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Exchange Rates, Foreign Currency Exposure and Sovereign Risk

Kerstin Bernoth* and Helmut Herwartz†

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Abstract

We quantify the causal link between exchange rate movements and sovereign risk of 16 major emerging market economies (EMEs) by means of structural vector autoregressive models (SVARs) and conditional on data from 10/2004 through 12/2016. We apply a novel data-based identification approach of the structural shocks that allows to account for the complex interrelations within the triad of exchange rates, sovereign risks and interest rates. The direction and size of the response of sovereign risk to FX rate movements depend on the considered type of exchange rate measure and on the size of the net foreign currency exposure of an economy. A depreciation of the domestic currency against the USD increases sovereign risk. In contrast, when looking at the nominal effective exchange rate that is unrelated to changes of USD exchange rate, we find in general no significant effect. We conclude that the ‘financial channel’ is more important in the transmission of exchange rate shocks to sovereign risk in comparison with the traditional ‘net trade channel’. Moreover, we confirm the prime role of the currency mismatch of the non-public sector for the strength of the ‘financial channel’.

Keywords: Exchange rates, sovereign risk, foreign currency exposure, structural VAR.

JEL Classification: C32, G12, G15, G23.

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1 Introduction

Financial market integration and liberalization have markedly increased with the rise of globalization. In this process, the external balance sheets and foreign currency exposures of public and private sectors have also increased in many countries. For emerging market economies (EMEs), in particular, such exposures are very often sizeable (Lane & Shambaugh, 2010; Bénétrix et al., 2015; Chui et al., 2016; Du & Schreger, 2016). As a result, the assets of the government and corporate sectors of these countries are exposed to valuation effects in response to exchange rate fluctuations. Therefore, not only domestic factors but also increasingly international factors, and in particular the value of their currency, are likely to determine the solvency of these countries.

In which direction FX rate adjustments affect sovereign risk is theoretically ambiguous. According to the traditional Mundell-Fleming model, a domestic currency depreciation is associated with a deterioration of a country's terms of trade. This increases net trades, which has a positive impact on a country's international competitiveness ('net trade channel'). As a result, aggregate output rises, thus reducing pressures on sovereign risk. However, a depreciation of the domestic currency can also have an adverse effect on a country's probability of default. When a sovereign holds an unhedged negative net foreign asset position, i.e., it has more foreign currency liabilities than foreign currency assets, a depreciation of its currency causes negative wealth effects. Similarly, when the balance sheet of the corporate sector has a currency mismatch, Krugman (1999) and Céspedes et al. (2004) show that a depreciation of the domestic currency worsens the credit position of firms by reducing the value of collateral and leading to an increase in the cost of credit and credit contraction.¹ This worsening of domestic financial conditions dampens economic activity, and increases ultimately the risk of default as perceived by investors. Moreover, Du & Schreger (2016) argue that, given a negative impact of currency depreciation on corporate balance sheets, the government is more inclined to explicitly default than to inflate in order to reduce the fiscal burden of government debt repayments. As a result, the described so-called 'financial channel' of FX rates adjustments can be regarded as a potential offset to the traditional

¹A credit contraction might be triggered by reduced cross-border capital inflows (Bruno & Shin, 2015a), or by a decline in the leverage of the domestic banking sector (Bruno & Shin, 2015b).

‘net trade channel’.² Whether a depreciation of the domestic currency increases or reduces sovereign risks depends on which of the two channels predominates.

In this work, we analyze empirically the effect of a currency depreciation on sovereign risk using monthly data of 16 major EMEs for the 10/2004 to 12/2016 period. We investigate the relative strength of the ‘net trade channel’ and the ‘financial channel’ by looking at alternative exchange rate measures and by differentiating countries with respect to their net foreign currency exposure. Similar to Hofmann et al. (2016, 2020), we argue that the response to a change in the nominal effective exchange rate (NEER) reflects the joint effect resulting from the ‘net trade’ and the ‘financial channel’ and could therefore be positive or negative. We argue that the size of a country’s net foreign currency exposure has a crucial role in determining the relative importance of these two channels and thus the direction in which an appreciation of the domestic currency affects sovereign risk. When a country is net short in a global currency, the response of sovereign risk measures to the bilateral exchange rate against that currency is expected to be dominated by the ‘financial channel’. We focus on the USD FX rate because the USD is the dominant medium of international financing for most EMEs and changes in the USD exchange rate are likely to have the largest valuation effects on sovereign and corporate balance sheets (Bénétrix et al., 2015; Kuruc et al., 2017; McCauley et al., 2015). Thus, in countries with negative net foreign currency exposure, the response of sovereign risk measures should increase with a depreciation of the domestic currency against the USD.³

While the empirical literature has been studying the balance sheet implications of FX rate movements comprehensively (see, amongst others, Lane & Milesi-Ferretti, 2007; Gour-

²The ‘financial channel’ in the related literature is also called the ‘risk-taking channel’ of FX rates introduced by Bruno & Shin (2015a).

³A recent strand of literature adds that also the export pricing paradigm matters for the transmission of exchange rate shocks (Georgiadis & Schumann, 2019). Effects of changes in the USD exchange rate on a country’s terms of trade should be small, if it has a large share of its exports denominated in USD. A data set collected by Gopinath (2015) lists the export invoice currencies for 11 of the 16 countries contained in our sample. Within this subsample, the mean export share in USD is around 50 percent and ranges between 13 and 98 percent. By pooling our estimation results across a panel of countries with very heterogeneous levels of exports shares in USD, we assure that our estimation results are not obscured by the effect of the potential export pricing paradigm. Moreover, while the pricing paradigm matters for the significance of the ‘net trade channel’, it does not so for the ‘financial channel’. Thus, independent of the dominant export invoicing currency of countries should the size of the net foreign asset position affect the *relative* importance of these two channels.

inchas & Rey, 2007), there are still gaps in both the content and the methodologies applied that we close with our work. First, the literature on the empirical determinants of sovereign risks of EMEs usually focuses on other macroeconomic and financial variables, like e.g., GDP growth, government debt, or the volatility of the terms of trade (Eichengreen & Mody, 1998; Ferrucci, 2003; Hilscher & Nosbusch, 2010). Only a few studies analyze to what extent FX rates affect the sovereign risks of EMEs. Longstaff et al. (2011) and Hofmann et al. (2016, 2020) find that the sovereign’s credit spread of a representative EME increases in response to a depreciation of its currency towards the USD, which underlines the relevance of the ‘financial channel’ in affecting a country’s financial conditions. However, to our knowledge, no existing study explicitly analyzes, if the size of net foreign currency position of domestic borrowers matters for the direction and magnitude of exchange rate effects on sovereign risk. By filling this gap, we complement the finding of Du & Schreger (2016) that a higher level of external foreign currency corporate debt is generally associated with larger sovereign risk.

Second, for the research question at hand, unrestricted bidirectional causalities that link FX rates and sovereign risks are immediate. As shown in Della Corte et al. (2016), an exogenous increase in sovereign risk leads to a significant depreciation of the FX rate. As such, this bidirectional relationship could invoke amplification effects and sizeable moves in FX rates and sovereign risk measures. Hence, self-reinforcing feedback loops between FX rate adjustments and domestic financial easing could develop. Fostering joint endogeneity even further, financial variables, like FX rates, interest rates, and CDS spreads, typically respond quickly to news entering integrated markets. Common triangular identification schemes, such as those used in Hofmann et al. (2016), do not account for such potentially bidirectional causalities and might provide biased structural estimates. We quantify the causal link between FX rate movements and sovereign risk of major EMEs by means of a panel of structural vector autoregressive models (SVARs). Our identification approach takes advantage of recent contributions to data-based identification of SVARs that build upon the uniqueness of independent structural shocks in linear non-Gaussian systems (Comon, 1994). The advantage of this novel approach is that it allows the evaluation of unrestricted transmission channels.⁴

⁴Various approaches to the extraction of independent components in linear non-Gaussian systems in SVARs have been suggested (Moneta et al., 2013; Herwartz & Plödt, 2016; Lanne et al., 2017; Gouriéroux et al.,

While our identification scheme is mainly of statistical nature, the economic labelling of the identified shocks benefits strongly from unrestricted structural parameter estimates which imply that exogenous increases of country risk invoke a contemporaneous exchange rate depreciation. Within the framework of triangular model structures this particular impact effect would be subject to a zero restriction. After statistical and economic identification, we use the cross sectional dimension of the data for purposes of mean group (MG) estimation (Pesaran & Smith, 1995; Pesaran et al., 1999).⁵ In particular, after identifying unrestricted structural models, MG inference allows to test the triangularity restriction (i.e. the Cholesky factor) rather than assuming it from the outset.

We find that the direction and size of the response of sovereign risk measures to a depreciation shock depend on the type of the considered exchange rate measure and on the size of the net foreign currency exposure of an economy. In general, a depreciation of the domestic currency against the USD increases sovereign risk. In contrast, a depreciation of the effective exchange rate lacks a significant effect on sovereign risk conditional on the full cross section of sampled EMEs. However, when focusing only on countries with large negative foreign currency exposures, we find that a depreciation of the NEER induces an increase in sovereign risk. In general, a change in the NEER that is unrelated to changes of USD FX rates does not affect sovereign risk. Thus, we conclude from these results that it is primarily the USD exchange rate that matters and that the ‘financial channel’ is more important in the transmission of exchange rate shocks to sovereign risk than the traditional ‘net trade channel’. Emphasizing the existence of the ‘financial channel’, this work is the first to show that the response of sovereign risk to a depreciation shock is indeed positively related to the net foreign currency exposure of a country. More specifically, in line with Bruno & Shin (2015a) and Du & Schreger (2016) we confirm the prime role of the currency mismatch of the non-public sector for the strength of the ‘financial channel’.

The estimation results of our study provide an important contribution to the actual discussion on monetary policy spillovers. It has been observed, for instance, that the correlation between long-term interest rates in the US and EMEs has increased since the beginning

2017; Matteson & Tsay, 2017; Herwartz, 2018, 2019). For a textbook treatment of data-based contributions to SVAR identification, see Kilian & Lütkepohl (2017).

