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# Asset Prices, Exchange Rates and the Current Account\*

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## 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 30% of the movements of the US trade balance at a horizon of 20 quarters. By contrast, shocks to the real exchange rate have been less relevant, explaining about 9% and exerting a more temporary effect on the US trade balance. Our findings suggest that large exchange rate movements may not necessarily be the key element of an adjustment of today's large current account imbalances, and that in particular relative global asset price changes could be a potent source of adjustment.

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

**JEL Classification:** F31; F32; F40; G10.

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# 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 before improving to about 5% of US GDP in 2008. One camp of this debate points at the US as a driver of the 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 (Engel and Rogers, 2006), and a reduction in income volatility and uncertainty (Fogli and Perri, 2006). Another camp focuses 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.<sup>1</sup>

A related strand of the literature concentrates 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.<sup>2</sup> This rise in asset prices has been more pronounced in the US than in

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<sup>1</sup>Related studies point at the rapidly increasing degree of global financial integration and ensuing valuation effects on gross international asset positions (Lane and Milesi-Ferretti, 2005; Gourinchas and Rey, 2007), while others underline the role of precautionary 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). Taking a different perspective, Lane and Perotti (1998) find that fiscal policy, in conjunction with exchange rate adjustment, is important for the behavior of the trade balance, while Milesi-Ferretti and Razin (1998) stress the role of trade openness and reserves in this regard.

<sup>2</sup>We distinguish the exchange rate from other asset prices throughout the paper in order to stress that it affects

other advanced economies and many EMEs. However, following the subprime crisis that began in the summer of 2007, asset prices have experienced sharp declines in the US and, to a lesser extent, in other major economies. These developments in global and relative asset prices are likely to have impacted significantly global current account positions. Empirically, such a connection has not been formally investigated, and we know little about the potential for asset prices to induce current account movements.<sup>3</sup> In principle, asset prices are relevant for 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 thus consumption, while also making it easier for firms to finance investment opportunities, inducing a deterioration in a country's trade balance.<sup>4</sup>

The objective of the present paper is to quantify the importance of asset prices versus exchange rate shocks for the current account, and to identify the channels through which they operate from the perspective of the US. We focus our analysis on the exogenous component of exchange rates and asset prices, rather than their endogenous component, because it is well documented that only a relatively small part of the fluctuations in these variables can be explained empirically by economic fundamentals. Our empirical methodology is based on a Bayesian structural vector autoregressive (VAR) model, stemming from the work of Uhlig (2005) and Mountford and Uhlig (2009). This approach requires imposing only a small number of sign restrictions that have an economically meaningful interpretation, while avoiding some of the identification problems present in more traditional structural VAR models. We choose identifying restrictions on the basis of existing empirical literature rather than a structural model, partly because theoretical 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.<sup>5</sup> Importantly, our methodology using sign

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the current account through fundamentally different channels than, for example, equity prices.

<sup>3</sup>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).

<sup>4</sup>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).

<sup>5</sup>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

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.

The empirical findings indicate that equity market shocks and housing price shocks have been important determinants of US current account developments in the past. In the variance decomposition, the two shocks account for up to 30% of the movements of the US trade balance at a horizon of 20 quarters. By contrast, shocks to the real exchange rate explain about 9% and have exerted 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 0.85% 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 3 years. 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.

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, monetary and fiscal policy shocks have been highly relevant in the past. However, the importance of asset prices is robust to the inclusion of additional shocks, underlining that they have indeed been a major determinant of current account movements.

The findings of the present paper are linked to several contributions in the literature on current account dynamics. In particular, the focus on asset prices in the present paper is linked to the work

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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 (Nason and Rogers, 2006).

on news shocks by Beaudry and Portier (2006), in which expectations about future productivity affect current equity prices. As to the work on current account dynamics, Blanchard et al. (2005) and Obstfeld and Rogoff (2005) show that a reversal of the US current account deficit is likely to be associated with a large depreciation of the effective US dollar exchange rate, possibly in excess of 50 to 60 percent. Our paper differs from this strand of the literature at a methodological level, since we use a structural VAR to examine the drivers of the fluctuations in the US current account and their relative contribution to current account adjustment. In this sense, the contribution of our paper is distinctively empirical, and lies in the estimation of the relative importance of different sources of shocks for current account determination. While confirming the standard result that a currency depreciation is associated with part of the adjustment process in the current account, our results also suggest that a massive exchange rate adjustment of the kind suggested by the above literature may not be necessary for current account imbalances to adjust. In fact, movements in asset (equity and housing) prices have been a key 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.<sup>6</sup>

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 sign restrictions approach. Section 2 also discusses 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.

## 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

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<sup>6</sup>Also, recall that the focus in much of this literature is on the endogenous response of the current account and the exchange rate to exogenous shocks in the economy. Because exchange rates (and asset prices) are generally difficult to explain in terms of conventional economic fundamentals and the bulk of their fluctuations is hardly attributable to structural forces, we focus in this paper on the exogenous shocks to exchange rates and asset prices rather than their endogenous component.

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.<sup>7</sup>

## 2.1 VAR model

Consider the reduced form VAR

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

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  $t = 1, \dots, T$ . In practice, we also allow for an intercept term in the VAR, which is omitted in equation (1) for simplicity.

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

$$\Sigma = E[u_t u_t'] = AE[e_t e_t']A' = AA'. \quad (2)$$

We need at least  $n \times (n-1)/2$  restrictions on  $A$  to achieve identification. A 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 Sign restriction approach

Uhlig (2005) and Mountford and Uhlig (2009) achieve identification of the above VAR model imposing sign restrictions on the impulse responses of a set of variables. Uhlig (2005, Prop. 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  $a = \tilde{A}q$ , where  $\tilde{A}\tilde{A}' = \Sigma$ , and  $\tilde{A}$  is the lower triangular Cholesky factor of  $\Sigma$ .

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<sup>7</sup>Ideally one would 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.

