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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) using 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. We find that the direction and size of the response of sovereign risk to FX rate movements depend on the type of exchange rate measure we look at 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, 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 of the private sector. In this case, a depreciation of the NEER also induces an increase in sovereign risk. 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'.

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 many countries have also increased. For emerging market economies (EMEs), in particular, such exposures are very often sizeable (Bénétrix et al., 2015; Lane & Shambaugh, 2010), affecting both the public and private sectors (Chui et al., 2016).[1] As a result, the wealth of these countries is exposed to valuation effects in response to foreign exchange (FX) rate movements. Put differently, these countries' solvency may not just be determined by domestic factors, but also by international factors and the value of their currency.

In which direction FX rate adjustments affect sovereign risk is theoretically ambiguous. According to the traditional Mundell-Fleming model, a domestic currency depreciation increases net exports and exerts a beneficial effect on a country's competitiveness ('net export channel'). As a result, aggregate output rises, thus inducing downward pressures on sovereign risk. However, a domestic currency depreciation may also have the opposite effect on a country's default probability. If a country 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. This situation is of particular relevance for many EMEs. The balance sheets of domestic borrowers in foreign currency are weakened, which deteriorates their credit position, leading to a credit contraction,[2] and increases borrowing costs. This worsening of domestic financial conditions dampens economic activity, ultimately increasing the risk of default as perceived by investors. As a result, the described so called 'financial channel' of FX rates adjustments can be regarded as a potential offset to the traditional net export channel.[3] Hence, if a depreciation of the domestic currency increases or reduces sovereign risks depends on the predominance of either the net export channel or the financial channel. This predominance, in turn, depends on the sensitivity of domestic balance sheets to FX rate changes.

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 export 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), we argue that the response to

[1] Many corporations, particularly in EMEs, borrow in international currencies because they find it difficult to borrow abroad in their own currencies (Eichengreen & Hausmann, 1999).

[2] A credit contraction might be triggered by reduced cross-border capital inflows in response to an appreciation of the US Dollar (USD) (Bruno & Shin, 2015a) or by a decline in the leverage of the domestic banking sector (Bruno & Shin, 2015b).

[3] The financial channel in the related literature is also called the 'risk-taking channel of FX rates' introduced by Bruno & Shin (2015a).

a change in the nominal effective exchange rate (NEER) reflects the joint effects resulting from the 'net export' and the 'financial channel'. Thus, the overall effect of a currency depreciation on domestic sovereign risk could be positive or negative, depending on the size of a country's net foreign currency exposure and the trade intensity with countries providing global funding currencies, primarily the United States.[4] In contrast, when a country holds a negative foreign currency position, the response of sovereign risk measures to the bilateral FX rate against the global currencies, is expected to be dominated by the 'financial channel'. We focus on the USD FX rate, since the USD is the dominant medium of international financing for most EMEs (Bénétrix et al., 2015; McCauley et al., 2015). Therefore, changes in the USD FX rate are supposed to exert the largest valuation effects on EMEs balance sheets. For countries that are net short in USD, sovereign risk should increase with an increase in the bilateral FX rate (depreciation) against the USD.

While the empirical literature studies the balance sheet implications of FX rate movements comprehensively (see, amongst others, Lane & Milesi-Ferretti, 2007; Gourinchas & Rey, 2007), 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) 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.

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). Rather than using ad-hoc Cholesky decompositions, 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.[5]

[4] Kuruc et al. (2017) show that the USD is still the most important global currency with respect to foreign currency investment positions. Bénétrix et al. (2015) find for a sample of 117 economies that a minority (i.e. 36 percent) is net short in foreign currency, but almost one half of these economies were short in USD.

[5] Various 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., 2017; Matteson & Tsay, 2017; Herwartz, 2018a,b). For a textbook treatment of data based contributions to SVAR identification, see Kilian & Lütkepohl (2017).

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. Ignoring such a bidirectional causality might lead to biased structural estimates. Fostering joint endogeneity even further, financial variables, like FX rates, interest rates, and CDS spreads, typically respond quickly to news entering integrated markets.

While our identification scheme is mainly of statistical nature at the level of a country specific SVAR, the economic labelling of the identified shocks is strengthened by the fact that labeled identification outcomes respect the coherence of two models that we quantify for a given economy. After statistical and economic identification, we use the cross sectional dimension of the data for purposes of mean group (MG) estimation (Pesaran & Smith, 1995).[6] In particular, after identifying unrestricted structural models, MG inference allows to test the triangularity restriction rather than assuming it from the outset as, for instance, in Hofmann et al. (2016).

We find that the direction and size of the response of sovereign risk measures to a depreciation shock depend on the type of exchange rate measure we look at 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 whose private sector has 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.

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

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

emerging literature that explains this phenomenon with the existence of a 'global financial cycle' (Rey, 2015; Bruno & Shin, 2015a). It is argued, 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 decrease of sovereign risk in countries that tend to have a negative net USD exposure. As a result, the default risk premia contained in long-term interest rates decline while 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 of the private sector 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.

2 Structural VARs with independent shocks

This section provides an outline of the approach adopted for SVAR identification. We (i) briefly sketch the VAR in reduced and structural form and state the identification problem; (ii) mention the sufficient conditions for uniqueness of independent structural shocks; (iii) provide implementation details for the extraction of independent components; and (iv) explain how we use 'external' information for model identification. As 'external' information, on the one hand, we consider the availability of cross sectional results that deserve a cross sectionally coherent identification. On the other hand, we consider the three complementary country specific SVARs to process similar (if not identical) shocks providing both external information and an additional issue of coherent identification.

2.1 The VAR in reduced and structural form

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.[7] We consider country specific VARs throughout. To simplify notation we refrain

[7]We 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 exchange rate regime. Due to limited data availability, our cross section has been reduced further to overall 16 EMEs.