⁵See also Gambacorta et al. (2014) for an application of MG inference in panel SVARs.

of the new millennium (Sobrun & Turner, 2015; Miyajima et al., 2014; Takats & Vela, 2014; Shambaugh, 2004; Obstfeld et al., 2005; Klein & Shambaugh, 2015). There is a sizeable and emerging literature that explains this phenomenon with the existence of a ‘global financial cycle’ (Rey, 2015; Bruno & Shin, 2015a). It is argued that financial conditions in financial centers - primarily the United States - increasingly spill over to other economies through cross-border capital flows and general risk aversion. Our study confirms the existence of valuation effects shaping an additional channel, through which monetary policy of major economies affects the financing conditions in EMEs, i.e. sovereign risk premia. A depreciation of the USD that results from expansionary US monetary policy induces a decline of sovereign risk premia in long-term rates of countries that are net borrowers in USD. As a result, the correlation between long-term interest rates in these countries and the United States increases. Thus, monetary policy autonomy in countries with large foreign currency exposures is constrained.

In the next Section, we outline the employed SVAR models and our approach to identification. Section 3 discusses the structural transmission patterns that we expect economically with reference to the stylized theoretical frameworks mentioned above. Our panel database is described in some detail in Section 4. Section 5 comprises estimation results. Section 6 highlights some policy implications of this study and concludes. The Appendix provides implementation details on the estimation of the structural parameters.

2 Structural VARs with independent shocks

This section provides an overview of the approach to SVAR identification used in this study. We (i) briefly sketch our VAR model in reduced and structural form, (ii) state the identification problem, and (iii) provide the sufficient conditions for uniqueness of independent structural shocks. For a more detailed exposition of the identification approach and its implementation we refer the reader to the Appendix.

Our analysis builds on sample information from 2004M10 to 2016M12. The cross section comprises 16 major EMEs, i.e. Brazil, Chile, Colombia, Croatia, Czech Republic, Hungary, Indonesia, Korea, Malaysia, Mexico, Philippines, Poland, Russia, South Africa, Thailand

and Turkey.⁶ In case of the European countries in our sample, the Euro turns out to be the dominant currency in terms of foreign denominated asset and liability holdings and not the USD. However, for the ‘financial channel’ to work, it is the *net* foreign currency position that matters. And here we confirm the finding of Bénétrix et al. (2015) that all countries except South Africa show a (sizeable) net short position in USD, including the European countries. In contrast, the net short position in Euro is very small for most countries except for Croatia, Hungary and Poland. That means that for these three countries the ‘financial channel’ should also become active in case of currency movements against the Euro and not only the USD.

We consider country specific VARs throughout. To simplify notation we refrain from using an additional country indexation of the (S)VAR model. When it comes to the specification of a ‘global’ (i.e. mean group) structural model, however, we introduce an additional index. We consider trivariate VARs ($K = 3$) of order p augmented with exogenous influences in reduced form and structural form, respectively, as

$$y_t = \nu + A_1 y_{t-1} + \dots + A_p y_{t-p} + Gx_t + u_t, \quad (1)$$

$$= \nu + A_1 y_{t-1} + \dots + A_p y_{t-p} + Gx_t + D\xi_t, \quad t = 1, 2, \dots, T, \quad (2)$$

where $y_t = y_t^{(m)} = (y_{1t}^{(m)}, y_{2t}, y_{3t})$ is a vector of endogenous variables comprising a measure of the change in exchange rates ($y_{1t}^{(m)}$), the monthly changes of the 5-year credit default swap spread, as a measure for sovereign risk (CDS, y_{2t}), and the first difference of the domestic 3-month money market rate, which controls for the stance of domestic monetary policy (y_{3t}). We distinguish three alternative exchange rate measures: (i) the first difference of the log USD FX rate ($m = 1$); (ii) the first difference of the NEER ($m = 2$) and, (iii) similar to Hofmann et al. (2016, 2020) and Avdjiev et al. (2018) an estimate of NEER changes that are considered unrelated (orthogonal) to changes of USD FX rates (NEER_{orth} , $m = 3$). These orthogonal components are obtained as the residuals of country-level linear regressions of first differences of the NEER on first differences of nominal (log) FX rates.⁷ Throughout, we refer

⁶We started with focusing on the same 30 major EMEs as Georgiadis & Mehl (2016). We dropped from their list all countries that have not implemented a ‘de jure’ floating FX regime. Due to limited data availability, our cross section has been reduced further to overall 16 EMEs.

⁷Alternatively, one could consider the use of a financially weighted exchange rate index, as this should reflect

to exchange rates of a domestic currency in direct quotation such that upward adjustments of the exchange rate correspond to a depreciation of the domestic currency.⁸

Thus, overall we evaluate three trivariate VARs for each economy that differ with regard to its first element $y_{1t}^{(m)}$.⁹ Similar to the country index, we omit the model index m from the notation whenever appropriate. The Q -dimensional vector x_t comprises a set of exogenous covariate information. In our case, the exogenous vector consists of five variables. The first variable is the volatility index of the Chicago Board Options Exchange (VIX). To account for the economic and the financial situation in the US, we control for US industrial production, US CPI (all measured in log-changes), and the change in the US 3-month money market rate. The inclusion of these four exogenous variables controls for global risk-aversion and the global financial cycle that might simultaneously affect exchange rates and CDS spreads.¹⁰ Finally, to approximate the fiscal stance of an EME, we use data on the level of government debt as a ratio of GDP.¹¹ Since debt data are only available at an annual frequency, we include 1/12 of the annual change of the debt to GDP ratio in x_t .¹² Moreover, u_t is a serially uncorrelated vector process with mean zero and covariance Ω , ν is a vector of intercepts, which assures that our estimation results are not contaminated with (neglected) fixed effects, A_1, A_2, \dots, A_p are $K \times K$ parameter matrices, and G is a $K \times Q$ parameter matrix. Presample values y_0, y_1, \dots, y_{1-p} are assumed available.

The autoregressive representation in (1) characterizes the jointly endogenous variables

the dynamics in international trade in financial assets, which should ultimately also influence sovereign bond yields (Lane & Shambaugh, 2010). However, an important contribution of our work to the existing literature is the innovative identification of structural shocks that takes into account potential bidirectional causalities of the model variables considered. By focusing on these three exchange rate measures, we can directly compare our results with those of Hofmann et al. (2016, 2020) and Avdjiev et al. (2018) who use the same exchange rate variables but more restrictive shock identification approaches. Furthermore, one objective of this study is to distinguish the ‘financial channel’ from the traditional ‘net trade channel’. For the latter, the NEER is commonly used.

⁸NEER data is provided by the Bank of International Settlement (BIS). The NEER index is calculated on basis of bilateral exchange rates in indirect quotation and normalized to 100 in a chosen base year.

⁹At first glance it appears tempting to consider one five-dimensional SVAR instead of three three-dimensional models. However, as our interest is in the specification of a structural model, we believe that the notion of ‘structural shocks’ is somehow ill defined in systems processing related information like FX rates and NEER, where the latter depend on the former by construction.

¹⁰Rey (2015) and Miranda-Agrippino & Rey (2015) show that the global financial cycle in capital flows, asset prices and in credit growth strongly co-moves with the VIX, and is partly dictated by monetary policy conditions in the United States.

¹¹Bernoth et al. (2012) find that the level of government debt is a key determinant of sovereign risk premia.

¹²All variables have been drawn from either from Bloomberg or Datastream.

y_t conditional on their history and the exogenous variables in x_t . By assumption, the model in (1) is causal, i.e., $\det(A(z)) \neq 0 \forall |z| \leq 1$, where $A(L) = I_K - A_1L - A_2L^2 - \dots - A_pL^p$ and L is the lag operator (e.g. $Ly_t = y_{t-1}$). Including covariance estimation, the reduced-form model in (1) can be quantified consistently by means of OLS/ML estimation.

The structural model in (2) formalizes a linear transmission, $u_t = D\xi_t$ from structural shocks ξ_t to reduced-form disturbances u_t . By assumption, the matrix D is nonsingular and its diagonal elements are positive.¹³ Hence,

$$\xi_t = D^{-1}u_t \text{ and } \text{Cov}[u_t] = DD' =: \Omega. \quad (3)$$

It is well known that in a Gaussian framework ($u_t \sim N(0, \Omega)$), the identification of D necessitates further information from outside the model or some *a-priori* choice (e.g. the assumption of a recursive causal structure; Sims, 1980). An important result in Comon (1994) states that the linear transmission scheme in the left hand side of (3) allows a unique recovery of D from (estimates of) u_t in non-Gaussian systems. To take advantage of this result for identification, the following distributional assumptions are added to the SVAR representation in (2):

- (i) $\xi_t \sim F(0, I_K)$,
 - (ii) ξ_{kt} , $k = 1, \dots, K$, are jointly independent, and
 - (iii) at most one marginal distribution in F is Gaussian.
- (4)

The data-based approach to identification that we pursue in this study (for details and an embedding into the related literature see the Appendix) has been suggested in Herwartz (2018). It builds upon the result of Comon (1994), i.e. the uniqueness of independent, non-Gaussian components, and consists in determining the matrix D in a way that joint dependence among the implied shocks $\xi_t = D^{-1}u_t$ is minimal in terms of a flexible non-parametric dependence measure, namely the so-called Cramér-von-Mises (CvM) distance of Genest et al. (2007). For a formal representation of this estimator see equation (7) in the Appendix. Since the statistical identification does not build upon behavioural economic

¹³Assuming positive diagonal elements is conventional in SVAR analysis and implies that the structural analysis focusses at the effects of positive structural shocks.

relationships, the ‘shock labelling’ analysis in Section 5 provides important complementary insights into the properties of the statistically identified independent shocks.

3 Impact effects under joint endogeneity

In this Section, we discuss the theoretical transmission channels that link the contemporaneous relations among exchange rates, sovereign risk and interest rates, and subsequently discuss their empirical evidence. This guides us in plausibly determining the impact effects that one expects for exogenous shocks hitting our SVARs through exchange rates, monetary policy (3-month interest rate), and measures of sovereign risk (CDS spreads). These considerations are essential for the identification of the ‘structural shocks’ in our model. To summarize the theoretical considerations, finally, Table 1 collects impact effects of shocks that we consider to be plausible. Although our stylized SVARs are unlikely to be fully informative for the identification of monetary policy shocks or shocks to sovereign risk,¹⁴ it is instructive to distinguish these shocks from exchange rate shocks. Noticing that our identification approach is unrestricted and largely data driven, expected impact effects are at the core of the economic labeling of the statistically detected structural shocks in Section 5.

