig, 2005).

In the case of a 5-variable VAR, to carry out the minimization of the $\Psi(a)$ function for each draw from the posterior we parameterized the space of vectors $(q_j)_{j=1}^{5} \in \mathbb{R}^4$ of unit length such

[16] In the case where a variable should respond positively according to the sign restriction imposed for the identification of a shock, the indicator function flips its sign in order to maintain the system of penalties and rewards of the function in (A.1). This makes sure that the function rewards a positive response in a linear proportion and punishes a negative response more severely (in a linear proportion times 100).

that

$$q = \begin{bmatrix} \cos(\gamma_1)\cos(\gamma_2)\cos(\gamma_3) \\ \cos(\gamma_1)\cos(\gamma_2)\sin(\gamma_3) \\ \cos(\gamma_1)\sin(\gamma_2) \\ \sin(\gamma_1)\cos(\gamma_4) \\ \sin(\gamma_1)\sin(\gamma_4) \end{bmatrix}. \tag{A.3}$$

Similar procedures can be adopted to implement the penalty function approach for larger VARs with more than one structural shock, where the only further restrictions needed is that any two structural shocks are orthogonal to each other when minimizing the penalty function. The impulse responses computed using the penalty function approach generally have smaller standard errors than the ones that are found with the pure-sign restriction approach. This happens because the penalty function approach exactly identifies a unique impulse vector a which characterizes the shock of interest. In contrast, the pure-sign approach finds a range of impulse vectors that are consistent with the sign restrictions imposed. See Uhlig (2005) and Mountford and Uhlig (2005) for further details.