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

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

where  $r_s$  is the matrix of impulse responses at horizon  $s$ . Sign restrictions can be imposed on a subset of the  $n$  variables over the horizon  $0, \dots, S$  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  $\tilde{A}q$ .<sup>8</sup> 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).<sup>9</sup> Using each  $q$  draw, we compute the associated  $a_j$  vectors and calculate the impulse responses as described in equation (3). If all 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 until we obtain 1000 draws that satisfy the restrictions, and error bands are calculated based on the draws kept.

Let us now turn to the more general case where we wish to identify multiple shocks, say  $m$ . In our empirical work, we identify up to three structural shocks, so that  $m = 3$ . In this case, we can characterize an impulse matrix  $[a^{(1)}, a^{(2)}, a^{(3)}]$  of rank 3. This can be accomplished by imposing

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<sup>8</sup>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.

<sup>9</sup>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, 2009).



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  $[a^{(1)}, \dots, a^{(m)}] = \tilde{A}Q$ , with the  $m \times n$  matrix  $Q = [q^{(1)}, \dots, q^{(m)}]$  of orthonormal rows  $q^{(j)}$ , i.e.  $QQ' = I_m$ . Mountford and Uhlig (2009) 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  $\Sigma$  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  $\tilde{A}$ , and the  $n$ -dimensional column vector  $r_{is} = [r_{1i}, \dots, r_{ni}]$ . The  $n$ -dimensional impulse response  $r_{as}$  at horizon  $s$  to the impulse vector  $a^{(s)}$  is given by

$$r_{as} = \sum_{i=1}^n q_i r_{is} \quad (4)$$

where  $q_i$  is the  $i$ -th entry of  $q = q^{(s)}$ .

To identify an impulse matrix  $[a^{(1)}, a^{(2)}, a^{(3)}]$ , identify  $a^{(1)}, a^{(2)}$  and  $a^{(3)}$  using the relevant sign restrictions  $a^{(1)} = \tilde{A}q^{(1)}$ ,  $a^{(2)} = \tilde{A}q^{(2)}$  and  $a^{(3)} = \tilde{A}q^{(3)}$ , 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 are not satisfied is discarded. This procedure is repeated until 1000 draws are obtained that satisfy the restrictions; error bands are calculated based on the draws that are kept.

From a methodological perspective, the sign restriction approach has several advantages. In particular, the results are independent of the chosen decomposition of  $\Sigma$ . This means that a different ordering of the variables does not alter the results. In addition, this method involves simultaneous estimation of the reduced-form 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 - *inter alia* - Faust (1998), Canova and Pina (1999), and Canova and De Nicoló (2002, 2003). Faust (1998) imposes sign restrictions only on impact. In contrast, Canova and Pina (1999) and Canova and De Nicoló (2002,

2003) 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, 2003), the identification in Uhlig (2005) is based on impulse responses and not on cross-correlations.

It is important to note that the implementation of the sign restrictions approach in Uhlig (2005) and in this paper has some shortcomings due to the key role of the median response used to examine the dynamics of the VAR model. Median responses for different shocks and horizons may combine information from several identification schemes. This implies that the required orthogonality condition of the shocks may not be satisfied when using the median response, potentially invalidating the variance decomposition. Fry and Pagan (2007) proposed an alternative method to overcome this problem, by choosing a response as close as possible to the median while imposing that the responses are generated from one single identifying matrix  $Q$ . We adopt the Fry-Pagan approach in the robustness analysis later in the paper.

## 2.4 The empirical model

Using quarterly data over the sample period 1974Q1-2008Q2, we consider the VAR model for

$$Y_t = [ c - c^* \quad p - p^* \quad i - i^* \quad reer \quad eq - eq^* \quad h - h^* \quad TB ]', \quad (5)$$

where an asterisk refers to a variable calculated for the G7 countries without the US (*G7 ex US*). We define  $c - c^*$  as the difference of the log private consumption of the US and *G7 ex US*;  $p - p^*$  is the difference between the log price level in the US and *G7 ex US*;  $i - i^*$  is the difference between the short-term interest rate in the US and the short-term interest rate in *G7 ex US*;  $reer$  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 log of equity prices in the US and the log of equity prices in *G7 ex US*;  $h - h^*$  is the difference between the log of nominal US housing prices and the log of nominal housing prices in *G7 ex US*; and  $tb$  is the US trade balance divided by GDP. A detailed description of these variables (and the weights used) is provided in the next section. 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.

Our modelling strategy involves starting from a simple 5-variable VAR (“benchmark” VAR)

that includes the exchange rate but initially excludes equity and housing prices. We then examine larger VARs where the two asset price variables are added to the model specification (“augmented” VAR). We use 3 lags for each VAR, 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.<sup>10</sup>

The crucial issue is the identification of the three shocks of interest, i.e. the exchange rate shock and the shocks to equity prices and housing prices. We focus our analysis on the exogenous component of exchange rates and asset prices, rather than their endogenous component, because it is well documented that only a relatively small part of the fluctuations in these variables can be explained empirically by economic fundamentals. Put another way, a large part of the variation in exchange rates and asset prices is driven by their own shocks, which cannot be easily linked to the state of the economy. For example, Artis and Ehrmann (2006) show that the exchange rate is largely driven by its own shocks for several major economies in the context of a structural VAR, while the vast literature on explaining movements in equity prices indicates that economically meaningful variables do not explain more than 10-15% of the variation in equity returns (e.g. Cochrane, 2005, and the references therein). Moreover, large and persistent bubbles (deviations from fundamentals) in house prices have long characterized the real estate markets of the US, the UK and other major economies (e.g. Case and Shiller, 1989; Muellbauer and Murphy, 1997).