The theory of the uncovered interest rate parity (UIP) suggests that the interest rate differential between two countries will be a function of the expected change of the bilateral FX rate and a risk premium associated with holding foreign currency denominated assets. This risk premium incorporates a compensation for FX rate volatility and a compensation for credit risk of domestic currency liabilities. Thus, one expects that (i) increases in domestic interest rates relative to foreign interest rates invoke an appreciation of the domestic currency; and (ii) an increase in the country risk premium comes with a depreciation of the domestic currency (i.e. an increase of the current FX rate). However, as pointed out by Gould & Kamin (2000), potential endogeneity problems may obscure the relationship between interest rates, exchange rates, and country risk premia. More precisely, domestic interest rates may be endogenous with respect to current adjustments of the FX rate, while sovereign risk and FX

¹⁴For instance, for the identification of Taylor-rule type monetary policy shocks, a standard New-Keynesian model would suggest including a measure for economic activity (e.g. output gap) and inflation in the set of variables.

rates are endogenous to adverse monetary policy shocks. As one might argue, the described and interrelated transmission channels are of particular relevance during financial crises. In addition, empirical studies show that the UIP very often fails and that exchange rates are moving in the opposite direction than suggested by this theory, which also questions the relationship between exchange rates and interest rates (Engel, 2014). The following overview emphasizes the complex interrelationships within the triad of exchange rates, sovereign risk, and interest rates, that complicates the simultaneous identification of transmission channels.

Exchange rate shocks (depreciation)

The effect of a positive exchange rate shock (depreciation) on sovereign risk is our main research question, which we state in terms of hypotheses below.¹⁵ Probably the most established theory is that exchange rate movements affect the economy through trade effects. Along the ‘net trade channel’ a currency depreciation improves an economy’s international competitiveness and, thereby, its trade balance as long as a share of its exports is priced in domestic currency. As a result, production increases, which induces a reduction of sovereign risk. However, a depreciation of the domestic currency can also have a negative impact on a sovereign’s probability of default. When a sovereign holds an unhedged negative net foreign asset position, a depreciation of its currency causes negative wealth effects, lowering its credit worthiness. Similarly, when the domestic currency depreciates, a currency mismatch in corporate balance sheets leads to a deterioration in the credit position of firms by reducing the value of collateral that could result in an increase of borrowing costs and a credit contraction (Bruno & Shin, 2015a,b). This deterioration in domestic financial conditions dampens economic activity and ultimately increases investors’ perceived default risk. Moreover, Du & Schreger (2016) argue that a government is more inclined to default than to inflate when the currency mismatch of the corporate sector implies large adverse balance sheet effects from a currency depreciation. Using a new data set, they empirically confirm that higher external foreign currency debt of firms in EMEs is associated with higher sovereign credit risk. These related transmission channels are grouped under the term ‘financial channel’ of exchange

¹⁵A foreign exchange rate shock is best described as an exchange rate adjustment that occurs independent of monetary policy shocks; i.e. in response to foreign exchange market disturbances.

rates and can be considered as a potential offset to the traditional ‘net trade channel’.

This paper investigates the relative strength of the ‘net trade channel’ and the ‘financial channel’ by analyzing the effect of a currency depreciation on sovereign risk. We argue that the response of sovereign risk measures to the bilateral FX rate against the global currencies – primarily the USD – is expected to be dominated by the ‘financial channel’, when domestic borrowers hold a significant currency mismatch on their balance sheet. For a country that is short in a global currency, sovereign risk should increase with a depreciation shock of the FX rate. In contrast, the response of sovereign risk to a change in the NEER reflects the joint effect resulting from the ‘net trade’ and the ‘financial channel’ and, therefore, can be positive or negative depending on which channel dominates. Moreover, we assume that the reaction of sovereign risk measures to the change in the NEER that is orthogonal to the change in the bilateral FX rate ($NEER_{orth}$) should reflect, to a large extent, the impact of the ‘net trade channel’. Therefore, the response of sovereign risk to a depreciation shock to the $NEER_{orth}$ is expected to be negative. These considerations result in the following three hypotheses to be tested:

H₁: A depreciation shock to the FX rate towards a global currency against which an economy is short increases sovereign risk.

H₂: The response of sovereign risk to a depreciation shock of the NEER is ambiguous and is positively related to the degree of foreign currency exposure.

H₃: A depreciation shock to the $NEER_{orth}$ decreases sovereign risk.

The response of short-term interest rates to an exogenous increase in the exchange rate is ambiguous and might be zero or positive. According to the IMF’s Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER),¹⁶ the official monetary policy strategy of most EMEs is inflation targeting paired with a ‘de facto’ floating FX rate agreement. A major implication of inflation targeting is that price stability has priority over other goals, such as FX rate stabilization. In this case, monetary policy should react to FX rate movements only if they result in inflationary pressures. However, despite an inflation tar-

¹⁶See IMF (2016).

getting monetary strategy, Central Banks might also have an incentive to target FX rates. Garcia et al. (2011) use a DSGE model to examine if including FX rates next to inflation and output in the central bank’s policy rule can improve economic performance. They conclude that financially vulnerable economies are especially likely to benefit from FX rate smoothing. Respective empirical findings are documented by Ebeke & Azangue (2015). They find that particularly central banks of those EMEs are prone to FX interventions that show a high import dependence, and hold a large share of public and private assets and liabilities denominated in foreign currencies. In contrast, and consistent with typical commitments applying under inflation targeting, monetary authorities in countries with a more developed financial system and a credible financial openness policy intervene much less in FX markets. This points to a positive relationship between the size of the currency mismatch of a country’s external balance sheet and respective monetary policy responses to exogenous exchange rate shocks. Accounting for the described ambiguity, we expect a zero or positive response as stated in Table 1.

Shocks to sovereign risk

Noticing that a lack of confidence prompts investors to sell domestic assets, an exogenous increase in sovereign risk should lead to a depreciation of the domestic currency (i.e. to a positive impact effect as stated in Table 1). Only a few studies empirically analyze the effect of sovereign risk on exchange rates, generally confirming the positive relationship between these two variables. Della Corte et al. (2016) find that an increase in a country’s CDS spread is accompanied by a contemporaneous depreciation of its currency. Coudert & Mignon (2013) show for a sample of 18 EMEs that the FX rate premium increases with credit risk. Similarly, Keblowski & Welfe (2012) find that an increase in the Polish CDS exerts a significantly positive effect on the real exchange rate. This result is also confirmed by Ojeda-Joya & Sarmiento (2017) for five Latin American economies, namely Brazil, Chile, Colombia, Mexico, and Peru.

An exogenous increase in sovereign risk invokes a capital account deficit, which exerts depreciation pressures on the domestic exchange rate. Whether a central bank responds to such events and increases interest rates depends most likely (i) on the pass-through of FX

rate movements to inflation; and (ii) on the central bank's willingness to intervene in FX markets to smooth FX rate adjustments (see above). Hence, as stated explicitly in Table 1, the respective theoretical effect is either zero or positive.

Monetary policy shocks

The response of the exchange rate to an exogenous increase in interest rates is ambiguous, since one finds theoretical arguments for effects operating in both directions. From UIP arguments, one might expect a decrease of the FX rate, i.e. an appreciation of the domestic currency, in response to a contractionary monetary policy impulse. Several channels might be behind this positive effect: First, an interest rate increase signals the central bank's commitment to price stability, which reduces inflation expectations and prevents a vicious cycle of inflation and currency depreciation. Second, monetary tightening mutes economy activity by reducing the level of aggregate demand. This induces a reduction in imports, which improves the trade balance and investor confidence in the countries' prospects for external viability and debt repayment. Third, increasing domestic interest rates reduce FX rates (i.e. invokes a currency appreciation) by making domestic currency assets more attractive relative to foreign assets, thereby stimulating capital inflows.

Pointing to eventual state dependence of the overall effects of rising interest rates on the FX rate, there are also arguments for a positive effect of a monetary policy tightening on the FX rate, which become relevant during crisis periods. Specifically, higher interest rates may increase the debt service burden for firms, reduce investment incentives, and add to pressures on the banking system. This further raises prospects of financial collapse and external debt default, potentially undermining investor confidence. As a result, the currency depreciates. Summarizing the former considerations, Table 1 leaves the theoretical effect of exogenous interest rate shocks on exchange rates unspecified.

The ambiguity of the effect of a monetary policy shock on exchange rates also finds reflections in the empirical literature. Zettelmeyer (2004) finds for Australia, Canada and New Zealand in the 1990s and Dekle et al. (2001) for Korea during the Asian crisis that an increase in interest rates led to currency appreciations. Gould & Kamin (2000) and Kim & Ratti (2006) focus on a sample of emerging markets during the Asian financial crisis. While

Gould & Kamin (2000) do not find a remarkable effect of monetary policy and interest rates on exchange rates, Kim & Ratti (2006) document that national currencies depreciate in response to an increase in interest rates. Caporale et al. (2005) investigate the effects of an interest rate rise on the nominal exchange rate during tranquil and turbulent periods for a sample of four Asian countries. According to their findings, monetary policy helped to defend the FX rate during tranquil periods. However, during the Asian crisis monetary policies showed opposite effects.

The effect of a monetary policy shock on sovereign risk is also ambiguous. On the one hand, one might argue that the central bank signals with an increase in its policy rate its commitment to price stability. This fosters credibility of the government and the monetary authority with both effects reducing sovereign risk. On the other hand, if public debt is perceived to be at an unsustainable level, raising the domestic interest rate could increase the default probability of domestic borrowers, reduce capital inflows, and weaken the asset quality of domestic banks. Therefore, sovereign risk would increase in response to a contractionary monetary policy shock.

That the effect of an interest rate shock on sovereign risk is ambiguous is supported by mixed results in the empirical literature. For instance, Andrade & Teles (2006) estimate a negative relationship between interest rates and sovereign risk for Brazil, while Blanchard (2004) and Favero & Giavazzi (2004) find a positive effect of a contractionary monetary policy shock on sovereign risk, given that financial markets believe that the economy has moved from a regime of ‘monetary dominance’ to one of ‘fiscal dominance’.¹⁷

Table 1 summarizes the impact effects of shocks that we consider to be plausible on basis of theoretical considerations and empirical evidence.

¹⁷‘Fiscal dominance’ is an economic condition in which a central bank’s objective for price stability might come into conflict with a dominant fiscal policy stance (Sargent & Wallace, 1981; Woodford, 2001).

Variable	Shocks		
	FX/NEER	Sovereign risk	Interest rates
FX/NEER	+	+	?
Sovereign risk	+/-/?	+	?
interest rate	≥ 0	≥ 0	+

Table 1: Theoretical sign pattern of shocks in FX and NEER system. Both SVARs share information on interest rates and CDS. The notation ‘+/-/?’ indicates the main focus of our analysis, as stated in terms of the hypotheses 1, 2 and 3.