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

$$\begin{aligned} y_t &= \nu + A_1 y_{t-1} + \ldots + A_p y_{t-p} + G x_t + u_t, && (1) \\ &= \nu + A_1 y_{t-1} + \ldots + A_p y_{t-p} + G x_t + D\xi_t,\ t = 1, 2, ..., T, && (2) \end{aligned}$$

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) and Avdjiev et al. (2018), an estimate of NEER changes that are considered unrelated to changes of USD FX rates (NEER$_{orth}$, $m = 3$). These orthogonal components are calculated by means of linear regressions of first differences of the NEER on first differences of nominal (log) FX rates and subsequent residual extraction. Throughout, we refer to exchange rates of a domestic currency in direct quotation. Hence, 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 particular, it contains five variables measuring global financing conditions and risk aversion as well as the economic and the financial situation in the US and the fiscal stance of the domestic government ($Q = 5$). The first four factors are approximated by the volatility index of the Chicago Board Options Exchange (VIX), US industrial production, US CPI (all measured in log-changes), and the change in the US 3-month money market rate. To approximate the fiscal stance of an EME,[10] we use data on the level of government debt as a ratio of

[8]NEER 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. Therefore, we calculate the inverse of the NEER and multiply it with 100, such that the scaling and sign of its changes is comparable to that of the first difference of the log USD FX rate.

[9]At 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.

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

GDP. 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.[11] 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, $A_1, A_2, \ldots, A_p$ are $K \times K$ parameter matrices, and G is a $K \times Q$ parameter matrix. Presample values $y_0, y_1, ..., y_{1-p}$ are assumed available.

The autoregressive representation in (1) characterizes the jointly endogenous variables 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 $A(L) = I_K - A_1 L - A_2 L^2 - \ldots - A_p L^p$ and L is the lag operator (e.g. $Ly_t = y_{t-1}$). Hence, the model implied moving average representation consists in the form of a convergent series and the structural shocks are fundamental. Including covariance estimation, the reduced-form model in (1) can be quantified consistently up to an $o_p(1)$ error by means of OLS/ML estimation.

The structural model in (2) formalizes the transmission of structural shocks $\xi_t \sim (0, I_K)$ to reduced-form disturbances u_t, i.e.

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

By assumption, latent structural shocks ξ_t and reduced-form disturbances u_t obey a linear relation, $u_t = D\xi_t$, where D is nonsingular. In the remainder of this Section we discuss the *data based* determination of D. Since the statistical identification does not build upon behavioural economic relationships, it is not clear if the effects summarized in D also allow for *theory based* characterization of the identified shocks (Herwartz & Lütkepohl, 2014). Therefore, the issue of 'shock labelling' addressed in the empirical Section 5, provides important complementary insights into the properties of the statistically identified independent shocks.

2.2 Uniqueness of independent shocks

In Gaussian VARs ($u_t \sim N(0, \Omega)$), all factors D of the covariance matrix Ω (e.g. a lower triangular Cholesky factor along with all its rotations) are observationally equivalent, since rotations of Gaussian random vectors do not affect the joint distribution. Hence, in a Gaussian framework, 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, the following distributional assumptions are added

[11] All variables have been drawn from either from Bloomberg or Datastream.

to the SVAR representation in (2):

$$
\begin{aligned}
(i)\ & \xi_t \sim F(0, I_K), \\
(ii)\ & \xi_{kt},\ k = 1, \ldots, K, \text{are jointly independent, and} \\
(iii)\ & \text{at most one marginal distribution in } F \text{ is Gaussian.}
\end{aligned}
$$

Building upon the result of Comon (1994), newer approaches to data based SVAR identification suggest alternative ways to extract independent components from the data.[12] For instance, Moneta et al. (2013) suggest loss statistics to determine optimal orderings in recursive systems. Maximum likelihood (ML) and Pseudo ML techniques of Lanne et al. (2017) and Gouriéroux et al. (2017), respectively, cope with fully interdependent structural models. A consistent nonparametric approach to identification that relies on independence testing is suggested by Matteson & Tsay (2017). In this study, we follow a similar identification scheme suggested in Herwartz (2018a). As a particular distinction between the identification schemes in Matteson & Tsay (2017) and Herwartz (2018a), the loss function employed in the context of the latter approach is scale free and, hence, is supposed to be (more) robust in light of data and model heterogeneities that characterize our panel.

2.3 A nonparametric approach to extract independent components

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 ($\widehat{\Omega}$) decompositions, i.e. 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 $\widetilde{D}_\theta = QR_\theta$ 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

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

[12] At first glance, the explicit assumption of independent structural shocks appears as an 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.

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 a space of potential structural shocks, $\tilde{\xi}_t = \widetilde{D}_\theta^{-1}\hat{u}_t$. To detect the particular member of this space that consists of least dependent (or independent) and, hence, unique shocks, model identification requires, secondly, a statistical measure of (in)dependence. In light of cross sectional and within country data heterogeneity, the detection of independent components in this work follows Herwartz (2018a) and relies on the minimization of the rank based Cramér-von-Mises (CvM) distance (Genest et al., 2007)

$$\mathcal{B} = \int_{(0,1)^K} \left[\sqrt{T}\left(C(\tilde{\xi}) - \prod_{k=1}^{K} U(\tilde{\xi}_k)\right)\right]^2 d\tilde{\xi}. \tag{4}$$

In (4), random vectors $\tilde{\xi}$ are obtained from orthogonalizing reduced form model disturbances, C and U denote the empirical copula of orthogonalized model disturbances and the implied copula under independence, respectively.[13]

Conditional on the dependence criterion in (4), an 'estimator' of D (or θ) solves the minimization problem

$$\widehat{D} = \widetilde{D}_{\hat{\theta}}, \text{with } \hat{\theta} = \text{argmin}_\theta\{\mathcal{B}|\tilde{\xi}_t = \widetilde{D}_\theta^{-1}u_t\}. \tag{5}$$

The solution of (5) requires nonlinear optimization. The estimated covariance decomposition obtains unique structural shocks $\hat{\xi}_t = \widehat{D}^{-1}u_t$ under the assumptions in (**??**).[14]