It is worth noting, however, that in theory current account adjustment could simply occur through an endogenous response of, say, the real exchange rate. To give an example, even if there are no exogenous shocks to the real exchange rate at all (where shocks are conceptualized as sudden changes in portfolio preferences or nonfundamental fluctuations, related to noise trading or bubbles), endogenous exchange rate changes may still play an important role in current account adjustment. Such endogenous changes might reflect, for instance, a currency risk premium that is negatively related to the size of the current account balance; in the context of equation (1), this endogenous dynamics would be captured by  $B(L)$  instead of  $u$ . In empirical work, the explanatory power in equations for asset prices is generally miniscule, implying that the systematic response

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<sup>10</sup>The SIC generally suggests 1 or 2 lags, whereas the AIC usually selects 2 to 3 lags. We decided to estimate all VAR models with 3 lags to ensure that the residuals are well behaved. Qualitatively, however, the impulse responses calculated when using 2 lags or 4 lags are identical to the ones generated from the VAR with 3 lags. In addition, we also experimented for slightly different values of  $S$ , obtaining very similar results.

of asset prices and exchange rates does not, in the absence of shocks, exhibit strong stabilizing properties. Therefore, to reiterate, our focus on exogenous shocks rather than the endogenous response to disequilibria is rationalized purely on the basis of these stylized empirical facts.

There are two conceptual challenges in identifying shocks to exchange rates and asset prices. First, the sign restrictions imposed to identify these shocks should be economically meaningful. To this end, we impose sign restrictions that have received 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 an appreciation of the real effective exchange rate are that the real effective exchange rate increases (i.e. appreciates), the short-term interest rate differential between the US and the other G7 countries decreases, the price differential between the US and the other G7 countries decreases, and relative domestic consumption increases. The rationale for these restrictions stems from the perspective of a monetary policy reaction function: an appreciation should reduce import prices and domestic inflation, thus requiring a decrease in domestic short-term interest rates and thereby stimulating consumption.

**Table 1. Identification of shocks through sign restrictions**

Shock:	$c - c^*$	$p - p^*$	$i - i^*$	$reer$	$eq - eq^*$	$h - h^*$	$TB$
Appreciation	+	-	-	+			
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 shock, reflecting a canonical wealth channel (e.g. Di Giorgio and Nisticó, 2007).

Similarly, a positive relative housing price shock is identified by restricting the relative housing price index, the price differential, relative consumption and relative interest rates not to decrease. These restrictions are consistent with the impulse responses and the theoretical model of Iacoviello

(2005), which shows that shocks to house prices induce an increase in interest rates and inflation. However, Iacoviello's (2005) theory does imply that a positive house price shock can induce either an increase or a decrease in consumption depending on the parameter values of the calibration, although empirically he finds that consumption responds positively to house price shocks, consistent with our identifying assumption and with the basic intuition of wealth effects.<sup>11</sup>

It is important to note that, from a theoretical point of view, rising house prices have quite different effects than rising equity prices. The main reason is that housing does not represent just an asset, but also a significant share of households' future expenditure burdens. Thus, while there are agents who benefit from rising house prices (if they have overinvested in housing relative to their longer-term demand for housing services) there are also those in the opposite camp. This precludes uncontroversial predictions about the sign of the income or wealth effect, although there are likely to be positive effects on aggregate consumption through the impact of housing wealth on household collateral and, hence, liquidity constraints.

The impulse responses for variables on which sign restrictions are not imposed are unrestricted; these are the blank cells in Table 1. In particular, the response of the trade balance is unrestricted, which is the main focus of our analysis. Since equity and housing shocks are asset prices shocks, the signs of the responses of relative consumption and relative interest rates are the same. However, the sign of the response of equity prices after a housing shock is uncertain and viceversa, and the responses of the VAR variables could be the same for equity and housing shocks. In order to uniquely identify the two asset prices shocks in the cases when all of the responses to these two shocks have the same sign, we identify the equity shock as the shock with the largest contemporaneous impact on relative equity prices. The latter seems a straightforward and uncontroversial assumption to disentangle the two shocks, and follows the approach used by Peersman (2005) in the context of identifying supply and oil shocks.

It is well known that identification of shocks is a very difficult task in the structural VAR literature. We impose a small number of sign restrictions and avoid imposing zero restrictions, which would be difficult to rationalize in our context on the basis of conventional theories of the current account. However, a specific concern for this paper is that the shocks we identify may at

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<sup>11</sup>Note also that Iacoviello (2005) adopts a closed-economy model, and hence the endogenous variables are not defined with respect to another country. However, the basic results of the analysis are unaltered when generalizing to an open-economy setting (e.g. Punzi, 2006).

least partly reflect other shocks; for instance an increase in equity prices may be due to a positive productivity shock. We argue that other shocks differ fundamentally from equity shocks. In the case of a productivity shock, equity prices may also increase, but contrary to equity price shocks, a productivity shock should lower domestic prices and domestic interest rates, rather than raise them. We will return to a detailed robustness analysis in Section 5, where we address these issues by allowing for additional shocks into the model.

### 3 Data

We use quarterly data over the period 1974Q1-2008Q2. The “rest of the world” series are identified by an asterisk in our notation and are calculated as weighted averages of *G7 ex US*, except for consumption, which is defined as the sum of consumption in *G7 ex US*.

Figure 1 (Panel A) shows the time series of the log of consumption in the US ( $c$ ) and *G7 ex US* ( $c^*$ ). The data on real consumption are seasonally adjusted in local currencies at 2000 price levels, and taken from the OECD. We convert the consumption series to US dollars using the average market exchange rate of 2000.<sup>12</sup>  $c^*$  is the log of the sum of consumption in *G7 ex US*. Prices and interest rates are from the *International Financial Statistics* (IFS) of the International Monetary Fund (IMF), while the real effective exchange rate is taken from the US Federal Reserve Board Statistics. The price series are based on the consumer price index (CPI), and are presented in logs in Figure 1. The short-term interest rates are 3-month money market rates. The evolution of prices ( $p$  and  $p^*$ ) and interest rates ( $i$  and  $i^*$ ) in the US and *G7 ex US* may be seen in Panels B and C, respectively, of Figure 1. Prices in the US and *G7 ex US* move together for the whole period. 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 largest differential of around 700 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 to mid 1980s, then depreciated until 1995 before appreciating again till the early 2000s; it then depreciated steadily until recent times. Note that the REER in our VAR specification is

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<sup>12</sup>We do this to preserve consistency with the prices base year.

measured in logs.