A note on triangularity and cross sectional heterogeneity

In light of the schematic display of expected contemporaneous effects within the triad of exchange rates, sovereign risks and interest rates in Table 1, two observations are worth mentioning. First, consistent with the fact that price formation in financial markets tends to occur immediately after the arrival of news, theoretical and empirical evidence points to a bi-directional and contemporaneous relationship between exchange rate changes and sovereign risk. In contrast, by imposing a triangular structure, Hofmann et al. (2016) assume that exchange rates do not respond to the effects of shocks that enter the system through sovereign risk or interest rates. Their identification method must therefore be regarded as critical. Hofmann et al. (2020) address the endogeneity issue by means of external information. Without an attempt to adjust for predictable information, the authors construct exchange rate *shocks* that are equal to *observed* log FX rate changes on days with monetary policy messages from the European Central Bank and zero otherwise. While such an instrument likely processes exogenous information, Hofmann et al. (2020) do not discuss the critical issue of instrument relevance.

Second, from the schematic display in Table 1 it is most likely to expect that even if in a cross sectional perspective structural model parameters are identical in terms of their signs, these parameters are likely country specific in terms of magnitudes. Hofmann et al. (2020) do not account for possible country-specific heterogeneity of model parameters. Instead, they estimate linear projection panel regressions with fixed effects to derive the dynamic impulse responses of sovereign risk measures to an exchange rate shock.

Our novel identification scheme applied in this paper has the major advantage of not imposing restrictions on the estimated (country-specific) structural parameters. Thus, we

account for both the bi-directional relationship between exchange rates and sovereign risk and the potential cross-sectional heterogeneity in the model parameters. Accordingly, we can formally test, if country specific estimates of the structural parameter matrix are in line with the theoretical patterns in Table 1. If so, this would provide strong support for the economic labelling of the statistically identified shocks. Moreover, unlike, the pooled fixed effect panel regressions in Hofmann et al. (2020), the MG estimator has a well specified population counterpart (i.e. the cross sectional mean of the true parameters), and MG inference has been shown to consistently respect nominal significance levels under an increasing cross sectional data dimension (see also Rebucci, 2010).

4 Cross sectional subsampling for testing H_2

With structural estimates of the impact of exchange rate shocks on sovereign risk at hand, we can test our hypothesis H_2 from Section 3, claiming that the response of sovereign risk to a depreciation shock of the NEER is positively related to the degree of foreign currency exposure. Assessing country risks posed by foreign currency exposures requires looking beyond aggregate numbers, since they may conceal sectoral differences. As shown by Chui et al. (2016) and McCauley et al. (2015), the public sector’s net foreign currency exposure has continuously declined in most EMEs since the Great Financial Crisis of 2008 due to higher official currency reserves and less foreign currency denominated government debt. In contrast, the private sector still borrows in global currencies, such as the USD, to a large extent. Arguments and evidence of Bruno & Shin (2015a) and Du & Schreger (2016) point to a prime role of the currency mismatch of the non-public sector for sovereign risk. They show that in countries where the private sector largely relies on foreign currency denominated debt, governments are relatively more inclined to explicitly default than to inflate away the debt because of adverse effects of a depreciation on the private sector. Thus, irrespective of the government’s net foreign currency position, a sovereign credit risk remains if there is a significant private sector currency mismatch. This suggests that looking at foreign currency exposures at the country level could be very misleading and that the response of sovereign risk measures to a depreciation shock of the NEER is predominantly positively related to the

size of the private sector’s net foreign currency position. To test hypothesis H_2 , we therefore focus on the net foreign currency exposure of the non-public sector. However, as a robustness check, we also look at the aggregate level of external currency position, including the government sector.

To our knowledge, information on foreign currency asset and liability holdings split up by sectors and currencies is not provided in publicly available databases. Official balance of payments and international investment position data usually do not record the currency composition of foreign currency assets and liabilities. Moreover, these statistics only provide information on country’s external foreign currency debt and ignore internal foreign currency denominated liabilities, i.e., bank and bond financing in foreign currencies from one resident to another in the same country. Other databases, like e.g. statistics provided by the Bank of International Settlement, provide detailed information on the currency composition of non-government sectors’ liabilities. However, to calculate the private sector’s net currency exposure, one also needs information on the asset side, e.g., how many assets households and corporations hold in a specific currency. Such information, however, is missing in the database of the Bank of International Settlement.

In coping with these difficulties, we use a data set provided by Chui et al. (2016). They combine various data sources to construct a variable that measures the net foreign currency assets of the non-government sector as a percentage of exports. Total exports of goods and services are used as a reference in their calculation to take account of the argument that countries with high export to GDP ratios can sustain larger foreign currency mismatches. Unfortunately, information on Croatia is missing from their dataset. To fill this gap, we use for this country the net foreign currency exposure information provided by Bénétrix et al. (2015).¹⁸ A caveat of the data provided by Chui et al. (2016) is that it is available only between 2006 and 2014, while our structural estimates are determined from sample information between 2004 and 2016. However, plotting the variables over time reveals that volatility is rather low. For this reason, we take a quite ‘conservative’ approach to country grouping. We group the economies whose non-government average net foreign currency position is above the sample median in one group (Group 1) and the remaining countries in

¹⁸Unlike Chui et al. (2016), Bénétrix et al. (2015) ignores local debt denominated in foreign currencies.

another group (Group 2).

Insert Figure 1 around here

Figure 1 shows the average net foreign currency exposure of the non-government sector relative to exports of the two country groups. The non-government sector of all sample countries but South Africa shows, on average, a negative net foreign currency position, meaning that they hold less foreign currency assets than liabilities. In this case, a depreciation of the domestic currency induces negative valuation effects. For countries in Group 2 (i.e. Brazil, Chile, Mexico, Korea, Hungary, Poland Turkey, and Croatia), the currency mismatch is larger in magnitude than the sample median. We therefore expect that the ‘financial channel’ in this subsample is stronger than for the countries in Group 1, which show more balanced foreign currency exposure (i.e., Colombia, Indonesia, Malaysia, Philippines, Thailand, Czech Republic, Russia and South Africa). As expressed in hypothesis H₂, therefore, in economies that belong to Group 2 sovereign risk likely respond more sensibly to a depreciation shock of the NEER in comparison with Group 1 economies.

5 Estimation results

5.1 Reduced form analysis and structural identification

For a given economy ($i = 1, 2, \dots, 16$), lag order selection by means of the BIC obtains mostly identical model orders for VARs comprising FX ($m = 1$), NEER ($m = 2$) or NEER_{orth} ($m = 3$) data. With a few exceptions (i.e. at most 4 out of 16 economies), the chosen lag order is one.¹⁹ VARs for Korea are specified with the largest autoregressive order of three.

Targeting independent structural shocks delivers unique estimates in non-Gaussian dynamic systems. We refrain from providing detailed results on normality tests for reduced form VAR residuals and their single components for each country in the sample. Overall, at the system level, Jarque-Bera (JB) statistics indicate non-Gaussianity of reduced form VAR

¹⁹At the MG level core modelling outcomes are largely robust, if the selection of single economy VAR orders relies on the AIC.

residuals for each country with strong significance (p -values below 0.1%).²⁰ With regard to respective diagnostics for single components of estimated VAR residuals, it turns out that in some rare cases one out of three reduced form VAR residual processes shows JB diagnostics that lack significance. From diagnosing strong evidence against normality for both marginal and joint residual processes, we conclude that the detection of shocks with weakest dependence can be suitably adopted for statistical identification in our panels of SVAR models.

Insert Table 2 around here

As outlined in Section 2, our structural identification is based on maximizing the effects of structural shocks on those variables to which they are primarily associated. MG results for the country specific structural parameter matrices are displayed in Table 2. The table shows the means and their corresponding t -ratios as well as descriptive statistics (minimum, maximum and three quartiles) for the impact effects of the structural shocks. MG coefficient estimates for all three groups of SVARs indicate that, on average, the (absolute) largest effects by row can be found in the diagonal positions of the structural matrix estimates. Hence, the statistical identification criteria highlight three shocks that exert their strongest impact on the exchange rates, CDS spreads, and short-term interest rates. In terms of an economic labelling one might refer to these shocks as exchange rate shock (1st column), risk shock (2nd column), and interest rate or domestic monetary policy shock (3rd column). Apart from MG results, it is worth emphasizing that the documented characteristic quantiles are also supportive of such a labelling for the vast majority of economy specific results. For instance, the 25% quartiles of diagonal estimates in a given row exceed (with one exception, i.e., \hat{d}_{33} for $m = 2$), the absolute value of both the first and the third quartile of the remaining row specific off diagonal estimates.

²⁰Throughout, if not stated otherwise, our discussion of effect significance refers to a 5% nominal level.

5.2 Impact and dynamic effects of exchange rate shocks

Quantile estimates in Table 2 indicate a marked cross sectional heterogeneity for most structural parameters. In the systems including FX rates and the NEER ($m = 1, 2$), the only exceptions are the contemporaneous impacts of exchange rate shocks on sovereign risk (elements \hat{d}_{21}) and of risk shocks on exchange rates (elements \hat{d}_{12}), which are estimated with common signs at the displayed 25% and 75% quantiles, throughout. Hence, for these effects one might consider the identified models to hint at important and (almost) homogeneous transmission channels. As we have noted before, seeing that structural parameter estimates \hat{d}_{12} are (almost) homogeneously positive provides a strong economic underpinning of the statistically identified shocks. By implication, imposing a triangular model structure would be significantly at odds with the data at the MG level.

The estimated contemporaneous effects of all three shocks, i.e. an exchange rate shock, a risk shock and an interest rate shock on the model variables are in line with our theoretical considerations described in Section 3 and summarized in Table 1. For the sake of brevity, we only concentrate on a detailed description of MG estimates of d_{21} , which quantify the on-impact response of our sovereign risk measure to an exchange rate shock. These effects are of core interest for investigating the hypotheses H_1 to H_3 raised in Section 3.