2.4 Global identification

While the minimization problem in (5) obtains a unique solution for the sample of implied structural shocks, their joint independence is invariant with regard to exchanging one shock $\{\hat{\xi}_{kt}\}_{t=1}^T$ for another $\{\hat{\xi}_{jt}\}_{t=1}^T$ or multiplying an estimated shock with minus unity ($\{-\hat{\xi}_{kt}\}_{t=1}^T$). Put differently, with $d_{.,k}$ denoting the k-th column of D and noticing that $\Omega = \sum_{k=1}^K d_{\cdot k}d'_{\cdot k}$, uniqueness of any decomposition $\Omega = DD'$ holds only up to the ordering and the sign of the columns $d_{\cdot k}$. For our panels of SVARs, we adopt the following scheme to establish sign uniqueness and model coherence:

[13] As 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 ($i = 1, 2, \ldots, 17$) 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).

[14] We 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).

(i) Identification in single country models

To establish uniqueness of sign and column ordering for each of the economy and model specific estimates $\widehat{D}_{i,m}$, we compare six alternative column orderings of initial solutions obtained from optimization problem in (5). In the case that one of these six alternative matrices has diagonal entries that are negative, we multiply the respective column with minus unity.[15] Then, we choose from the six alternatives the particular candidate having the maximum sum of diagonal entries.[16] This matrix is denoted $\widetilde{D}_{i,m}$.

(ii) Coherent identification in sets of SVARs $\widetilde{\mathcal{D}}_i = \{\widetilde{D}_{i,1}, \widetilde{D}_{i,2}\}$ (and $\widetilde{\mathcal{D}}_i = \{\widetilde{D}_{i,1}, \widetilde{D}_{i,3}\}$)

From country specific identification we obtain two models ($m = 1, 2$ and $m = 1, 3$) for each country which share the same second and third variable (y_{2t}, y_{3t}). Hence, it is most likely that their dynamics are driven by two similar (if not common) shocks. Noticing that, by construction, the first element in $\xi_t^{(m)}$ exerts its strongest effects on the exchange rate variables, we assume that the cross model similar shocks are in the second and third position of $\xi_t^{(m)}$, $m = 1, 2, 3$. Hence, these shocks correspond to the second and third column of the structural matrices $\widetilde{D}_{i,1}$ and $\widetilde{D}_{i,2}$ ($\widetilde{D}_{i,1}$ and $\widetilde{D}_{i,3}$). To establish model coherence between two country specific SVARs in $\widetilde{\mathcal{D}}_i$, we determine implied structural shocks

$$\xi_t^{(m)} = \widetilde{D}_{i,m}^{-1} u_t^{(m)},\ m = 1, 2, (m = 1, 3,)$$

and investigate the correlation patterns among $\xi_t^{(1)}$ and $\xi_t^{(2)}$ (among $\xi_t^{(1)}$ and $\xi_t^{(3)}$). 'Common shocks' shocks across two model specifications should exhibit strongest correlations. We therefore adopt sign manipulations and column reorderings of the elements in $\xi_t^{(2)}$ (or $\xi_t^{(3)}$) and, hence, $\widetilde{D}_{i,2}$ ($\widetilde{D}_{i,3}$) to minimize the Frobenius distance between the lower right 2×2 block of $\Gamma = \mathrm{Cor}[\xi_t^{(1)}, \xi_t^{(2)'}]$ ($\Gamma = \mathrm{Cor}[\xi_t^{(1)}, \xi_t^{(3)'}]$), denoted $\Gamma_{2,2}$ and an identity matrix.[17] By means of this minimization step we ensure a maximum correlation among shocks in distinct models measuring similar (if not identical) 'surprises'.

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, then discuss their empirical evidence. This guides us in plausibly determining the impact effects that one

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

[16] Maximizing the diagonal entries in absolute value is in line with the presumption that a given structural shock exerts the strongest effect on the variable to which it is primarily associated (Lütkepohl & Netšunajev, 2014).

[17] The Frobenius norm F reads as $F = \sum_{k=1}^{2} \sum_{j=1}^{2} H_{kj}^2$ with $H = \Gamma_{2,2} - I_2$.

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,[18] 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 default 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. a reduction 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 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.

Noticing the interrelations within the triad of exchange rates, sovereign risks and interest rates, it is particularly difficult to identify the contemporaneous transmission channels. Hofmann et al. (2016) suggest solving the identification problem by means of Cholesky factors and assume that an exchange rate shock shapes all variables in the model on impact. In turn, exchange rates do not respond on impact to shocks entering the system through sovereign risk or interest rates. Noticing that pricing on financial markets including money markets usually happens instantaneously in response to the arrival of news, this identifying assumption might be critical.

[18] For 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.

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 distinguished hypotheses below.[19] Probably the most established theory is that exchange rate movements affect the economy through trade effects. Along the 'net export channel' a currency depreciation improves an economy's international competitiveness and, thereby, its trade balance. As a result, production increases, which induces a reduction of sovereign risk. However, when a country is a net borrower in a foreign currency – most likely the USD in the case of EMEs – currency depreciations could also have adverse consequences for sovereign risks. Assume that domestic borrowers are short in a global currency. Then, a depreciation of the domestic currency against this international currency would induce negative valuation effects on borrowers' balance sheets (Bénétrix et al., 2015), weaken their credit position, exacerbate existing banking sector distress, and induce a credit contraction (Bruno & Shin, 2015a). As a result, sovereign risk increases. This so called 'financial channel' of exchange rates can be regarded as a potential offset to the traditional net export channel.