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Figure 1. Data

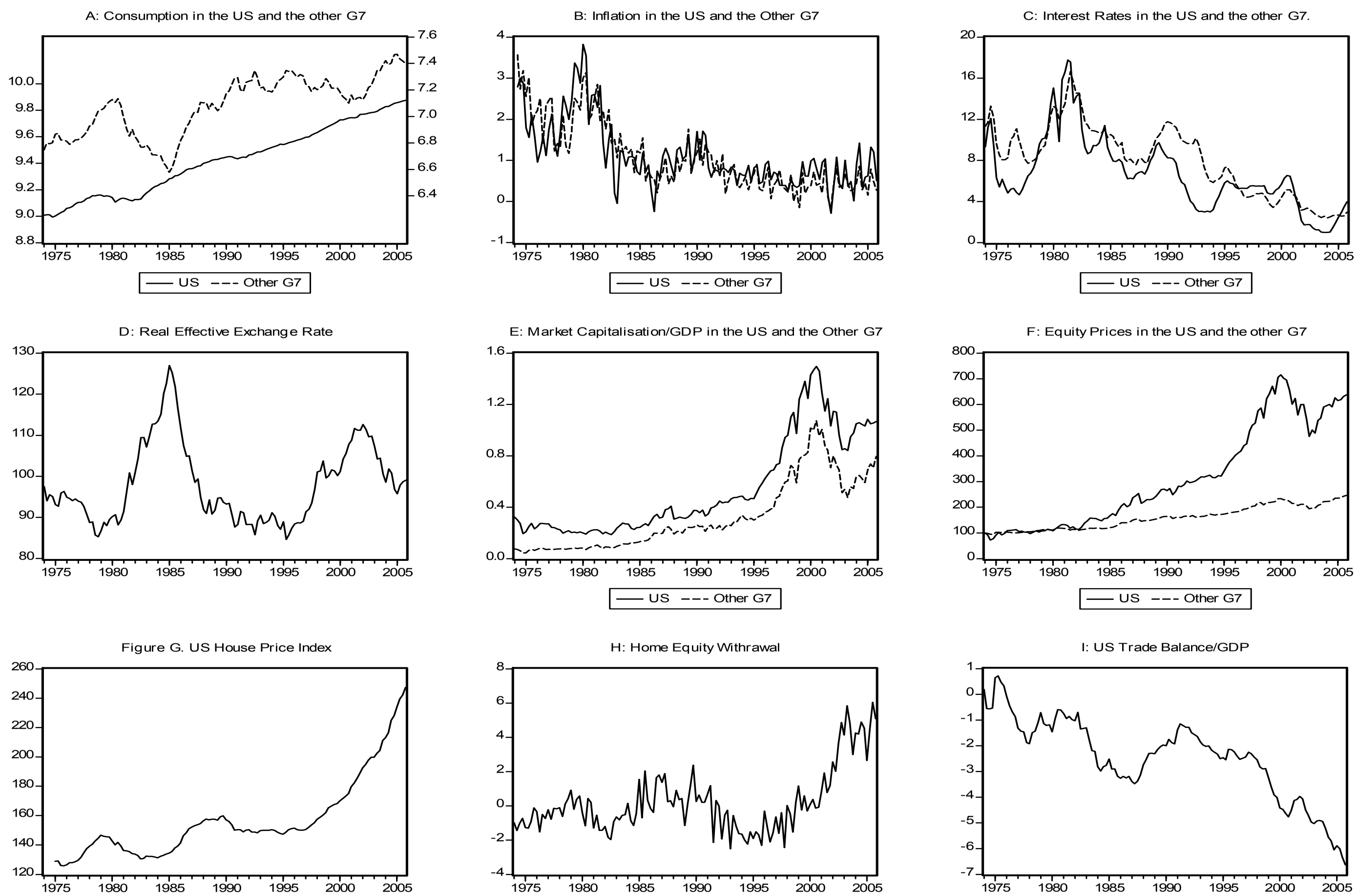
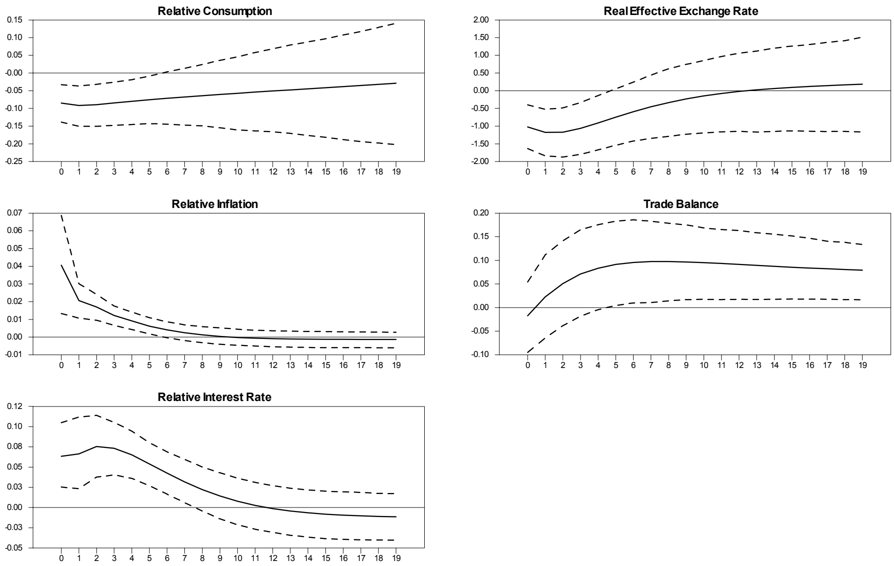
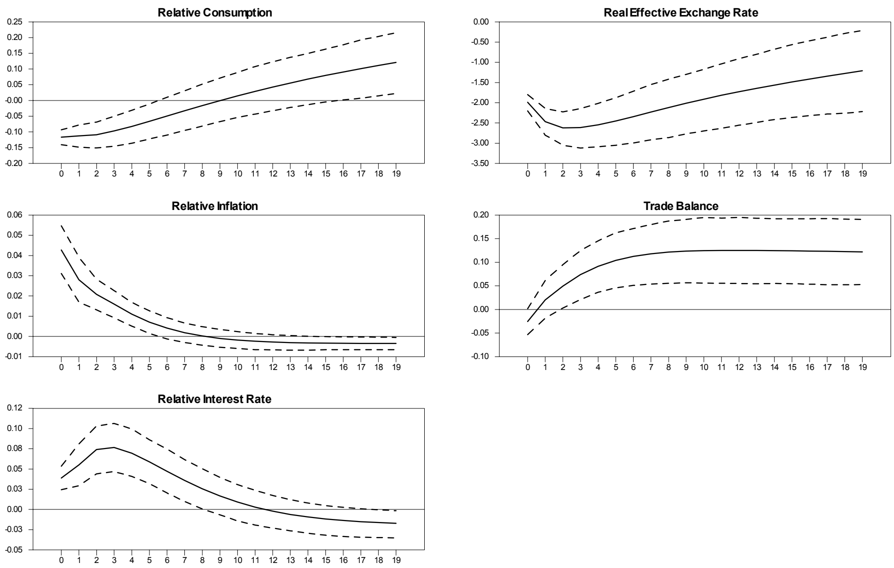


Figure 2. Impulse Responses to Exchange Rate Shock (5-variable VAR)
[Pure-Sign Approach]