Insert Figure 2 around here

MG estimates for the SVARs containing the USD FX rate are summarized in the first panel of Table 2. We find that a depreciation shock to the FX rate of one standard deviation results in an instantaneous upward adjustment of CDS spread by 5.55 basis points.²¹ Figure 2 shows the corresponding cumulated impulse response functions (IRFs). The dynamic effect characteristics suggest that the positive impact of an exchange rate shock on country risk is highly persistent over time, reaching about 10 basis points after about one month, which is

²¹When interpreting the estimate of the structural parameter d_{21} as an impulse response, it is important to notice that the MG estimate averages across non-linear effects and, hence, provides an approximate quantification. Regarding the shock size, we notice that the respective MG variance estimate is 6.35E-04. From this, one may consider $0.025 = \sqrt{6.35E-04}$ as a reasonable approximation of the magnitude of a shock of size one standard deviation.

comparable in magnitude to the results of Hofmann et al. (2020).²² Hence, our estimation results confirm hypothesis H_1 . Sovereign risk of EMEs increase when the domestic currency loses value against the USD. This suggests that the response of sovereign risk of EMEs to a depreciation of the domestic currency against the USD is dominated by the ‘financial channel’.

Insert Figure 3 around here

Depreciation shocks to the NEER also increase sovereign risk, albeit the effect is somewhat smaller in magnitude. On average, a shock to the NEER of one standard deviation results in an instantaneous increase of the CDS spread by 2.51 basis points.²³ However, while the described on-impact effect is highly significant in the FX system, it lacks significance at conventional levels in the NEER system. Contrary to the result of Hofmann et al. (2020), the respective IRF shown in Figure 3 highlights that the accumulated effect of a NEER shock is insignificant at all horizons. These results confirm hypothesis H_2 that the response of sovereign risk to a change in the NEER reflects the joint effect resulting from the ‘net trade’ and the ‘financial channel’ and, therefore, can be positive or negative depending on which channel dominates. Both the heterogeneity and the overall insignificance of the estimated contemporaneous effects of exchange rate shocks on sovereign risk obtain from noticing that the relative importance of the two transmission channels is likely country specific.

Insert Table 3 around here

As formalized further in hypothesis H_2 , we assume that the ‘financial channel’ becomes more dominant when an economy’s net foreign currency exposure is large. Thus, the response of sovereign risk to a depreciation shock of the NEER should be positively related to the degree of foreign currency exposures. To test H_2 , we divide the sample economies on the basis of data on net foreign currency exposures relative to exports into two groups (see also

²²Identified IRFs show the cumulated response of endogenous variables to a structural shock of size one standard deviation. After the determination of the structural matrices D the extraction of identified IRFs is straightforward (Lütkepohl, 2007, Chapter 2)

²³The residual variance MG estimator for the NEER equation is 5.88E-04, resulting in an approximate size of a representative shock being 0.024.

the discussion in Section 4). The upper (lower) panel of Table 3 shows the MG results for countries whose private sector’s foreign currency exposure is on average below (above) the sample median. Subsample structural estimates are broadly in line with H_2 . When focusing only on countries with a large negative net foreign currency exposure (Group 2), we diagnose a significantly positive on-impact effect of a depreciation shock of the NEER on sovereign risk (4.96 basis points at the subsample MG level). Thus, the ‘financial channel’ is clearly dominant for these countries. However, when focusing only on countries whose private sectors’ balance sheet shows a relatively small currency mismatch, the response of sovereign risk to a depreciation shock is also positive (0.05 basis points at the subsample MG level), albeit close to zero in magnitude and insignificant at conventional levels. Thus, the transmission of exchange rate shocks via the ‘financial channel’ is weaker in this group of economies. The findings in favour of H_3 are also confirmed by the IRFs shown in Figure 4. As far as we know, we are the first to show that the sign and size of the response of sovereign risk measures to an exchange rate shock are moderated by the size of the net foreign currency exposure of a country.²⁴

Insert Figure 4 around here

To disentangle the effects resulting from the ‘net trade channel’ and the ‘financial channel’, we estimate the SVARs comprising NEER changes that are considered unrelated to changes of (log) USD rates ($NEER_{orth}$, $m = 3$).²⁵ According to H_3 , one would expect that a depreciation shock to the $NEER_{orth}$ decreases sovereign risk, since it filters out the ‘financial channel’ of exchange rate transmission such that only the effect of the ‘net trade channel’ should be left. Respective MG estimation results are summarized in the third

²⁴As a robustness test, we have grouped our sample of countries alternatively according their aggregate level of net USD exposure, thus, taking not only the foreign currency position of the private sector into account, but also of the government. In line with the arguments and evidence of Bruno & Shin (2015a) and Du & Schreger (2016), we find no significant difference in the impulse responses, when looking at the total net USD position. Thus, this confirms the prime role of the currency mismatch of the non-public sector for the strength of the ‘financial channel’.

²⁵Alternatively, we re-calculated the nominal effective exchange rate for all countries contained in our sample excluding the USD. However, this newly calculated nominal exchange rate was still highly correlated with the overall NEER and the bilateral USD rate, indicating the strong correlation of all other exchange rates with the USD. Therefore, this approach turned out is not effective for isolating the ‘net trade channel’ from the ‘financial channel’.

panel of Table 2 and the corresponding IRFs are shown in Figure 5. In line with H_3 , the contemporaneous and dynamic response of CDS spreads to a depreciation shock is negative on average. A one standard deviation increase in the $NEER_{orth}$ induces the sovereign risk premium to decrease by 0.60 basis points.²⁶ However, as this effect lacks significance at conventional levels, we can conclude that the change in the NEER has no significant effect once adjusted for changes of the (log) USD rate. From these results, we conclude that it is primarily the USD exchange rate that matters and that the ‘financial channel’ is more important in the transmission of exchange rate shocks to sovereign risk in comparison with the traditional ‘net trade channel’. A potential explanation why the latter seems to play only a minor role in the transmission of exchange rate shocks is that for some countries in our sample a sizeable fraction of exports is denominated in USD. A deterioration in a country’s terms of trade in response to a depreciation of the domestic currency against the USD is therefore dampened.

Insert Figure 5 around here

5.3 Recursive structural relations (evidence and implications)

Our estimation results displayed in Table 2 show that for all structural models (with slightly weaker evidence for SVARs including $NEER_{orth}$), estimates located above the main diagonal of the structural parameter matrix are not uniquely insignificant at the MG level. Accordingly, although country specific results show sizeable variability throughout (see e.g. the often detected sign changes between the first and third quartile of non-diagonal estimates), MG estimates and diagnostics indicate non triangularity.²⁷ Hence, one might expect that a more restrictive identification via *a-priori* triangular covariance decompositions – as applied by Hofmann et al. (2016) – might come at the risk of biased structural estimates at both the aggregate and the single country level. More specifically, using lower triangular Cholesky factors for structural analysis might result in overestimating the relevance of those trans-

²⁶The residual variance MG estimator for the $NEER_{orth}$ equation is 1.18E-04 obtaining 0.011 as an approximate size of a representative shock.

²⁷We regard common signs of quantile estimates as informal hints at similarities among single economy estimates. By construction, the diagonal parameter estimates are positive within each structural matrix.

mission channels that remain active within the recursive model. Among these, one finds, in particular, the effects of exchange rate shocks on domestic CDS spreads, which are of interest when it comes to an inferential assessment of the hypotheses H_1 to H_3 .

To highlight the potential of effect overestimation, we impose recursive transmission patterns to each panel member. Respective results for SVARs (MG structural estimates are multiplied with 100, t -ratios in parentheses) read as

$$D_1^\Delta = \begin{bmatrix} 2.432 & - & - \\ (14.2) & & \\ 12.12 & 21.06 & - \\ (8.88) & (10.2) & \\ 5.799 & 4.294 & 41.73 \\ (1.69) & (2.39) & (4.17) \end{bmatrix}, D_2^\Delta = \begin{bmatrix} 2.213 & - & - \\ (8.63) & & \\ 10.55 & 21.86 & - \\ (7.013) & (10.0) & \\ 5.722 & 5.480 & 40.39 \\ (2.03) & (2.82) & (4.39) \end{bmatrix}$$

$$\text{and } D_3^\Delta = \begin{bmatrix} 1.006 & - & - \\ (9.34) & & \\ 1.012 & 24.82 & - \\ (0.92) & (11.6) & \\ 1.983 & 7.029 & 42.82 \\ (1.67) & (2.46) & (4.17) \end{bmatrix} \quad (5)$$

Comparing the results in (5) with those in Table 2 shows that the closure of impact transmissions from risk and interest rate shocks on exchange rates (i.e. the imposition of triangularity) comes with markedly stronger assessments of the effects of exchange rate shocks on both sovereign risk and short term interest rates in comparison with unrestricted structural estimates. In SVARs of nominal FX rates, $m = 1$, (NEER, $m = 2$) the unrestrictedly estimated impact effects of 5.55 (2.51) basis points shift up to 12.1 (10.6) basis points as implied by the triangular model.

6 Conclusions and policy implications

We explore the effects of an exchange rate shock on sovereign risk using monthly data covering the 10/2004 to 12/2016 period in a panel of 16 major emerging market economies (EMEs) by means of structural vector autoregressive models (SVARs). We differentiate between identified shocks hitting the bilateral exchange rate against the USD (FX rate), the nominal effective exchange rate (NEER), and the NEER that is unrelated to changes of USD

FX rates ($NEER_{orth}$). Regarding the transmission of exchange rate adjustments to the real economy and sovereign risk, this differentiation allows us to discriminate between the traditional ‘net trade channel’ and the ‘financial channel’. Instead of using ad-hoc Cholesky decompositions, as in Hofmann et al. (2016), we apply a novel identification approach of the structural shocks (Herwartz, 2018) that exploits the uniqueness of independent structural shocks in non-Gaussian SVARs. As a particular merit, this identification strategy accounts for the complex interrelations within the triad of exchange rates, sovereign risks, and interest rates in an unrestricted and purely data driven manner.

We find that FX rate movements are an important determinant of sovereign risk. However, the direction and size of the response of sovereign risk depend on the type of the exchange rate measure that we look at and on the size of an economy’s net foreign currency exposure. In general, a depreciation of the domestic currency against the USD increases sovereign risk. In contrast, a depreciation of the effective exchange rate turns out to have only a significant effect on sovereign risk for countries with large negative net foreign currency exposures. In this case, a depreciation of the NEER also induces an increase in sovereign risk. When looking at the NEER that is unrelated to changes of the USD FX rate, we find in general no significant effect of a depreciation shock on sovereign risk. Thus, we conclude from these results that it is primarily the USD exchange rate that matters and that the ‘financial channel’ is more important in the transmission of exchange rate shocks to sovereign risk in comparison with the traditional ‘net trade channel’. Moreover, we confirm the prime role of the currency mismatch of the non-public sector for the strength of the ‘financial channel’. Hence, only for countries, whose non-government sector is net short in USD, a depreciation of their domestic currency induces a tightening of their financial conditions.