The US equity measure,  $EQ$ , is the S&P500 price index, sourced from *Bloomberg*.  $EQ^*$  is an aggregate of the stock prices in the other G7 countries weighted according to their respective (time-varying) GDP shares at purchasing power parity (PPP) values.<sup>13</sup> Panel E of Figure 1 shows  $EQ$  and  $EQ^*$ . The general trend of  $EQ$  and  $EQ^*$  is largely common, reflecting the general strong performance of equity markets around the world during the sample period, with downwards corrections being especially apparent in the early 2000s (the dotcom bubble crash) and since 2007 (the subprime crisis).

The US housing price index is obtained from the Bank for International Settlements (BIS). The index for the other G7 countries also stems from the BIS and is calculated using time-varying GDP shares at PPP weights.<sup>14</sup> Panel F of Figure 1 plots the US housing price index ( $H$ ) and the corresponding housing price index for *G7 ex US* ( $H^*$ ). While both housing price indices trend upwards during most of the sample, it is clear that the US housing price index reveals a more pronounced cycle from the early 2000s until 2007, when the US housing market started declining in the aftermath of the subprime crisis.

Finally, the US trade balance,  $TB$  series was obtained from the IFS (seasonally adjusted) and is expressed as a ratio of GDP.<sup>15</sup> Panel G 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. Since then the trade balance has improved slightly, at the same time when the US equity market and, to a lesser extent, the housing market started to decline.

## 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

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<sup>13</sup>Specifically, we use the FTSE 100 for the UK, the CAC40 for France, the DAX for Germany, the MIB for Italy, the S&P/TSX for Canada, and the Nikkei 225 for Japan.

<sup>14</sup>Due to missing data from 2005 to 2008 for Germany and Italy we updated the database from Hypoport and ISTAT, respectively.

<sup>15</sup>The US GDP series is taken from the IFS, and it is expressed as seasonally adjusted.

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 consumption, relative prices and relative interest rates, i.e.  $Y_t = [c - c^* \ p - p^* \ i - i^* \ reer \ TB]'$ .<sup>16</sup>

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 affected by model uncertainty, it is fairly standard in this Bayesian VAR literature to report the 16th and 84th quantiles (e.g. Uhlig, 2005).

The impulse responses in Figure 2 suggest that a positive shock (an appreciation of the US dollar) improves the US trade balance slightly upon impact, consistent with a J-curve effect. After this initial reaction, the trade balance deteriorates from 2 quarters onwards and shows a statistically significant negative reaction between 5 and 11 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 1.0% real appreciation of the US dollar worsens the US trade balance by about 0.1% of US GDP. In the context of the current large deficit that characterizes the US trade balance, these estimates imply that even a relatively high real depreciation of the US dollar, for instance by 10%, would improve the US trade balance by a modest 1.0%.

Through what channels does the exchange rate influence the trade balance? The imposition of our sign restrictions implies that the appreciation shock generates a small decrease in the US interest rate relative to the average rest-of-the-world interest rate. 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 positive real exchange

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<sup>16</sup>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 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 fairly robust when increasing the size of the VAR in subsequent estimations.



rate shock induces a downward movement in interest rates, which according to uncovered interest rate parity (UIP) should imply expectations of a subsequent currency appreciation. However, the US dollar depreciates after the initial appreciation, 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 initially lowered by the monetary authority after a currency appreciation due to lower inflationary pressures, and are then subsequently raised as price pressures increase. Another interpretation is that the current account adjustment occurs via expenditure switching effects in response to the exchange rate shock. Although this effect is not immediately transparent in the VAR model because we do not disaggregate exports and imports in our measure of the trade balance, the results are indirectly suggesting that the relatively small response of the current account to exchange rate shocks is consistent with existing empirical evidence that price elasticities of US export and imports are small (Campa and Goldberg, 2009). Moreover, while it is worth mentioning expenditure switching effects in thinking about the channels of influence from the exchange rate to the trade balance, we prefer the first interpretation above since the expenditure-switching channel is likely to be largely endogenous, rather than driven by exogenous shocks.<sup>17</sup>

Next, we add relative equity prices and house prices in the VAR, which is now the “augmented” 7-variable VAR of equation (5). Figures 3.A to 3.C show the impulse response functions for real exchange rate shocks, relative equity shocks and housing price shocks, respectively. The trade balance impulse response for the real exchange rate shock (Figure 3.A) in this 7-variable VAR is similar to the VAR without asset prices. The impulse response to an equity market shock (Figure 3.B) reveals that the effect of an equity price shock on the US trade balance is quantitatively comparable to and more persistent than a real exchange rate shock. A positive relative equity shock of about 1.8% (10%) lowers the US trade balance by 0.15% (0.85%) after 10 quarters, which is comparable to the impact of the real exchange rate shock on the trade balance. This is particularly important given that relative equity market changes, especially throughout the 1990s, have been

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<sup>17</sup>For example, in the presence of real shocks that are specific to one country (such as productivity shocks, or fiscal shocks, etc.), nominal exchange rate changes can induce adjustment of relative prices of goods across countries. A country that experiences a productivity increase should experience a decline in the price of its output that induces a switch in expenditures toward the domestic product.

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 worsening the trade balance. Interestingly, a relative equity market shock does not have a significant effect on the real exchange rate of the US dollar. Finally, a positive US housing market shock (Figure 3.C) leads to a significant and persistent deterioration of the US trade balance, with a 1.0% (10%) rise in US housing prices lowering the US trade balance by 0.16% (1.6%) after 11 quarters. US housing price developments appear to affect the other variables in the VAR in a similar fashion as US equity market shocks. Also, note that a positive housing price shock raises relative equity prices somewhat. This is indicative of ensuing wealth effects of the housing market increase not only to raise consumption but also the demand for equities, thus exerting an upward pressure on equity prices.

## 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 30% of the trade balance is explained by equity market shocks and housing price shocks at a horizon of 20 quarters. By contrast, at most 9% of the US trade balance is accounted for by shocks to the US dollar real exchange rate at the 20-quarter horizon.

**Table 2. Variance decomposition for the US trade balance (7-variable VAR)**

Shocks			
Steps	Appreciation	Equity	Housing
4 quarters	14.5	9.6	16.4
8 quarters	12.1	10.7	14.4
12 quarters	10.6	14.1	13.5
16 quarters	9.8	16.2	12.7
20 quarters	9.2	17.3	12.4

In summary, the findings 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. However, as one would expect, real exchange rate movements have also contributed to the variation in trade balance, exerting a slightly more moderate effect.