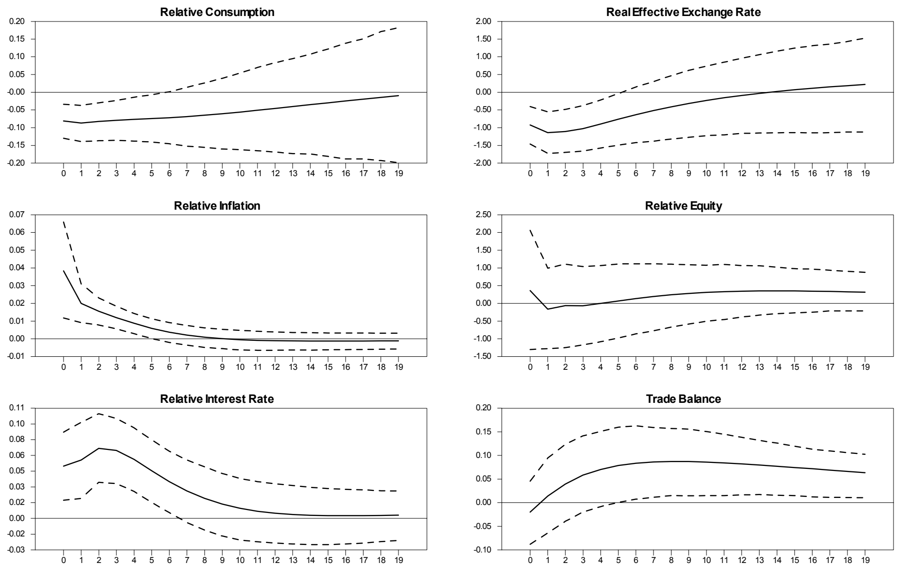
Notes: The figure shows the Impulse Responses to a depreciation shock using the pure-sign restriction approach. The response of the real effective exchange rate and relative consumption were restricted not to be positive and the responses of the relative inflation and the relative interest rate were restricted not to be negative for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

Figure 3. Impulse Responses to Exchange Rate Shock (5-variable VAR)
[Penalty Function Approach]



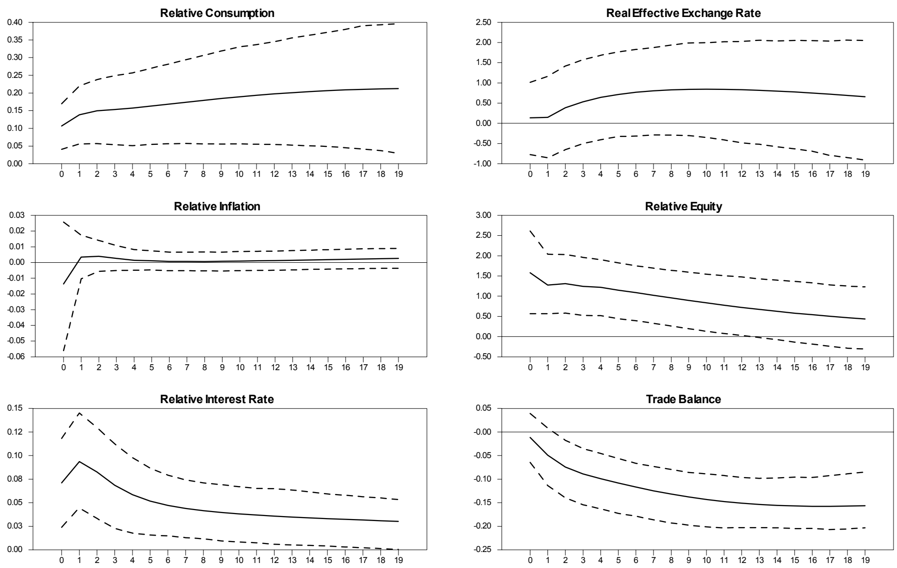
Notes: The figure shows the Impulse Responses to a depreciation shock using the penalty function approach. The responses of the real effective exchange rate, relative consumption, the negative of the relative inflation and the negative of the relative interest rate were penalised for positive values and rewarded for negative values for two quarters according to the penalty function described in the appendix. The shocks are identified by minimising the penalty function. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

Figure 4.A. Impulse Responses to Exchange Rate Shock (6-variable VAR)
[Pure-Sign Approach]