Our results suggest that valuation effects due to exchange rate movements play an important role for monetary policy spillovers from global economies, like the US, to EMEs. As an example, consider the low interest rate environment in the US in the aftermath of the Great Financial Crisis (GFC), which created currency appreciation pressures in EMEs. According to our estimation results, domestic currency appreciations compress sovereign risk and, hence, lower sovereign bond yields. Thus, even without any monetary policy coordination, cross country correlation of interest rates would increase due to changes in sovereign risk premia.

To counteract the eventual inflationary pressures, the Central Bank of an affected EME might consider raising policy rates. However, this would aggravate domestic currency appreciation further and a vicious circle would start. Alternatively, the Central Bank of an affected EME could respond to appreciation pressures by also lowering its policy rate, which would come at the cost of higher inflation. Thus, monetary policy autonomy in countries with large foreign currency exposures of the private sector is constrained independent of the choice of its exchange rate regime.

To regain more monetary autonomy in the process of globalization, it appears effective to establish policies that support a further reduction of a country's net foreign currency exposure. As shown by Chui et al. (2016) and McCauley et al. (2015), public sector net foreign currency exposure has already continuously declined since the GFC due to higher currency reserves and reduced foreign currency denominated government debt, while the private sector in most EMEs still borrows to a large extent in foreign, mostly global, currencies such as the USD. Hence, advisable policies might target at an improvement of a country's international credit standing in general, and promote the development of domestic bond and equity markets as an alternative source of financing for firms.

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Appendix - Estimation of structural model parameters

Building upon the result of Comon (1994), recent approaches to data-based SVAR identification (i.e the point estimation of D in (2)) follow a variety of avenues to extract independent components from the data (e.g., Moneta et al., 2013; Matteson & Tsay, 2017; Lanne et al., 2017; Gouriéroux et al., 2017, . See also Kilian & Lütkepohl (2017) for a textbook treatment).²⁸ In this study, we adopt an identification scheme suggested in Herwartz (2018). As a particular distinction between the identification schemes in Matteson & Tsay (2017) and Herwartz (2018), the loss function employed in the latter is scale free and, hence, is supposed to be (more) robust in light of data and model heterogeneities that characterize our panel. Detecting independent (or least dependent) shocks in a nonparametric manner requires two essential building blocks. First, one needs a well parameterized space of reduced form covariance ($\hat{\Omega}$) decompositions, i.e. a collection of potential structural matrices. For this purpose, we consider a decomposition space that obtains from systematic rotations of a baseline decomposition, $\Omega = QQ'$ (with Q being a lower triangular Cholesky factor). To be explicit, let $\tilde{D}_\theta = QR_\theta$ denote a candidate decomposition matrix with the rotation matrix R_θ ($R_\theta \neq I_K$, $R_\theta R'_\theta = I_K$) being the product of distinct forms of Givens rotation matrices. For instance, in our case of $K = 3$ R_θ reads as

$$R_\theta = \begin{bmatrix} 1 & 0 & 0 \\ 0 & \cos(\theta_1) & -\sin(\theta_1) \\ 0 & \sin(\theta_1) & \cos(\theta_1) \end{bmatrix} \times \begin{bmatrix} \cos(\theta_2) & 0 & -\sin(\theta_2) \\ 0 & 1 & 0 \\ \sin(\theta_2) & 0 & \cos(\theta_2) \end{bmatrix} \times \begin{bmatrix} \cos(\theta_3) & -\sin(\theta_3) & 0 \\ \sin(\theta_3) & \cos(\theta_3) & 0 \\ 0 & 0 & 1 \end{bmatrix},$$

with $0 \leq \theta_i < \pi$ and $\theta = (\theta_1, \theta_2, \theta_3)'$. The parameterized space of potential structural matrices allows the extraction of corresponding structural shocks, $\tilde{\xi}_t = \tilde{D}_\theta^{-1} \hat{u}_t$. Second, to detect the particular member of this space that consists of unique least dependent (or independent) shocks, model identification requires a statistical measure of joint (in)dependence. In light of cross sectional and within country data heterogeneity, the detection of independent compo-

²⁸At first glance, the explicit assumption of independent structural shocks appears as a restrictive 'add-on' to the common assumption of orthogonality. In this context, however, it is worth noticing that an independence (or Gaussian) assumption is also implicit in the common approach to determine IRFs from notions of isolated unit impulse hitting a system of interest. Characterizing isolated unit shocks, the restriction $E[\varepsilon_{jt}|\varepsilon_{it} = 1] = 0$, $i \neq j$, only applies under independence of the elements in ε_t or in a Gaussian context.

nents in this work follows Herwartz (2018) and relies on the minimization of the rank based Cramér-von-Mises (CvM) distance (Genest et al., 2007),

$$\mathcal{B} = \int_{(0,1)^K} \left[\sqrt{T} \left(C(\tilde{\xi}) - \prod_{k=1}^K U(\tilde{\xi}_k) \right) \right]^2 d\tilde{\xi}. \quad (6)$$

In (6), random vectors $\tilde{\xi}$ are obtained from orthogonalized reduced form model disturbances ($\tilde{\xi}_t = \tilde{D}_\theta^{-1} \hat{u}_t$), C and U denote the empirical copula of orthogonalized model disturbances and the implied copula under independence, respectively.²⁹ Conditional on the dependence criterion in (6), an ‘estimator’ of D (or θ) solves the minimization problem

$$\hat{D} = \tilde{D}_{\hat{\theta}}, \text{ with } \hat{\theta} = \operatorname{argmin}_{\theta} \{ \mathcal{B} | \tilde{\xi}_t = \tilde{D}_\theta^{-1} u_t \}. \quad (7)$$

The solution of (7) requires nonlinear optimization.³⁰ The estimated covariance decomposition obtains structural shocks $\hat{\xi}_t = \hat{D}^{-1} u_t$ under the assumptions in (4) which are unique except for the column ordering in \hat{D} . From the six alternative orderings of the columns in \hat{D} we select the particular candidate with maximum sum of diagonal entries to ensure that a given shock exerts the strongest effect on the variable to which it is primarily associated (Lütkepohl & Nets̄unajev, 2014).

²⁹As a particular merit of the CvM distance in detecting independent components, it is worth pointing out that this statistic is scale free and does not rely on the marginal distributions of the structural shocks. Genest et al. (2007) consider the CvM distance as an ‘ideal’ choice for nonparametric dependence diagnosis unless the analyst is sufficiently confident to opt for a particular local dependence alternative. Analyzing a set of heterogeneous EMEs ($i = 1, 2, \dots, 16$) and VAR models ($m = 1, 2$ and $m = 1, 3$), we refrain from suggesting any particular local alternative to be equally suitable for all models and economies. For an explicit representation of the CvM distance \mathcal{B} the reader should consult Section 4 of Genest et al. (2007).

³⁰We use a standard optimization routine based on simulated annealing as implemented in R. Similar procedures are implemented in the R package ‘svars’ (<https://cran.r-project.org/package=svars>) as provided by Lange et al. (2017).

Tables

	min	$q_{.25}$	$q_{.50}$	$q_{.75}$	max	mean	t -rat
	SVARs including bilateral exchange rates ($m = 1$)						
\hat{d}_{11}	1.21	1.73	2.04	2.57	3.12	2.12	14.20
\hat{d}_{21}	-10.11	2.84	5.26	8.46	14.72	5.55	3.45
\hat{d}_{31}	-93.04	-1.64	0.84	3.81	7.82	-4.60	-0.77
\hat{d}_{12}	-0.14	0.14	0.70	1.24	2.37	0.74	4.13
\hat{d}_{22}	12.88	16.40	22.36	25.52	38.39	22.54	12.20
\hat{d}_{32}	-42.80	-4.31	0.80	5.26	36.67	-0.03	-0.01
\hat{d}_{13}	-0.67	-0.14	0.05	0.57	1.80	0.30	1.77
\hat{d}_{23}	-11.46	-1.86	2.71	4.21	18.15	2.22	1.22
\hat{d}_{33}	6.12	11.28	25.04	55.25	111.84	39.98	4.62
	SVARs including NEER ($m = 2$)						
\hat{d}_{11}	0.58	1.25	1.62	2.26	3.68	1.86	8.74
\hat{d}_{21}	-24.13	0.56	3.84	7.50	15.76	2.51	1.07
\hat{d}_{31}	-47.73	-0.84	0.97	4.73	19.08	0.35	0.10
\hat{d}_{12}	-0.89	0.34	0.62	0.92	2.93	0.65	3.18
\hat{d}_{22}	12.81	15.57	20.57	24.31	42.48	21.26	11.35
\hat{d}_{32}	-18.06	-2.45	1.92	11.87	22.38	2.56	0.90
\hat{d}_{13}	-0.54	-0.17	-0.02	0.46	2.47	0.27	1.34
\hat{d}_{23}	-18.33	-6.57	0.33	5.87	19.57	1.02	0.39
\hat{d}_{33}	6.13	11.53	21.88	53.58	130.93	39.61	4.25
	SVARs including NEER _{orth} ($m = 3$)						
\hat{d}_{11}	0.42	0.65	0.77	1.06	1.92	0.94	8.82
\hat{d}_{21}	-9.40	-4.04	-0.19	2.12	9.68	-0.60	-0.43
\hat{d}_{31}	-95.16	-1.73	0.30	6.00	15.71	-4.00	-0.63
\hat{d}_{12}	-0.30	-0.03	0.06	0.25	0.62	0.12	1.94
\hat{d}_{22}	13.38	13.82	22.75	25.10	43.13	22.73	10.31
\hat{d}_{32}	-41.51	-2.86	2.38	5.90	23.07	-1.01	-0.26
\hat{d}_{13}	-0.47	-0.20	-0.04	0.09	0.59	-0.00	-0.01
\hat{d}_{23}	-9.71	-1.29	1.41	2.16	26.87	2.61	1.14
\hat{d}_{33}	5.81	10.20	25.75	49.95	115.24	39.54	4.42

Table 2: Descriptive statistics (min, max, and quantiles q_α at levels $\alpha = 0.25, 0.50, 0.75$) and mean group results (mean and t -ratio) for typical elements of structural matrix estimates (\hat{d}_{ij}) obtained from three alternative SVAR model compositions. All structural estimates have been multiplied with 100.