## 5 Robustness and Extensions

Empirical results are often dependent on underlying assumptions and variables definitions. A key advantage of the Bayesian VAR with sign restrictions is that it requires only a small number of identification restrictions, which are relatively uncontroversial. While we have described the identifying assumptions in detail above, we now turn to discussing various alternative variables definitions and also different VAR specifications.

### 5.1 Allowing for additional shocks

An important issue refers to the robustness of our findings to the inclusion of further variables in the VAR and the allowance for additional shocks. First, there is some evidence that productivity shocks have been an important determinant of current account positions (Bussiere et al., 2005; Corsetti et al., 2006). Also, 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. Second, an important source of fluctuations in exchange rates and, thereby, the current account stems from monetary shocks, which have been the subject of a large empirical literature (e.g. Bems et al., 2007). Third, fiscal policy shocks may also be an important driver of current account fluctuations (e.g. Corsetti and Müller, 2008).

We check the sensitivity of our core results to the inclusion of all these three shocks: productivity, monetary and fiscal. With respect to productivity shocks, we add relative GDP to the 7-variable VAR described in the previous section.<sup>18</sup> The short-run sign restrictions imposed for  $S = 2$  quarters to identify a positive productivity shock are that relative output and relative consumption increase, and that the price differential and relative interest rate decrease. The first two restrictions should be obvious. The third restriction is motivated by the “mainstream” model of price 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.

With respect to 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 has been mentioned in the debate on global imbalances as a relevant factor (e.g. Bems et al., 2007). To allow for this channel in the VAR framework, we introduce monetary policy shocks. The sign restrictions imposed are the same as those in the literature discussed above, such that a monetary tightening shock increases short-term interest rates, lowers prices, decreases relative output, and raises the US real effective exchange rate.

Finally, we identify fiscal policy shocks following Peersman and Straub (2009). Specifically, we impose the restrictions that a positive fiscal policy shock increases relative output, reduces private consumption expenditure, increases relative prices and relative interest rates. Note that Peersman and Straub (2009) do not impose a sign restriction on consumption but they document that empirically a positive fiscal policy shock reduces consumption expenditure, which is one of our sign restrictions to identify the fiscal policy shock.

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<sup>18</sup>The data are from the IFS, and the relative real GDP measure is calculated as the difference between the log of US real GDP and the log of the sum of real GDP for *G7 ex US*.

**Table 3. Variance decomposition for the US trade balance**

Steps	Shocks					
	Appreciation	Equity	Housing	Productivity	Mon. Pol.	Fiscal
4 quarters	12.8	10.6	17.5	7.2	8.5	13.2
8 quarters	11.5	10.9	17.1	8.1	9.0	12.4
12 quarters	8.9	12.1	16.8	9.3	10.4	11.3
16 quarters	8.5	13.7	16.5	12.8	10.9	10.2
20 quarters	7.9	16.1	15.0	14.2	11.2	9.7

Table 3 shows the variance decomposition for our extended, 8-variable VAR specification including also productivity, monetary and fiscal 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 and housing shocks amount to about 31% after 20 quarters, which is very similar to our baseline result obtained in the previous section. Second, we also find support for the findings of the previous literature in that productivity, monetary and fiscal policy shocks have been exerting substantial effects, explaining up to 14%, 11% and 10%, respectively, of the US trade balance after 20 quarters. Thus, overall, asset prices are confirmed as a key driver of the US trade balance while exchange rate shocks appear to have been somewhat less important, with about 8% of the variance of the trade balance being driven by them after 20 quarters.

## 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 (e.g. changes in returns) from trade in goods and services; thus, our preferred measure remains the trade balance.

Figure 4 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, 2007).

Second, we investigate the sensitivity of the results by using nominal (as opposed to real) effective exchange rate shocks. 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 5 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 exchange rates, with the elasticities being only marginally different.

### 5.3 Alternative Impulse Response Calculation

As noted by Fry and Pagan (2007), there are conceptual problems related to the selection of the median impulse response, resulting from the multiplicity of impulse vectors. In particular, they show that median impulse response functions may be generated by different impulse matrices  $Q$ . Thus, the median of the impulse responses may not be generated by a single model and, as a consequence, the shocks identified are no longer orthogonal to each other. Fry and Pagan (2007) propose to calculate the response that is as close as possible to the median response while being generated from one single matrix  $Q$ .<sup>19</sup>

We check the sensitivity of our results by applying the method suggested by Fry and Pagan (2007) to the 7-variable VAR in equation (5). The results are displayed by the dotted line in Figure 6, and the variance decomposition results are shown in Table 4. We find that our core results are not qualitatively changed when applying this method.

**Table 4. Variance decomposition using the Fry-Pagan method**

Steps	Shocks		
	Appreciation	Equity	Housing
4 quarters	13.0	10.4	17.2
8 quarters	11.3	11.3	18.3
12 quarters	12.2	15.2	16.8
16 quarters	11.0	17.3	14.3
20 quarters	10.5	18.7	13.8

<sup>19</sup>This approach has been applied by Enders, Müller and Scholl (2008). Note, however, that Enders et al. (2008) do not find a significant difference in results between their baseline specification based on the median response and the Fry-Pagan method, suggesting that the problems discussed here do not always affect empirical work.

## 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 both of 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 price shocks.

To address this question, the paper has employed a Bayesian VAR model, which requires imposing only a minimum of sign restrictions that have a meaningful economic interpretation. Our main finding is that asset prices are an important driver of the US trade balance. In fact, 30% of the movements of the US trade balance after 20 quarters can be accounted for by asset price shocks, and only about 9% by shocks to the US dollar real exchange rate. These results are robust to various extensions and alternative specifications. For instance, while also US productivity shocks, monetary and fiscal policy shocks are found to exert a significant effect on the US current account, asset prices remain a key driver of the US current account in all VARs estimated.

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 and may well 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 fall in US asset prices or a rise in asset prices outside the US.