This paper investigates the relative strength of the net export channel and the financial channel by analyzing the effect of a currency appreciation on sovereign default 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 export and the financial channel and, therefore, can be positive or negative depending on which channel dominates. One might expect that the described 'financial channel' is stronger, the larger is an economy's net liability position in international currency. In this respect, 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. In countries where the private sector largely relies on foreign currency denominated debt, the costs of inflating away sovereign debt are often larger than those of a sovereign default. Thus, independent of the net foreign currency position of the government, sovereign default risk remains when the currency mismatch of the private sector implies large adverse balance sheet effects from a currency depreciation. We therefore assume that the response of sovereign risk measures to a depreciation shock of the NEER is positively related to the size of the private sector's short position in international currency, i.e. the USD. Moreover,

[19] A 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 or adjustments in the risk premium.

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 export channel'. Therefore, the response of sovereign default risk to a depreciation shock to the NEER_{orth} is expected to be negative. These considerations give rise the following three hypotheses to be tested:

H_1: A depreciation shock to the FX rate towards a global currency against which the private sector of a country is short increases sovereign risk.

H_2: 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.

H_3: 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),[20] 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 targeting 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 foreign exchange 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 leave room for a zero or positive response stated in Table 1.

[20] `https://www.imf.org/en/Publications/Annual-Report-on-Exchange-Arrangements-and-Exchange-Restrictions/Issues/2017/01/25/Annual-Report-on-Exchange-Arrangements-and-Exchange-Restrictions-2016-43741`

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 the default 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 most likely depends (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 rise 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 lower 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 90s 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 its commitment to price stability with an increase in its policy rate. 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 the 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'.[21]

[21] '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).

	Shocks		
Variable	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.

4 Cross sectional subsampling for testing $\mathbf{H}_2$

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. As outlined in Section 3, theoretical and empirical evidence suggests that it is predominantly the private sector's net liability position that matters for sovereign risk, and looking at country level foreign currency exposures could be highly misleading. To our knowledge, this information 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 assets and liabilities. Moreover, these statistics only provide information on country's external foreign currency debt and ignores internal foreign currency denominated liabilities; that is, 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.

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, their data set lacks information on Croatia. To fill this gap, we refer to an alternative data set provided by Bénétrix et al. (2015). Their data set also allows for calculating the net foreign currency debt exposure, but differently to Chui et al. (2016) they ignore, due to a lack of data, local debt denominated in foreign currencies. However, 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 below the sample median in one group (Group 1) and the remaining countries in another group (Group 2).

Figure 1 shows the average net foreign currency exposure relative to exports of the two country groups. The non-government sector of all sample countries but South Africa shows, on average, a negative 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 contained in Group 2 (i.e. Brazil, Chile, Mexico, Korea, Hungary, Poland Turkey, and Croatia), the currency mismatch is larger in magnitude than the sample mean. We therefore expect that the 'financial channel' in this subsample is stronger than for the countries contained 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 H_2, 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 two step structural identification

For a given economy ($i = 1, 2, \ldots, 16$), 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.[22] 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

[22] At the MG level core modelling outcomes are 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%).[23] With regard to respective diagnostics for single components of estimated VAR residuals, it turns out for rare cases that 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.

As outlined in Section 2, our structural identification consists of two steps. (i) maximizing the effects of structural shocks on those variables to which they are primarily associated; and (ii) establishing model coherence of country specific estimates within sets of SVARs. Practically, the second step could be considered to be a correction of first step outcomes. It turns out for both sets of SVARs – $m = \{1, 2\}$ and $m = \{1, 3\}$ – that this correction applies for at most three economies. Structural estimates $\{\widehat{D}_{i,1}, \widehat{D}_{i,2}\}$, $\{\widehat{D}_{i,1}$, and $\widehat{D}_{i,3}\}$ are characterized by the fact that two implied shocks are very similar (i.e. highly correlated) across models. Contemporaneous correlations among the similar shocks are larger than 0.60 with one exception and exceed of 0.90 for the majority of model comparisons. Generally, the contemporaneous correlations among the similar shocks are slightly weaker between sets of SVARs determined from FX and NEER$_{orth}$ ($\{\widehat{D}_{i,1}$ and $\widehat{D}_{i,3}\}$) as it is the case for correlations between sets of SVARs determined from FX and NEER ($\{\widehat{D}_{i,1}$ and $\widehat{D}_{i,2}\}$).

Regarding the correlations between those shocks that have a prime effect on exchange rates, i.e. $\xi_{1t}^{(m)}$, $m = 1, 2, 3$, we diagnose for 11 economies and models $m = 1, 2$ contemporaneous correlations in excess of 0.9. These strong correlations likely reflect the strong impact of bilateral USD prices on the NEER. Accordingly, one might argue that the exchange rate shocks retrieved from SVARs with $m = 2$ are overly spoilt by nominal effects. Not surprisingly, for the identified shocks $\xi_{1t}^{(m)}$, $m = 1, 3$ (FX vs. NEER$_{orth}$), correlations are much less than, and for no economy beyond, 0.48.

In summary, we take these results as indicative for a successful discrimination of two common and one distinct shock among model variants with FX rates on the one hand ($m = 1$) and NEER or NEER$_{orth}$ ($m = 2, 3$) on the other hand. Since the effects of the distinct shock $\xi_{1t}^{(m)}$ are of particular interest for addressing the hypotheses H_1 to H_3 raised in Section 2, it is worth emphasizing that this shock is unique and well identified in a statistical sense and in terms of cross model coherence. Moreover, correlation patterns between $\xi_{1t}^{(1)}$ and $\xi_{1t}^{(3)}$ indicate that variables in NEER$_{orth}$ are suitably adjusted for nominal effects.

MG results for the full panel are displayed in Table 2. The tables show 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

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

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.

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.

The estimated contemporaneous effects of all three shocks, i.e. an exchange rate shock, a risk shocks, and an interest rate shocks 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, in the following. These effects are of core interest for investigating the hypotheses H_1 to H_3 raised in Section 3.

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.[24] Figure 2 shows the corresponding cumulated impulse response functions (IRF). The dynamic effect characteristics indicate that the positive impact of a FX rate shock on sovereign risk is very

[24] Table 2 documents the MG estimator of the structural parameter d_{21}. When interpreting the estimate 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 $0.025 = \sqrt{6.35\text{E-}04}$ as a reasonable approximation of the magnitude of a shock of size one standard deviation.

persistent over time.[25] 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'.