	min	$q_{.25}$	$q_{.50}$	$q_{.75}$	max	mean	t -rat
Group 1							
\hat{d}_{11}	1.01	1.16	1.30	2.28	3.03	1.74	5.97
\hat{d}_{21}	-24.13	-3.22	1.97	5.37	15.76	0.05	0.01
\hat{d}_{31}	-47.73	-1.75	0.06	2.70	12.07	-4.15	-0.64
\hat{d}_{12}	-0.89	0.23	0.51	0.95	2.93	0.67	1.69
\hat{d}_{22}	13.17	15.02	17.21	24.73	42.48	21.28	6.06
\hat{d}_{32}	-9.90	-0.00	1.92	11.87	12.54	3.62	1.27
\hat{d}_{13}	-0.54	-0.16	0.08	0.50	2.47	0.40	1.10
\hat{d}_{23}	-9.74	-6.90	-3.47	3.67	14.69	-0.57	-0.18
\hat{d}_{33}	6.13	9.44	18.30	42.84	72.79	28.03	3.31
Group 2							
\hat{d}_{11}	0.58	1.59	1.81	2.26	3.68	1.98	6.10
\hat{d}_{21}	-6.38	3.77	6.25	7.50	11.21	4.96	2.69
\hat{d}_{31}	-5.14	-0.17	2.70	7.20	19.08	4.84	1.64
\hat{d}_{12}	-0.13	0.46	0.68	0.92	1.03	0.62	4.44
\hat{d}_{22}	12.81	19.68	21.45	24.20	27.28	21.24	12.87
\hat{d}_{32}	-18.06	-6.29	2.15	9.86	22.38	1.50	0.30
\hat{d}_{13}	-0.49	-0.17	-0.09	0.46	1.12	0.14	0.72
\hat{d}_{23}	-18.33	-2.76	1.43	7.85	19.57	2.60	0.61
\hat{d}_{33}	10.03	17.14	33.05	76.38	130.93	51.18	3.16

Table 3: Descriptive statistics for structural estimates from SVARs comprising NEER ($m = 2$) for two groups of economies separated according to the currency mismatch of the non-government sector relative to exports (Chui et al., 2016). The cross sectional separation is displayed in Figure 1. Group 1 (Group2) consists of Colombia, Indonesia, Malaysia, Phillipines, Thailand, Czech Republic, Russia and South Africa (Brazil, Chile, Mexiko, Korea, Hungary, Poland, Turkey and Croatia). For further notes see Table 2.

Figures

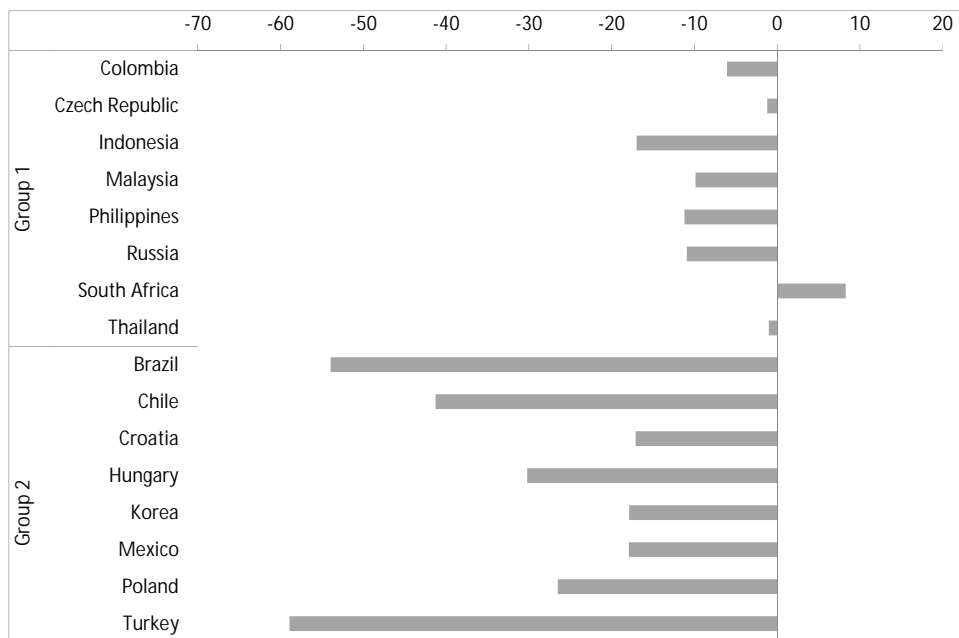


Figure 1: Currency mismatch of the non-government sector relative to exports. Sources: Chui et al. (2016), Bénétrix et al. (2015) and own calculations. This is also the subdivision behind results documented in Table 3.

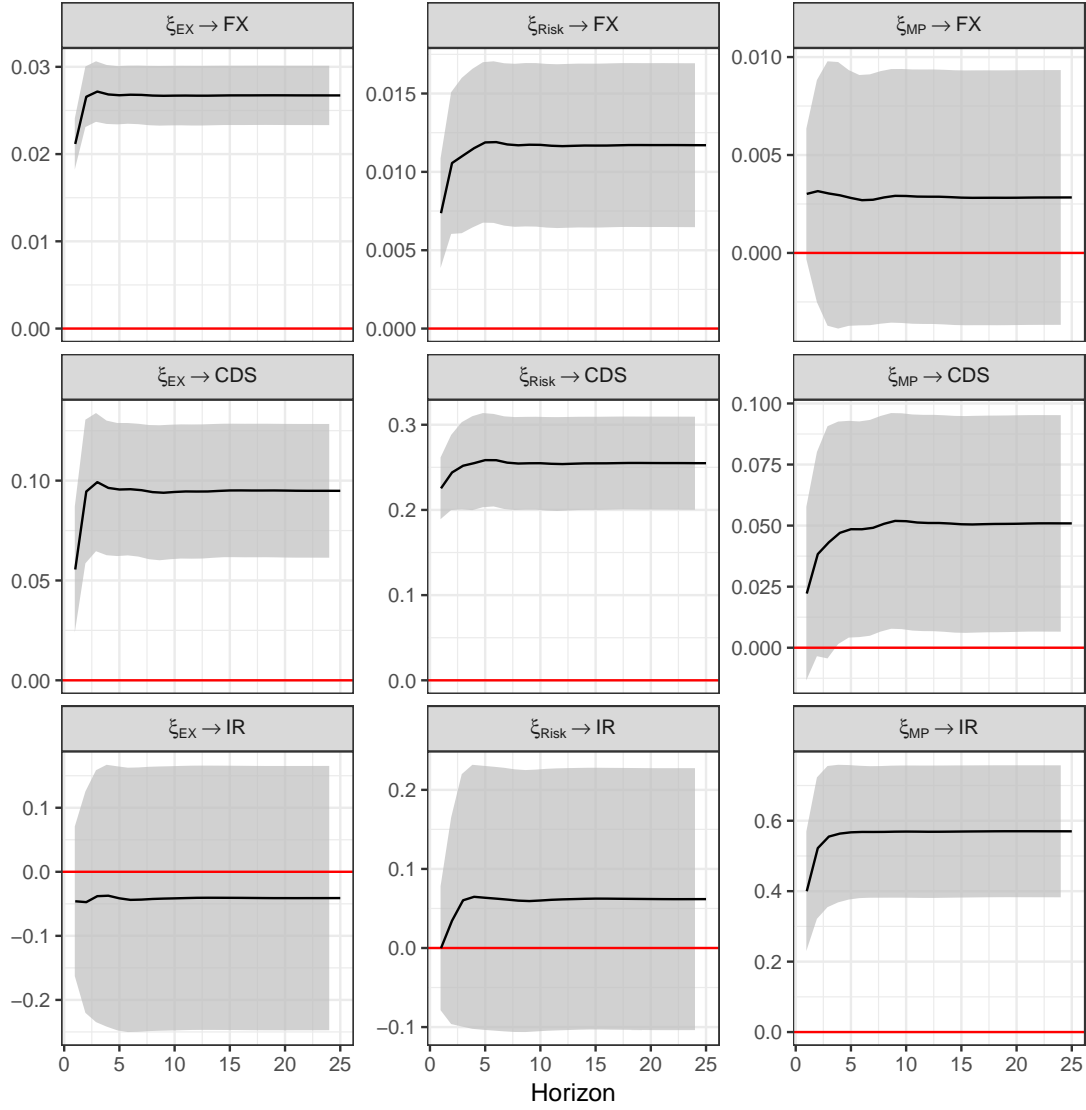


Figure 2: MG cumulated IRFs in FX systems ($m = 1$). Exchange rate, country risk and monetary policy shocks are indicated with ‘EX’ ‘Risk’ and ‘MP’, respectively. The shaded areas are pointwise MG confidence intervals with 95% coverage.