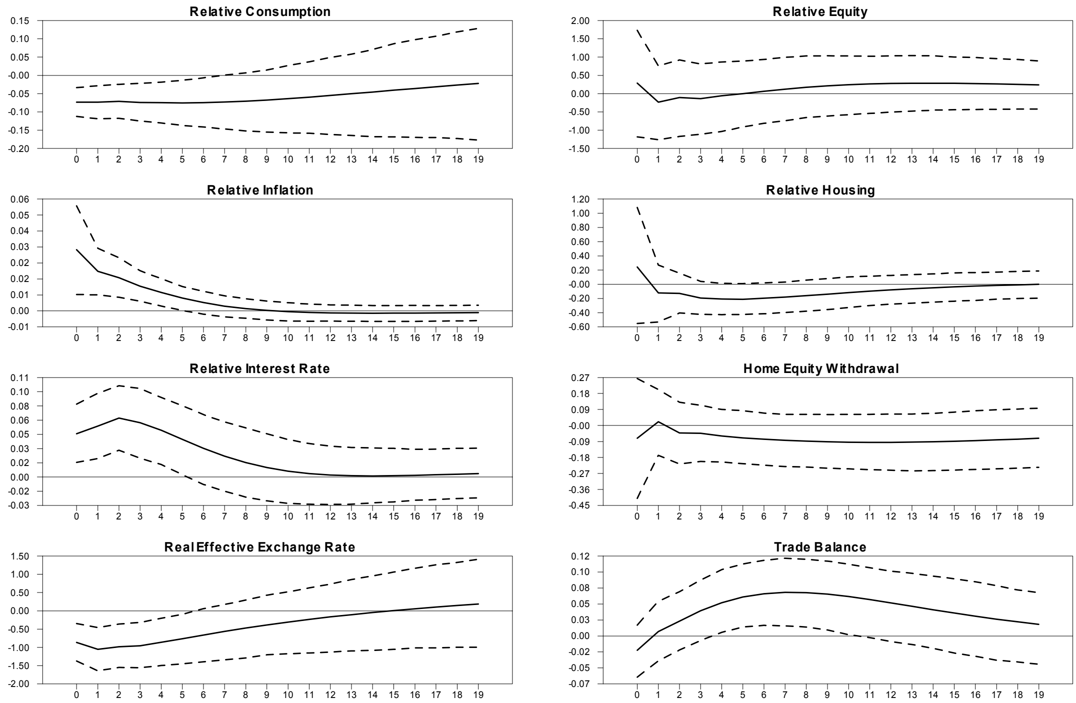
Notes: The figure shows the Impulse Responses to a depreciation shock using the pure-sign restriction approach. The responses of the real effective exchange rate and relative consumption were restricted not to be positive and the responses of the relative inflation and the relative interest rate were restricted not to be negative for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

Figure 4.B. Impulse Responses to Equity Shock (6-variable VAR)
[Pure-Sign Approach]



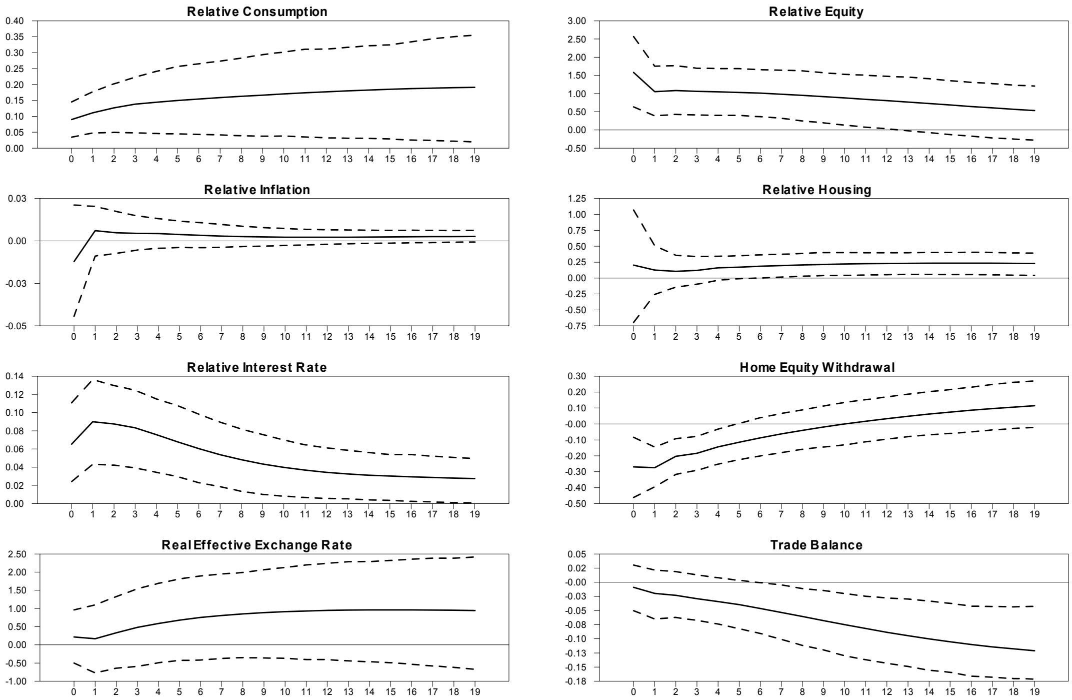
Notes: The figure shows the Impulse Responses to an equity shock using the pure-sign restriction approach. The response of the relative equity, relative consumption and the relative interest rate were restricted not to be negative for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