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.[26] However, while the described on-impact effect is highly significant in the FX system, it lacks significance at conventional levels in the NEER system. The impulse response function shown in Figure (Figure 3) confirms 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 export' and the 'financial channel' and, therefore, can be positive or negative depending on which channel dominates. Both 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.

As formalized further in hypothesis H_2, we assume that the financial channel becomes more dominant when the net foreign currency exposure of a country's private sector 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 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 mean. Subsample structural estimates are broadly in line with H_2. When focusing only on countries with a large foreign currency exposure (Group 2), we diagnose a

[25] Identified 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. Let $\Phi(L) = \sum_{i=0}^{\infty} \Phi_i L^i$ be an operator such that $\Phi(L)A(L) = I_K$ (Lütkepohl, 2007, Chapter 2). From systematic comparison of coefficients in $\Phi(L)A(L)$ one obtains (see Lütkepohl 2007, Chapter 2) $\Phi_0 = I_2$, $\Phi_1 = \Phi_0 A_1$, $\Phi_2 = \Phi_1 A_1 + \Phi_0 A_2, \ldots, \Phi_i = \sum_{j=1}^{i} \Phi_{i-j} A_j,\ A_j = 0$ for $j > p$. Owing to the autoregressive representation in (1), the (backward looking) moving average representation of the conditionally dynamic components of y_t exists and reads as

$$\tilde{y}_t := y_t - A(L)^{-1}\nu - A(L)^{-1}Gx_t = \sum_{h=0}^{\infty} \Phi_h u_{t-h} = \sum_{h=0}^{\infty} \Psi_h \xi_{t-h}, \quad \Phi_0 = I_K, \Psi_h = \Phi_h D. \tag{6}$$

Considering an isolated unit structural shock to occur in variable j, $\xi_{jt} = 1$, the (i, j)-th element of a matrix $\Psi_h = \Phi_h D$, $h = 0, 1, 2, \ldots$, in (6) is the h-step ahead response of the i-th variable, $\tilde{y}_{i,t+h}$. Estimated IRFs are obtained by replacing true model parameters in (6) by their respective estimators.

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

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 favor of H_3 are also confirmed by the IRFs shown in Figure 5. 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.

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}, $m = 3$).[27] 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 panel of Table 2 and the corresponding IRFs are shown in Figure 4. 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.[28] 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'.

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}), structural parameter located above the main diagonal are are not uniquely insignificant at the MG level. Accordingly, although country specific results show sizeable variability throughout (see e.g. the often detected sign

[27] Alternatively, 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'.

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

changes between the first and third quartile of non-diagonal estimates), MG estimates and diagnostics indicate non triangularity.[29] 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 transmission 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

$$D_1^{\Delta} = \begin{bmatrix} \underset{(14.2)}{2.432} & - & - \\ \underset{(8.88)}{12.12} & \underset{(10.2)}{21.06} & - \\ \underset{(1.69)}{5.799} & \underset{(2.39)}{4.294} & \underset{(4.17)}{41.73} \end{bmatrix}, \; D_2^{\Delta} = \begin{bmatrix} \underset{(8.63)}{2.213} & - & - \\ \underset{(7.013}{10.55} & \underset{(10.0)}{21.86} & - \\ \underset{(2.03)}{5.722} & \underset{(2.82)}{5.480} & \underset{(4.39)}{40.39} \end{bmatrix} \quad (7)$$

$$\text{and } D_3^{\Delta} = \begin{bmatrix} \underset{(9.34)}{1.006} & - & - \\ \underset{(0.92}{1.012} & \underset{(11.6)}{24.82} & - \\ \underset{(1.67)}{1.983} & \underset{(2.46)}{7.029} & \underset{(4.17)}{42.82} \end{bmatrix} \quad (8)$$

Comparing the results in (8) with those in Table 2 shows that for systems of SVARs ($m = 1, 2$) 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

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

[29] We 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.

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 default risk,; this differentiation allows us to discriminate between the traditional 'net export 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, 2018a), which 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 exchange rate measure we look at 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 turns out to have only a significant effect on sovereign risk for countries with large negative net foreign currency exposures of the private sector. 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 USD FX rates, 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'. Hence, only for countries that are net long 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 Global 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. The decrease in sovereign risk fuels capital inflows, thereby stimulating economic growth. To counteract the eventual inflationary pressures, the Central Bank of an affected EMEs might consider raising policy rates. However, this would aggravate domestic currency appreciation further and a vicious circle would start. Alternatively, the Central Bank of thean affected EME could respond to appreciation pressures by also lowering its policy rate, which would come at the cost of higher inflation and compressed economic activity. Thus, monetary policy autonomy in countries with large foreign currency exposures in 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 less foreign currency denominated government debt, while the private sector in most EMEs still borrows to a large extent in foreign, mostly global, currencies like such as the USD. Hence, advisable policies might target at an enhancingement 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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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