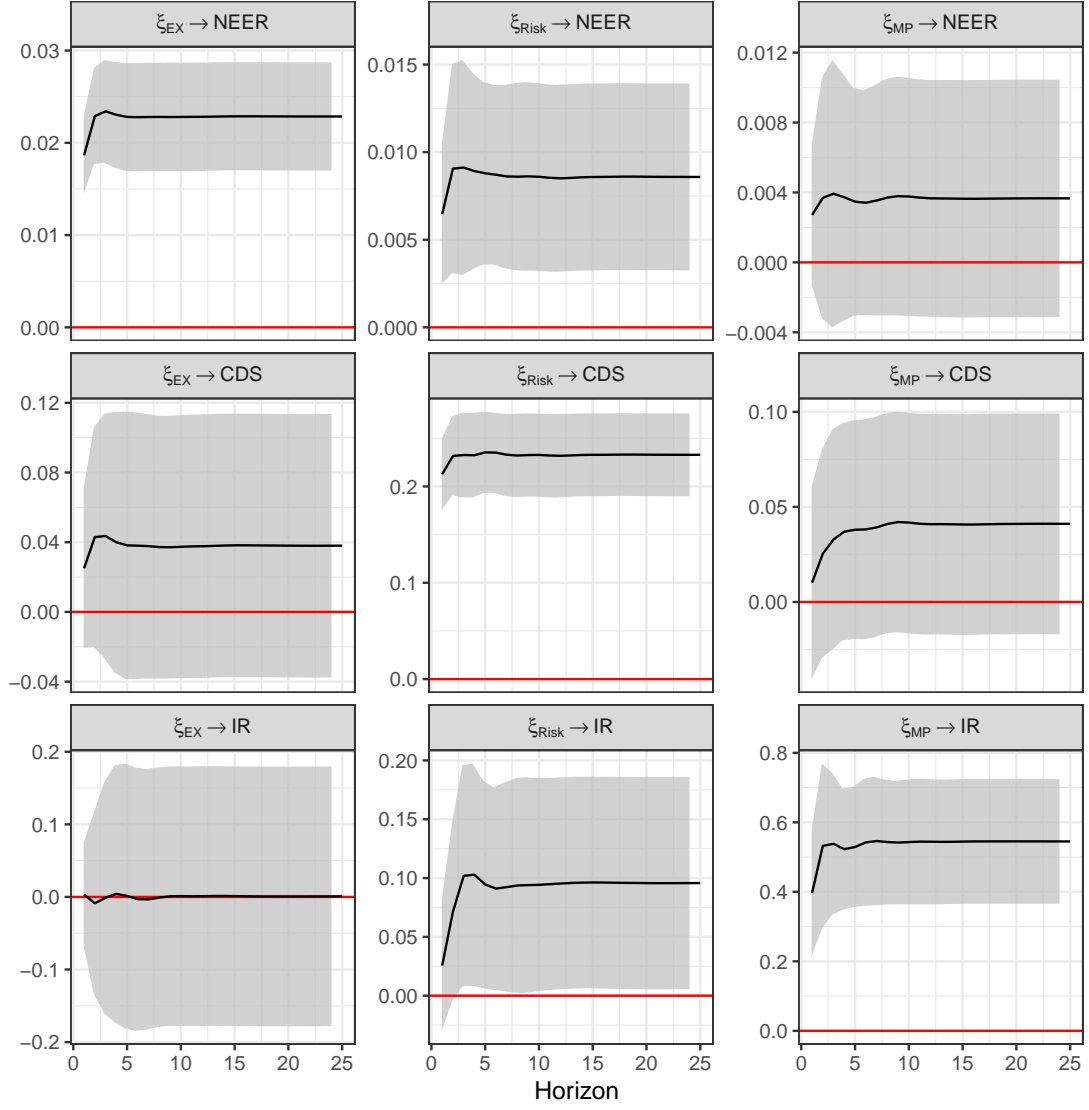


Figure 3: MG cumulated IRFs in NEER systems ($m = 2$). For further notes see Figure 2.

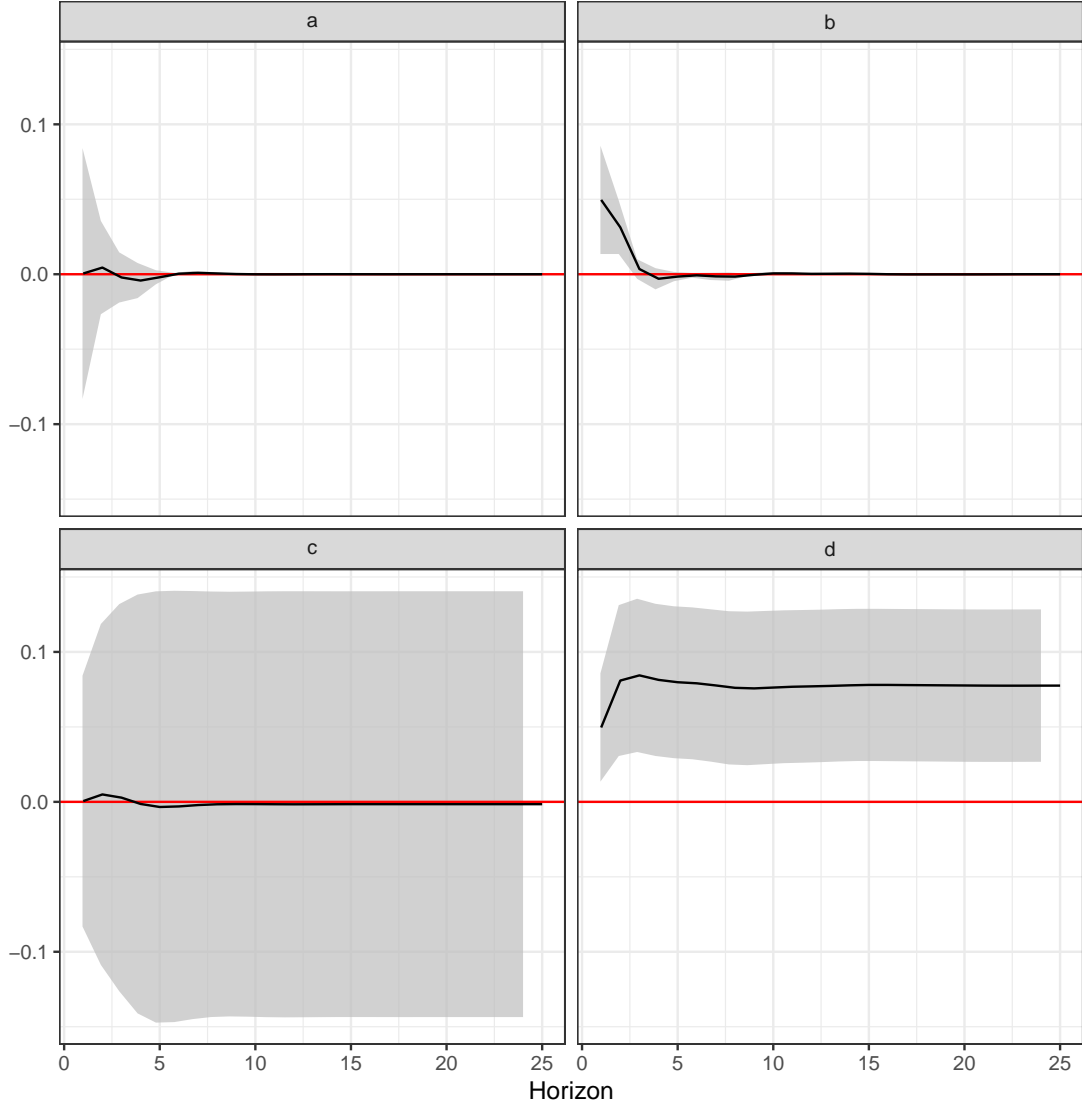


Figure 4: Subsampled effects of effective exchange rate shocks on country risk (SVARs with $m = 2$). Cross sectional subsamples are build with respect to the currency mismatch of the non-government sector relative to exports (Chui et al., 2016) (see also Figure 2). Results for Group 1 (Group 2) are displayed in panels a,c (b,d). IRFs (cumulated IRFs) are displayed in panels a,b (c,d).

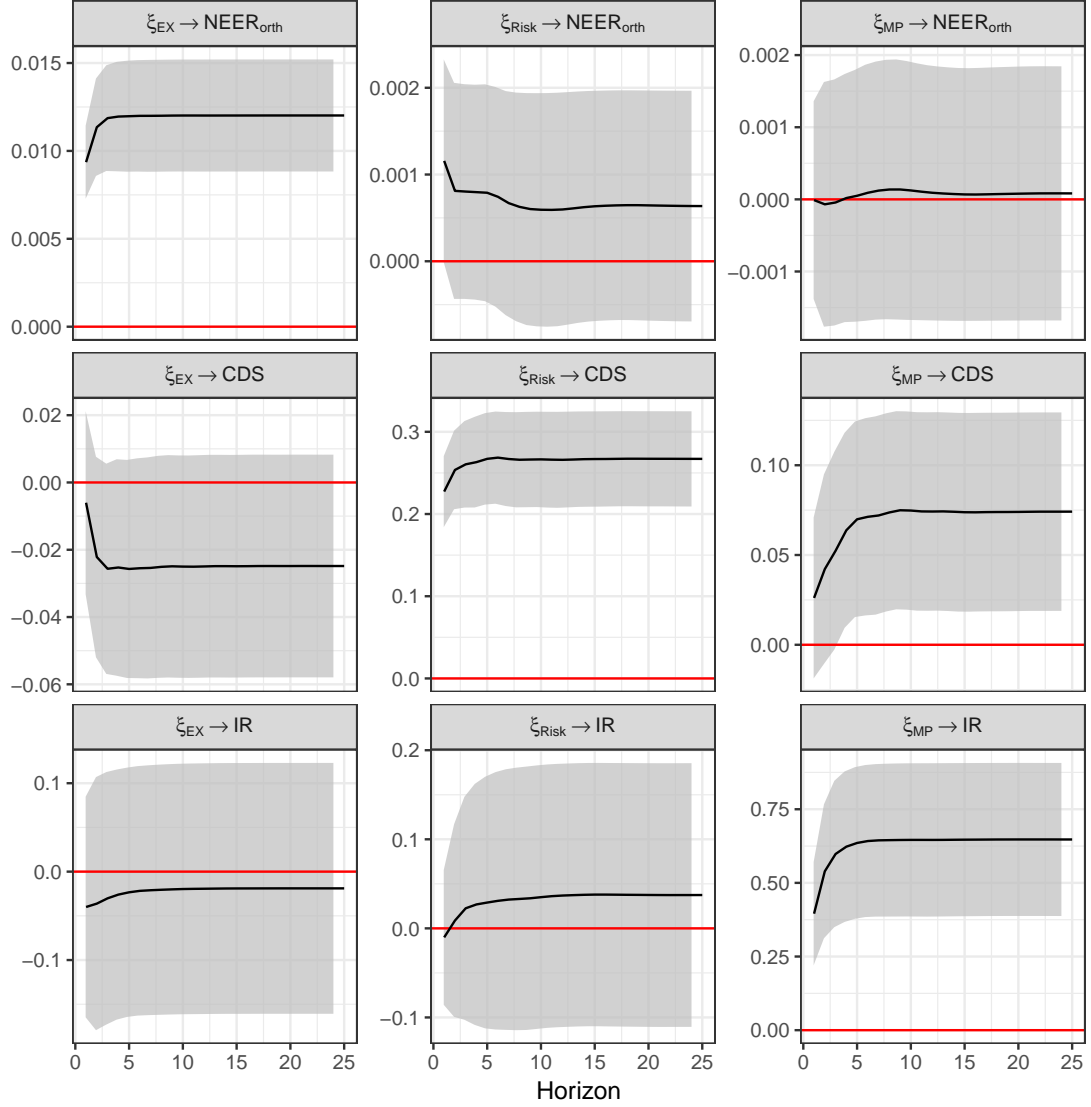


Figure 5: MG cumulated IRFs in NEER_{orth} systems ($m = 3$). For further notes see Figure 2.