Country Risk Premia, Endogenous Collateral Constraints and Non-linearities: A Threshold VAR Approach

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Abstract

The notion of occasionally binding constraints has been used in macroeconomic models to generate amplified financial accelerator effects - in particular for emerging market business cycles. As much as these models have to use global solution techniques, empirical models have to resort to non-linear estimation techniques to capture asymmetries. Using a threshold vector autoregression approach, I analyze the effect of shocks to the country risk premium in different regimes which are interpreted as states of the economy where collateral constraints bind to a different degree. Amplification coefficients measuring the non-linearity of responses are computed across various emerging market economies. First, the results show that there is large heterogeneity in the responses and the size of amplification coefficients. Second, these cross-country differences can be associated with characteristics such as liability dollarization or external leverage. This validates the underlying conceptual framework where vulnerability at the country level is assumed to depend on the degree of financial frictions. Third, this paper shows that both a debt-deflation mechanism which causes asset price spirals as well as pecuniary externalities stemming from exchange rate depreciation can lead to non-linearities; however, the former is associated with a higher likelihood of leading to regime switches.

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1 Introduction

Financial accelerators have gained tremendous popularity and are now an integral part of most macroeconomic models ever since their introduction in business cycle models à la Bernanke and Gertler (1989) and Bernanke *et al.* (1999). Financial accelerators amplify shocks whenever financial frictions generate inefficiencies that shorten the supply of credit. In particular, the importance of financial frictions has been emphasized in models of emerging market (EME) business cycles.[1] However, the quantitative importance of the simple financial accelerator stemming from an *always binding* credit constraint has been found to be small (Kocherlakota, 2000; Christensen and Dib, 2008). In contrast, *occasionally binding* constraints have been introduced to macroeconomic models in order to study asymmetries in business cycles (see Bianchi, 2010 for a short introduction). In an EME context, these models capture typical phenomena of asymmetry such as credit booms and sudden stops.

The present paper addresses the issue of occasionally binding constraints *empirically*. If periods of binding and non-binding constraints can be identified in the data, an estimate of the pure amplification effect due to binding constraints is feasible. In particular, I analyze two channels that have been discussed in the literature as triggers of non-linearities. In the context of EME crises, occasionally binding credit constraints can lead to non-linearities whenever: (1) asset price reductions deplete entrepreneurs' net worth and result in further asset price collapses (Mendoza, 2010) and (2) contractionary exchange rate depreciations reduce net worth in the presence of liability dollarization and lead to additional depreciation (Bianchi, 2011; Benigno *et al.*, 2011). I want to test to what extent each of these channels contributes to non-linear financial accelerator effects and therefore use asset prices and the exchange rate as the respective variables via which non-linear effects can be tracked. By doing so, I can identify whether asset price spirals or depreciation spirals (or both) are causing boom-bust cycles.

I am specifically looking at the impact of country risk premium shocks as they represent a type of shock that has the potential to make credit constraints bind and thus trigger non-linear reactions. In particular for emerging market economies, country risk premium shocks are a very relevant source of business cycle fluctuations as interest rates have been shown to be countercyclical (Neumeyer and Perri, 2005; Uribe and Yue, 2006). Fluctuations in the country risk premium can be assumed to be largely exogenous (Calvo *et al.*, 2006): Sudden stop events appear "in bunches" across a large number of heterogeneous economies. Nonetheless, country fundamentals remain important in determining the effect that country risk premium shocks have on the economy.

Credit constraints can be justified from microeconomic principles such as contractual arrangements stemming from information asymmetry. Their straightforward inclusion in macroeconomic models has become a conventional habit while the empirical validation of the relevance of financial accelerators is lacking behind their use in theory. This is partly due to the difficulty in quantifying financial accelerator effects. In *theoretical* models, the amplification effect from (always) binding credit constraints following an adverse

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[1]Any dynamic business cycle model which seeks to capture structural features of EMEs incorporates financial frictions such as credit/collateral constraints, endogenous country risk or contractionary devaluations.

shock is measured in a quasi-experimental set-up by comparing outcomes under binding and non-binding constraints. When credit constraints bind only occasionally, their effect can be measured more directly without imposing exogenously if constraints bind or not. In particular, constraints bind endogenously in response to adverse shocks. As much as theoretical models with occasionally binding constraints have to use global solution techniques (see the literature review in Benigno *et al.*, 2019), empirical models have to resort to non-linear estimation techniques to capture non-linearities. Using a threshold vector autoregression (TVAR) approach, I estimate the behaviour of interest rate shocks, or more specifically shocks to the country risk premium, in different regimes and interpret these regimes as states of the economy with differing degrees of access to borrowing. The largely exogenous character of country risk premium shocks facilitates the identification of a shock process. Amplification coefficients which measure the non-linearity of responses are computed for different types of shocks.

The study of non-linearities and crisis phenomena has – up to the recent financial crisis – been confined to the realm of emerging markets. Balance of payments and currency crises in the 1980s and 1990s triggered a vast research agenda centered around the study of non-linear dynamics. However, the experience of a severe financial crisis that hit industrialized countries during the 2008–09 Global Financial Crisis, the transmission mechanisms and non-linear effects described in the EME literature have seen a revival. In particular, a growing literature relies on the use of non-linear time-series techniques to study, among others, how the transmission mechanisms of monetary and fiscal policy differ in times of crisis.