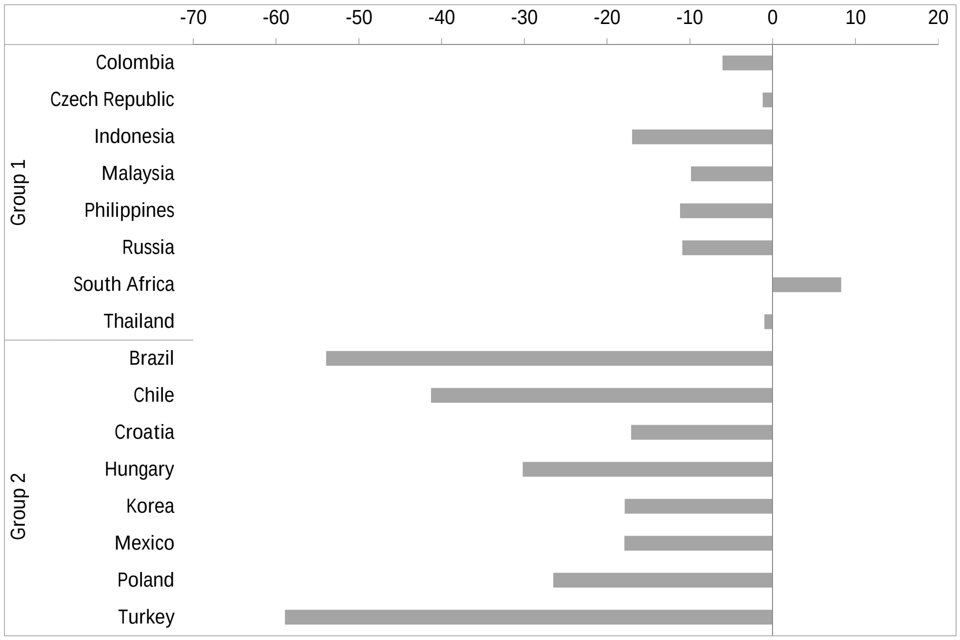

Figure 1: Currency mismatch of the non-government sector relative to exports (Chui et al., 2016). This is also the subdivision behind results documented in Table 3.

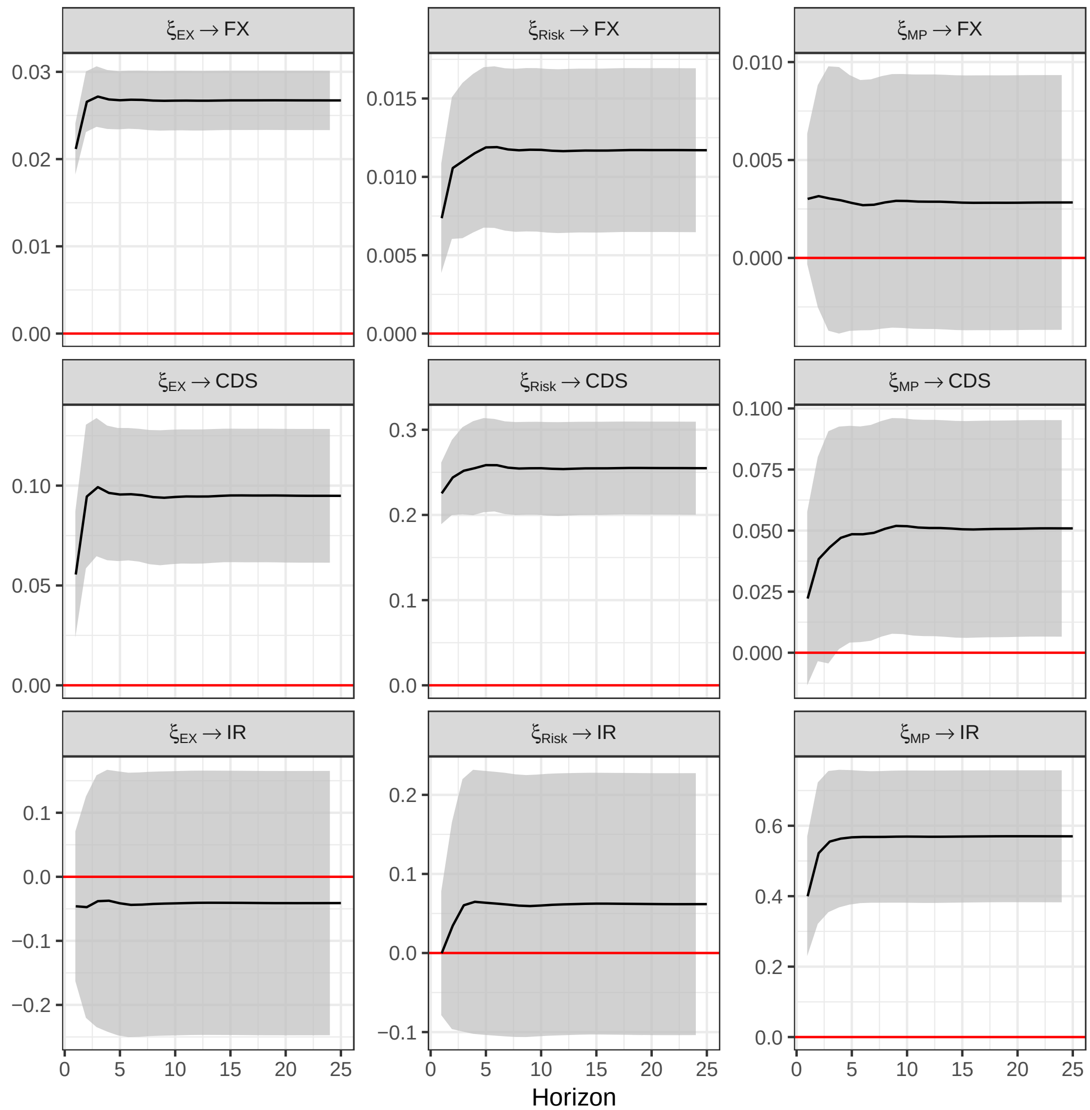

Figure 2: MG cumulated IRFs in FX systems ($m = 1$). Exchange rate, country risk and monetary policy shocks are indicated with 'EX' 'Risk' and 'MP', respectively. The shaded areas are pointwise MG confidence intervals with 95% coverage.