Figure 5.A. Impulse Responses to Exchange Rate Shock (8-variable VAR)
[Pure-Sign Approach]



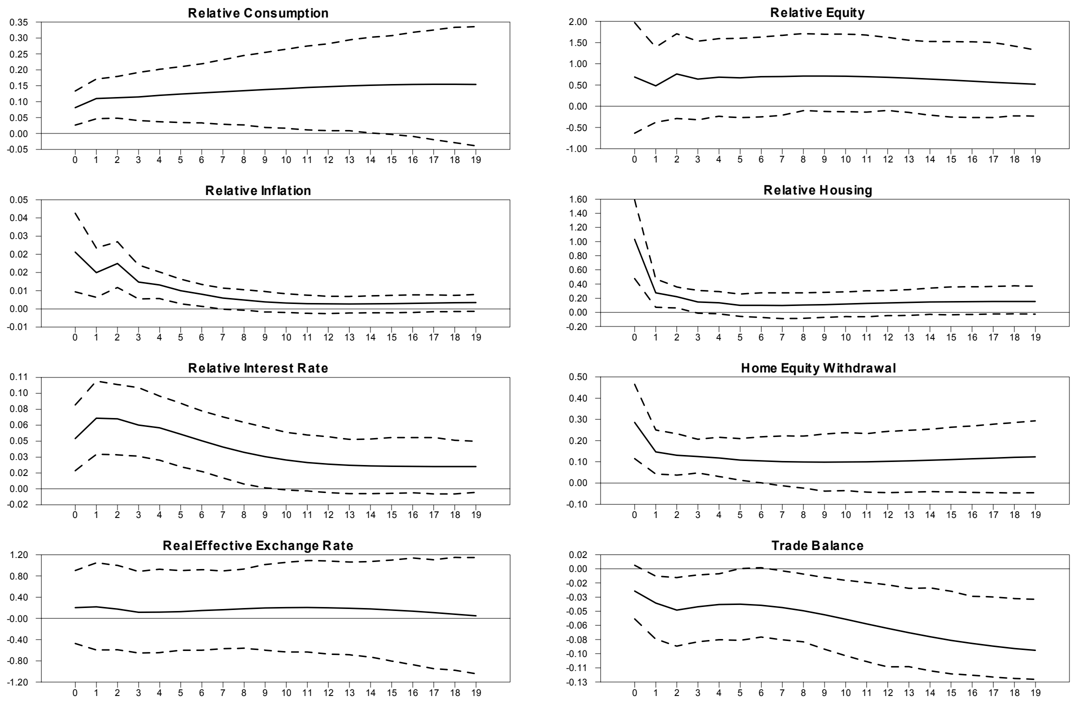
Notes: The figure shows the Impulse Responses to a depreciation shock using the pure-sign restriction approach. The responses of the real effective exchange rate and relative consumption were restricted not to be positive and the responses of the relative inflation and the relative interest rate were restricted not to be negative for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

Figure 5.B. Impulse Responses to Equity Shock (8-variable VAR)
[Pure-Sign Approach]



Notes: The figure shows the Impulse Responses to an equity shock using the pure-sign restriction approach. The response of the relative equity and the relative interest rate were restricted not to be negative for two quarters and the response of home equity withdrawal was restricted not to be positive for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

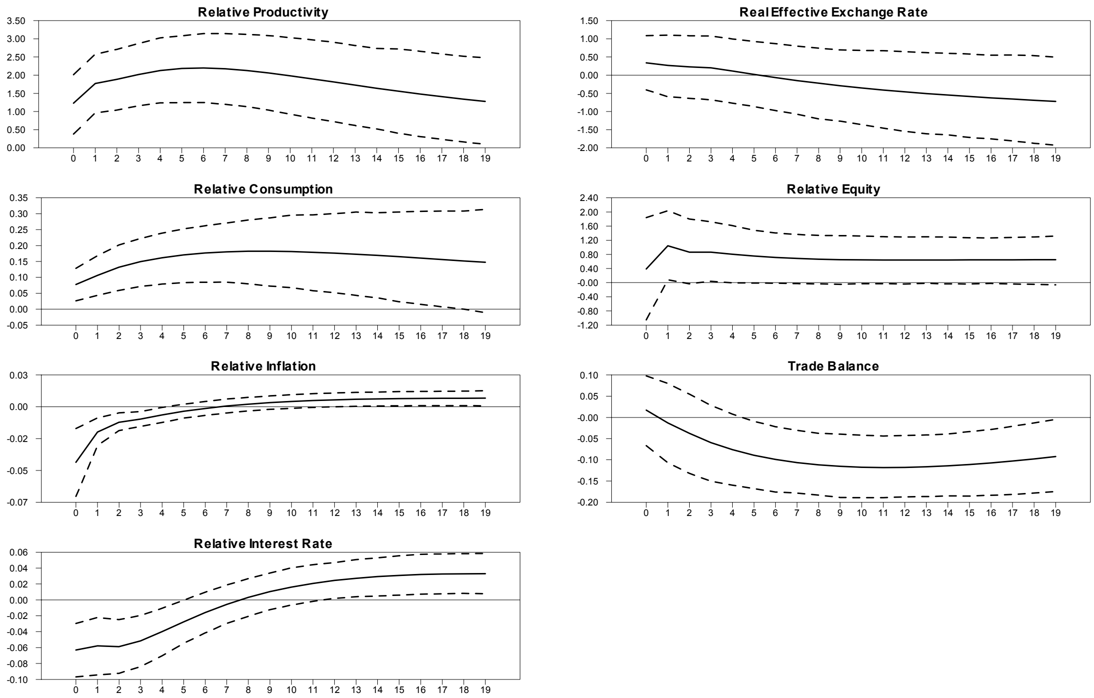
Figure 5.C. Impulse Responses to Housing Shock (8-variable VAR)
[Pure-Sign Approach]