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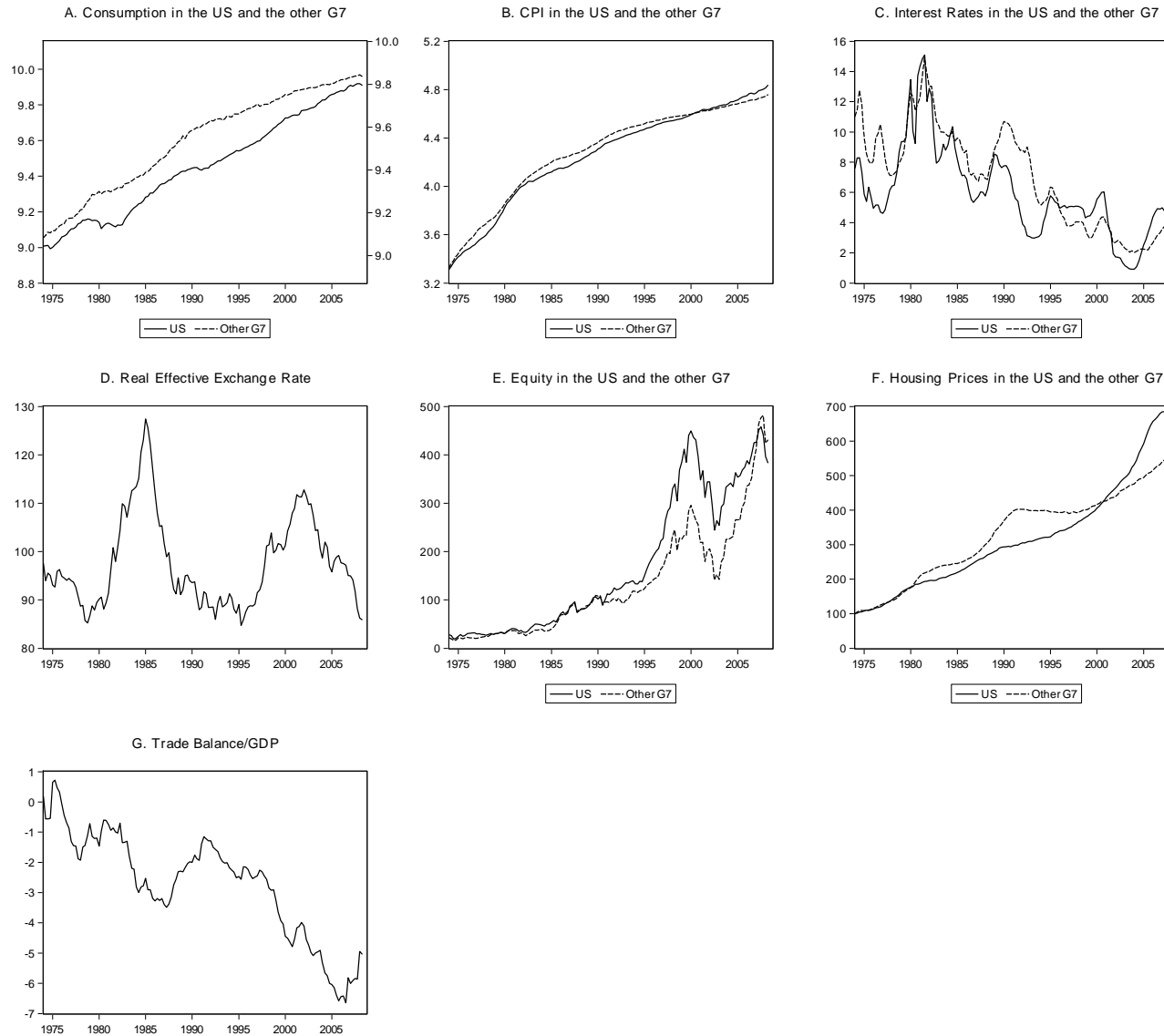


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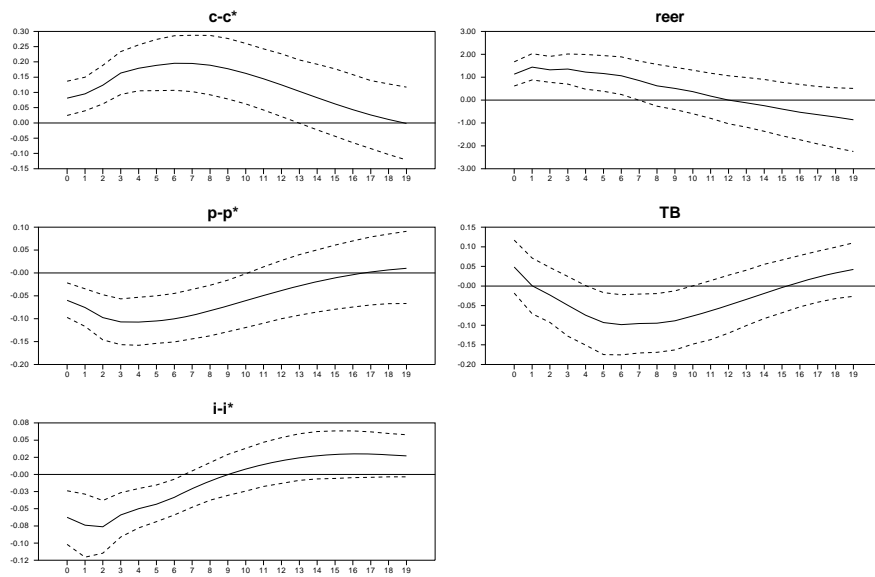
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**Figure 1. Data**