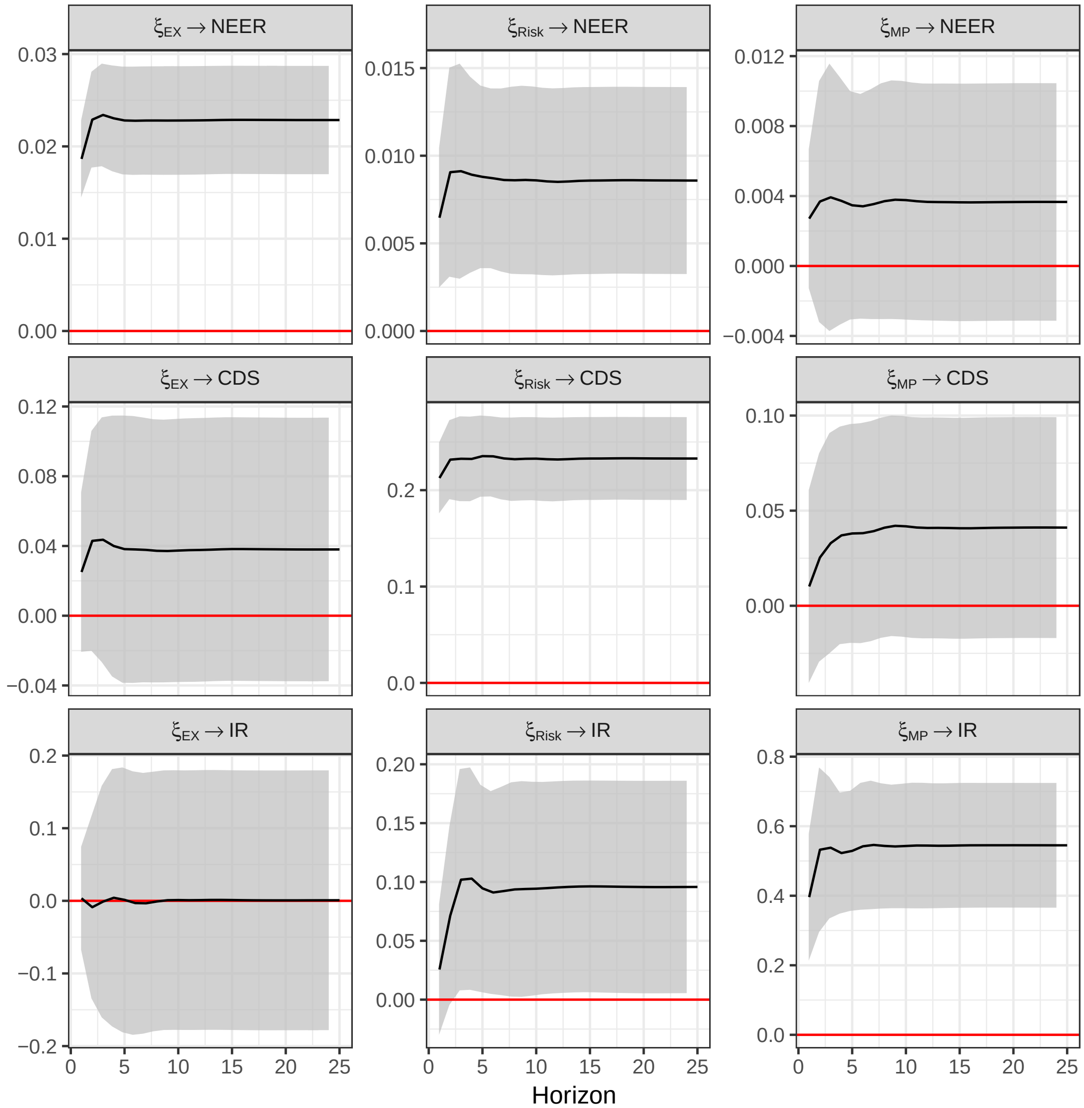

Figure 3: MG cumulated IRFs in NEER systems ($m = 2$). For further notes see Figure 2.