Notes: The figure shows the Impulse Responses to a housing shock using the pure-sign restriction approach. The responses of relative housing, relative consumption, relative inflation, relative interest rate and home equity withdrawal were restricted not to be negative for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

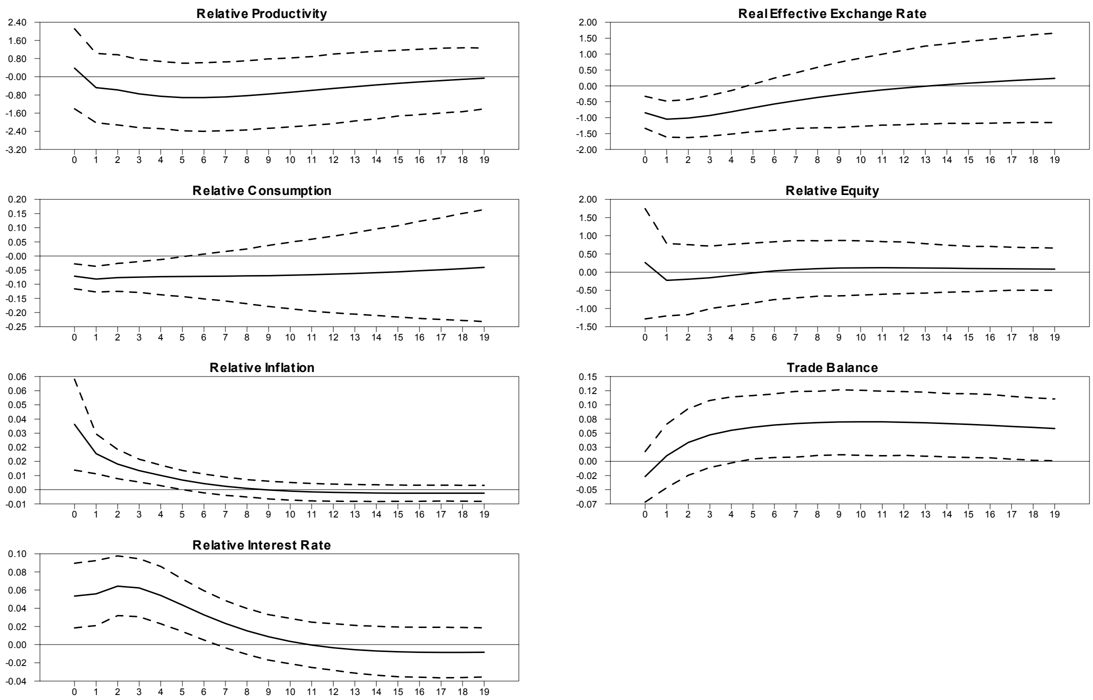
Figure 6.A. Impulse Responses to Productivity Shock (7-variable VAR)

[Pure-Sign Approach]