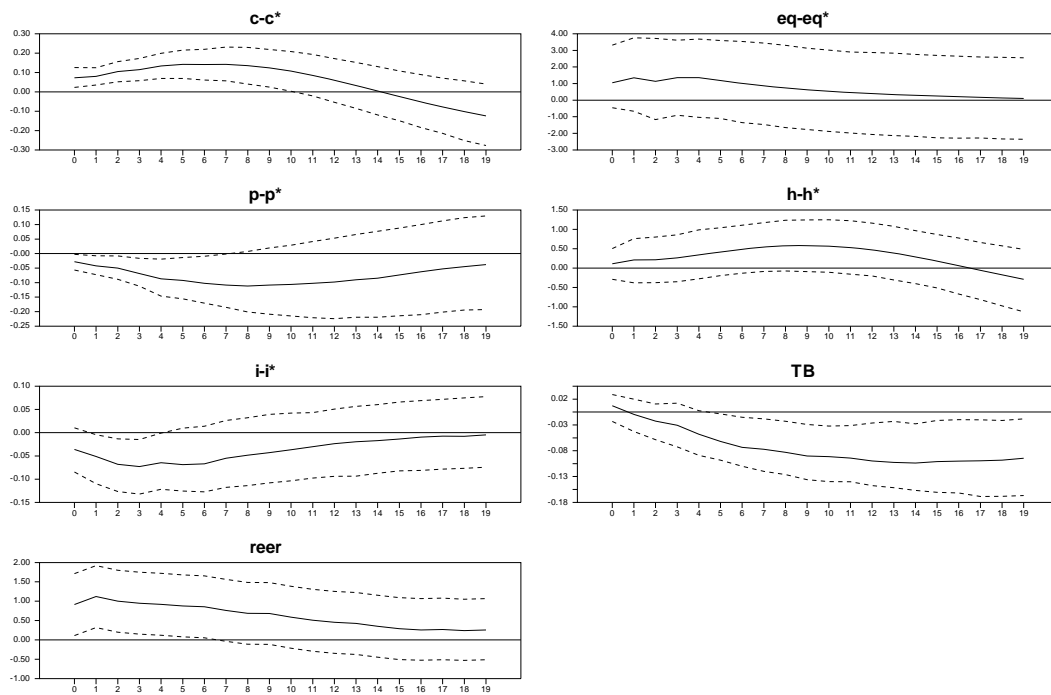


**Figure 2. Impulse Responses to Exchange Rate Shock (5-variable VAR)**



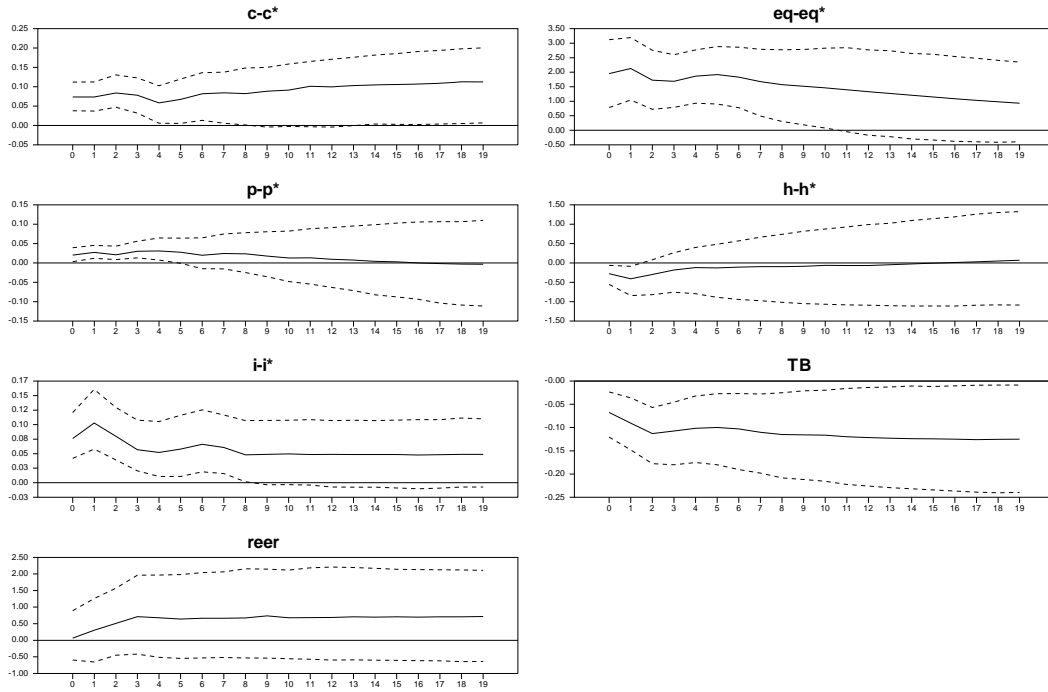
**Notes:** The figure shows the Impulse Responses to an appreciation shock. The responses of the real effective exchange rate and relative consumption were restricted not to be negative and the responses of relative prices and the relative interest rate were restricted not to be positive for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

**Figure 3.A. Impulse Responses to Exchange Rate Shock (7-variable VAR)**



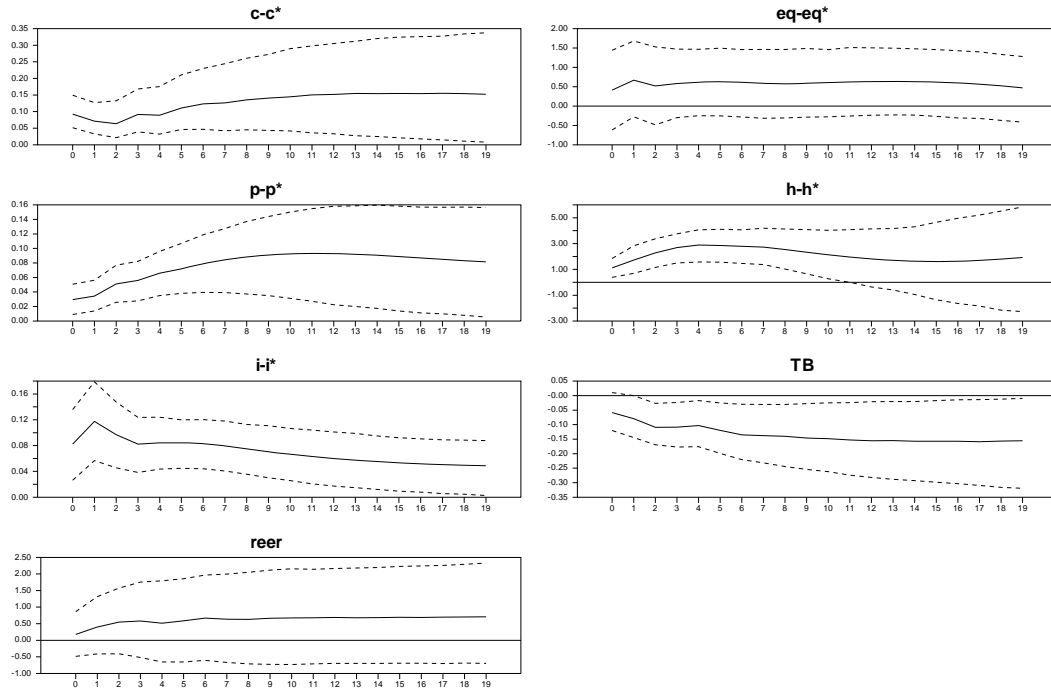
**Notes:** The figure shows the Impulse Responses to an appreciation shock. The responses of the real effective exchange rate and relative consumption were restricted not to be negative and the responses of relative prices and the relative interest rate were restricted not to be positive for two quarters. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