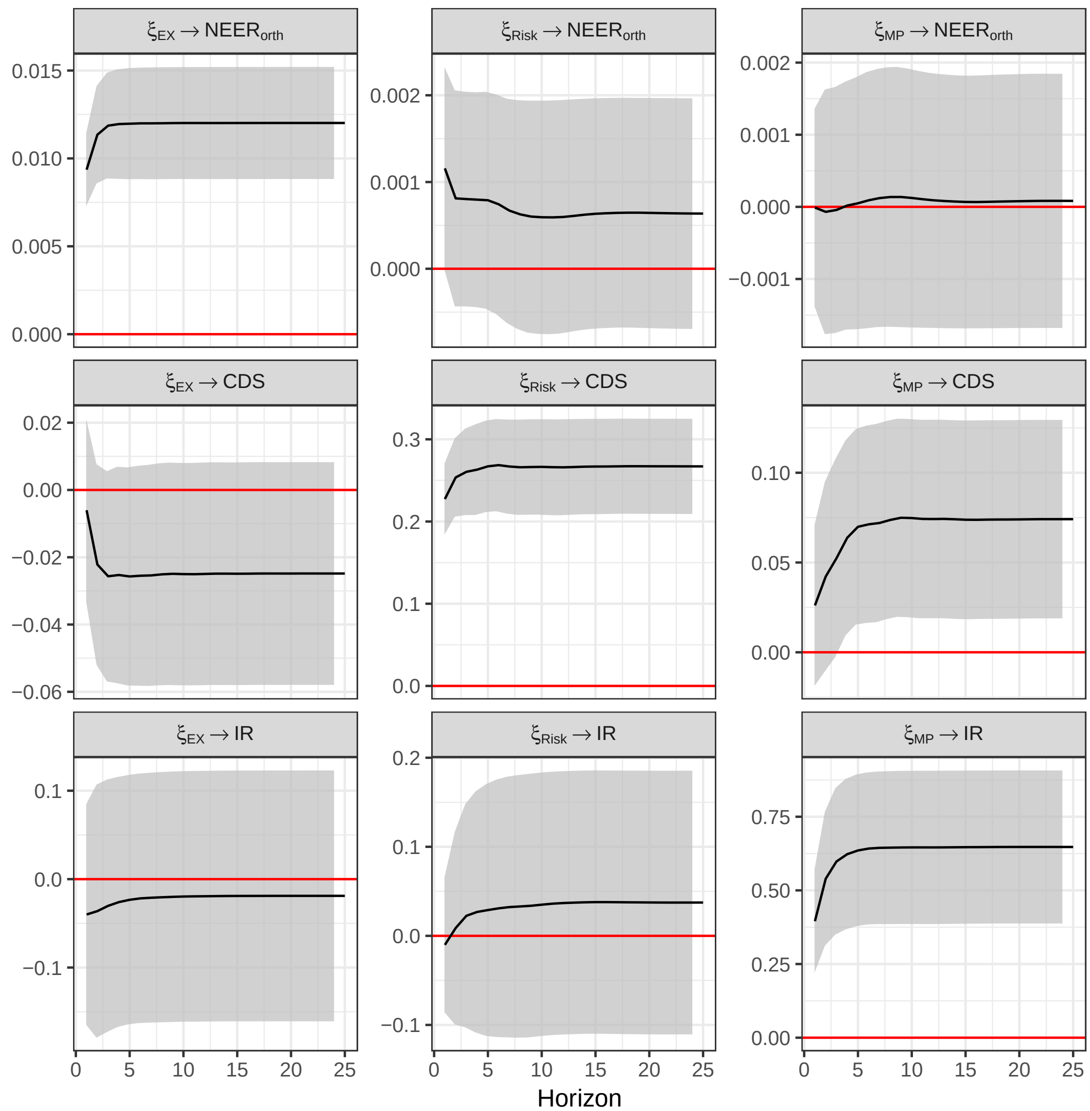

Figure 4: MG cumulated IRFs in NEER$_{orth}$ systems ($m = 3$). For further notes see Figure 2.

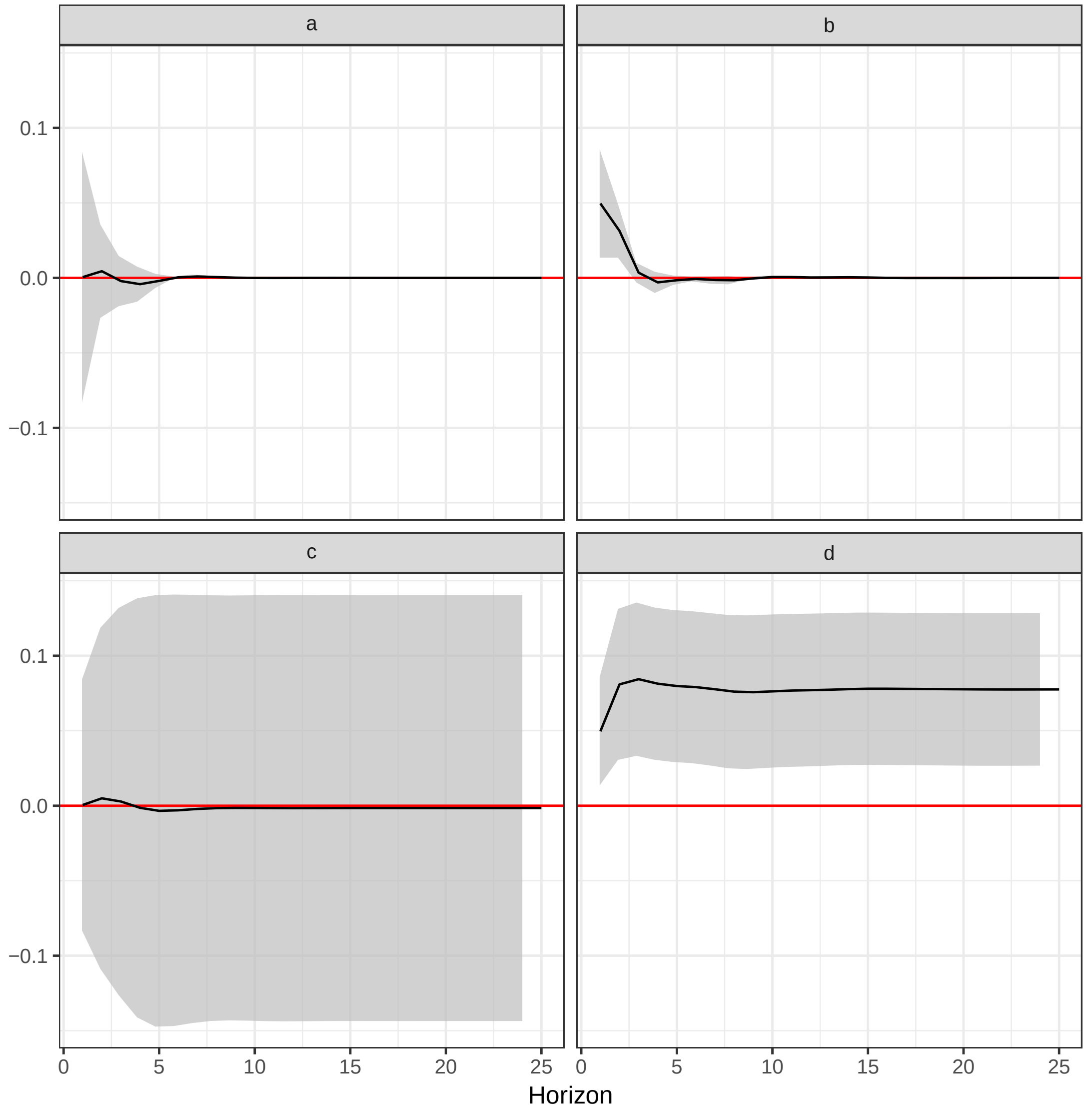

Figure 5: Subsampled effects of effective exchange rate shocks on country risk (SVARs with $m = 2$). Cross sectional subsamples are build with respect to the currency mismatch of the non-government sector relative to exports (Chui et al., 2016) (see also Figure 2). Results for Group 1 (Group 2) are displayed in panels a,c (b,d). IRFs (cumulated IRFs) are displayed in panels a,b (c,d).