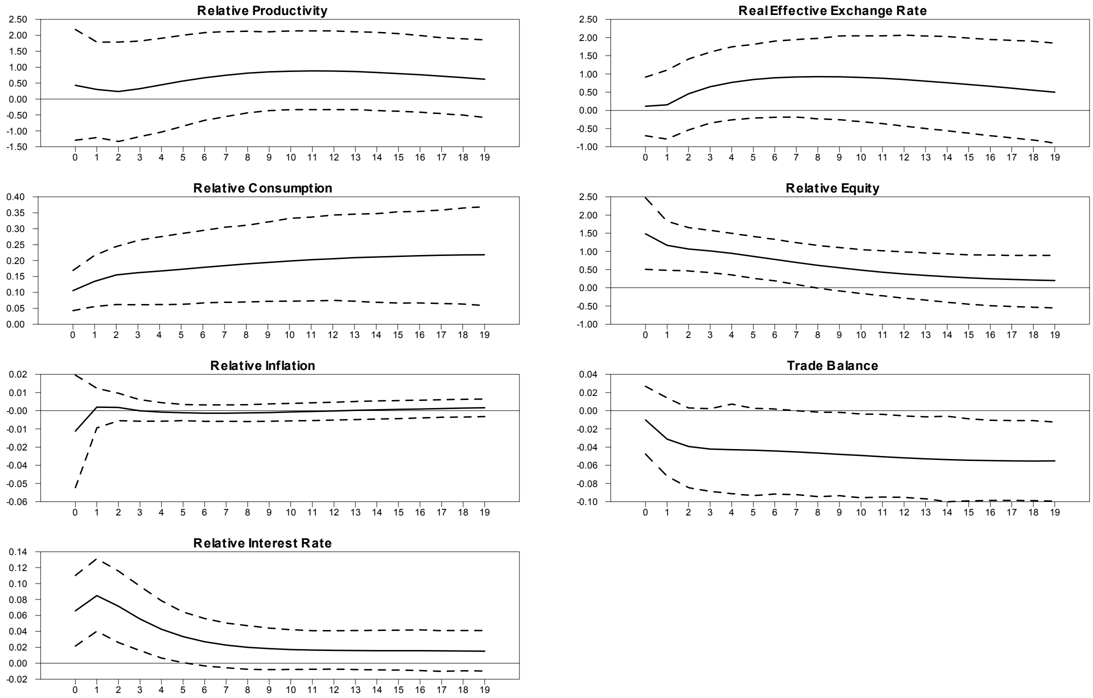
Notes: The figure shows the Impulse Responses to a productivity shock using the pure-sign restriction approach. The response of relative productivity and relative consumption were restricted not to be negative and the responses of the relative inflation and the relative interest rate were restricted not to be positive for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

Figure 6.B. Impulse Responses to Exchange Rate Shock (7-variable VAR)
[Pure-Sign Approach]



Notes: The figure shows the Impulse Responses to a depreciation shock using the pure-sign restriction approach. The responses of the real effective exchange rate and relative consumption were restricted not to be positive and the responses of the relative inflation and the relative interest rate were restricted not to be negative for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

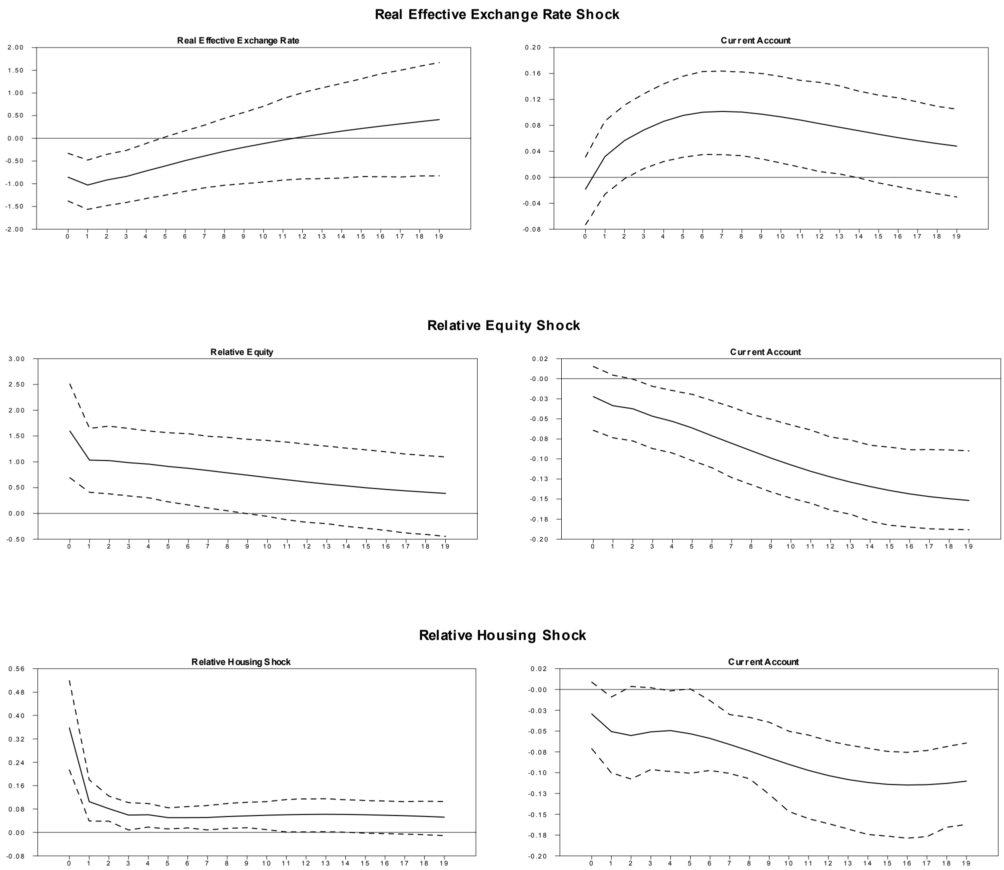
Figure 6.C. Impulse Responses to Equity Shock (8-variable VAR)
[Pure-Sign Approach]