**Figure 3.B. Impulse Responses to Equity Shock (7-variable VAR)**



**Notes:** The figure shows the Impulse Responses to an equity shock. The responses of 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.

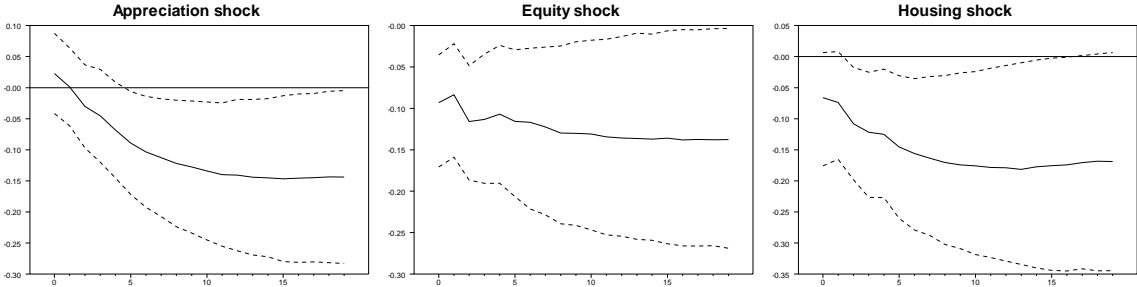
**Figure 3.C. Impulse Responses to Housing Shock (7-variable VAR)**



**Notes:** The figure shows the Impulse Responses to a housing shock. The responses of relative housing, relative consumption, relative prices 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.

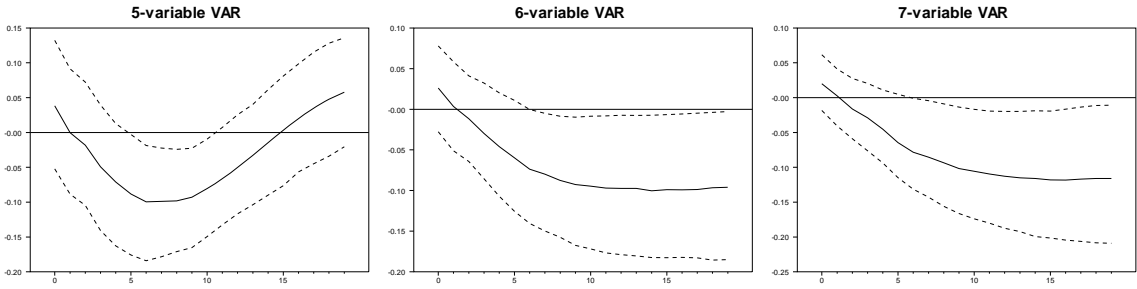


**Figure 4. Impulse Responses to the Current Account to Exchange Rate, Equity and Housing Shocks (7-variable VAR)**



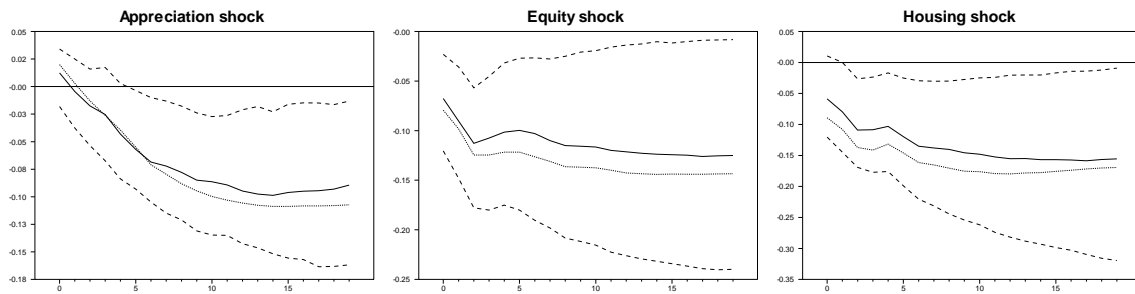
**Notes:** The figures show the Impulse Responses of the Current Account to an appreciation shock, an equity shock and a housing shock in a 7-variable VAR. The sign restrictions are the same as the ones in figures 4.A, 4.B and 4.C. The solid line is the median of the posterior distribution and the dashed lines represent the 16% and 84% quantiles.

**Figure 5. Impulse Responses to Nominal Effective Exchange Rate Shock (5-, 6- and 7-variable VAR)**



**Notes:** The figures show the Impulse Responses of the Trade Balance to an appreciation shock when the model is specified using the Nominal Effective Exchange Rate instead of the Real Effective Exchange Rate in a 5, 6 and 7 variables VAR. The sign restrictions are the same as in figures 2, 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.

**Figure 6. Comparison Median and Median Solution**



**Notes:** The figures show the Impulse Responses of the Trade Balance to an appreciation shock, an equity shock and a housing shock in a 7-variable VAR. The sign restrictions are the same as 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. The dotted line represents the impulse responses generated by the median solution proposed by Fry and Pagan (2007).