Notes: The figure shows the Impulse Responses to an equity shock using the pure-sign restriction approach. The response of the relative equity, relative consumption and the relative interest rate were restricted not to be negative for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

Figure 7. Impulse Responses of the Current Account to Exchange Rate, Equity and Housing Shocks (8-variable VAR)

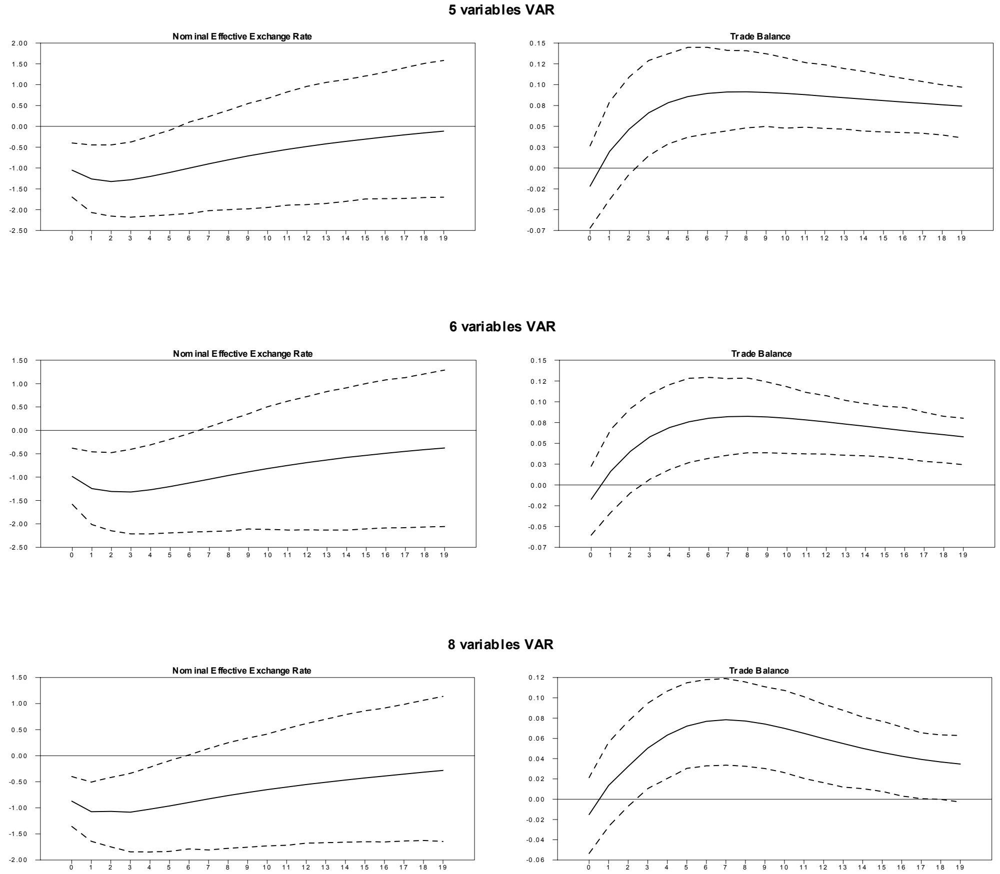
[Pure-Sign Approach]



Notes: The figures show the Impulse Responses of the Current Account to a depreciation shock, an equity shock and a housing shock using the pure-sign restriction approach in a 8 variables VAR. The sign restrictions are the same as the ones in figures 4.A, 4.B and 4.C. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

Figure 8. Impulse Responses to Nominal Effective Exchange Rate Shock (5-, 6- and 8-variable VAR)

[Pure-Sign Approach]



Notes: The figures show the Impulse Responses of the Trade Balance to a depreciation shock when the model is specified with the Nominal Real Effective Exchange Rate instead of the Real Effective Exchange Rate using the pure-sign restriction approach in a 5, 6 and 8 variables VAR. The sign restrictions are the same as the ones in figures 1, 3.A. and 4.A. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

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