The results can be summarized as follows. *First*, there is considerable heterogeneity in the responses and the size of amplification coefficients across countries. This finding questions the generality of results stemming from analyses of individual countries. Most studies on emerging markets that are of theoretical nature calibrate their models to Mexican or Argentine data. The present paper, however, shows that results differ to a large extent across EMEs. *In a second step*, this heterogeneity is therefore investigated and it is found that cross-country differences can be associated with characteristics such as liability dollarization or external leverage. It is thus structural features which govern to what extent financial frictions can unfold. *Third*, I find that asset price collapses are associated with higher dependence on regime differences than do exchange rate depreciations. Fire sales and deleveraging are more likely to lead to binding collateral constraints then do the pecuniary externalities of exchange rate depreciation arising in the presence of liability dollarization.

The contribution of this paper lies in its use of non-linear VAR techniques to uncover important non-linearities in emerging market business cycles and to relate them to the specific mechanisms and structural factors that cause non-linear reactions to detrimental shocks. More generally, this approach can advance our understanding of the channels through which financial accelerator effects materialize. Furthermore, the findings suggest that emerging market economies should not always be measured with the same yardstick, as this group of countries is very heterogeneous in its responses to macroeconomic shocks.

Related literature. This paper is related to two large strands of literature: The first investigates the role of financial frictions for EME business cycles and crisis episodes. The

second is in particular concerned with the economic channels that lead to asymmetries in EME business cycles.

Business cycle patterns in EMEs are different from the ones in industrialized countries in numerous aspects (see for example Agénor *et al.*, 2000 or Lane, 2003).[2] The literature has investigated the contribution of interest rates and country risk premia to explain these particular features of business cycles. Early open-economy RBC models concluded that interest rate shocks are negligible for *industrialized* SOEs.[3] This finding, however, does not hold for emerging markets. As documented by Neumeyer and Perri (2005), real interest rates are countercyclical and lead the cycle in EMEs whereas they are acyclical and lag the cycle in industrialized countries. In order to capture this countercyclical nature of real interest rates, EME business cycle models feature country risk premia coupled with financial frictions. In Neumeyer and Perri (2005) and Uribe and Yue (2006), firms are subject to a working capital constraint which enables interest rate movements to influence real activity. Fundamentals affect the country risk premium and, in turn, shocks are amplified via endogenous interest rates. Other papers studying the role of interest rates and country risk premia are García-Cicco *et al.* (2010), Chang and Fernandéz (2013) and Akinci (2013).

The above described country risk premium channel relies on a financial accelerator mechanism; yet, EMEs are often subject to an amplified version of this mechanism due to balance sheet effects following currency depreciation. The ambiguous effect of devaluations in the light of limited access to international capital markets and liability dollarization is a well known phenomenon (Céspedes *et al.*, 2002). It is also commonly accepted that these frictions tend to amplify the effects of various shocks on the economy (see for example Devereux *et al.*, 2006 or Berganza *et al.*, 2004). Various papers account for the special structural features of EMEs by including foreign currency debt as a financial friction in their respective models.

Financial frictions and vulnerabilities not only lead to volatile business cycles, but have also been studied extensively in the literature on banking and currency crises in EMEs. Analyzing the empirical regularities of EME crises, Mendoza (2006) finds that sudden stop episodes are often characterized by deep recessions, sharp current account reversals and price collapses. In order to explain these patterns, the literature on EME

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[2]Business cycles are characterized by higher volatility of macroeconomic aggregates in general compared to industrialized countries who have experienced a gradual decline in the magnitude of fluctuations in the last decades. Another feature of EME business cycles is their higher volatility of consumption vis-à-vis the volatility of output. In comparison to industrialized countries, interest rates (or the country risk premium) are countercyclical. The current account and the trade balance of EMEs tend to be more countercyclical than the ones of industrialized countries. The real exchange rate is procyclical and TOT and GDP volatilities are highly correlated. Limited leeway for macroeconomic policies as well as political economy reasons are often identified to rationalize the procyclicality of macroeconomic policies - in particular fiscal policy. However, the arguments also pertain to monetary policy which tends to be more constrained given the particular exchange rate policy in EMEs.

[3]The seminal paper by Mendoza (1991) argued on the contrary that productivity shocks can explain a large fraction of aggregate volatility in SOEs. His analysis, however, focused on industrialized economies. Quantitative models which estimate the importance of world interest rate shocks have found differing results. Kose (2002) finds a negligible role for EME business cycles whereas Lubik and Teo (2005) identify these shocks as the main driving force of business cycles in both industrialized and developing SOEs. Looking at industrialized SOEs, Blankenau *et al.* (2001) also attribute a rather large role to world interest rate shocks.

crises has introduced various frictions which lead to the amplification and propagation of shocks. For instance, Chari *et al.* (2005) show that sudden stops do not lead to recessions as such; this is only the case when important frictions such as advance payment constraints in combination with collateral constraints provide channels for the generation of output collapses.

Financial frictions thus have served as a source and amplification channel both in the literature on *business cycle* patterns in EMEs as well as in the literature on EME *crises*. That this is not a mere coincidence, but rather two sides of the same coin is emphasized by Mendoza (2010) and Gruss and Mertens (2010): Financial frictions matter for the explanation of EME business cycles exclusively due to their amplification effects in times of crisis. Periods of highly restricted access to international capital markets are embedded in business cycles which otherwise feature similar patterns as those of industrialized countries. As such, business cycles in EMEs are characterized by large asymmetries due to infrequent crisis events. As mentioned above, these asymmetries can best be captured in theoretical models by making collateral or credit constraints only bind occasionally.

Collateral constraints limit the amount of debt which agents can hold as a function of their value of collateral or net worth. Asset holdings and capital can serve as collateral. Equivalently, current income and the value of outstanding debt determine the ability to borrow. Following an adverse shock, several channels have been investigated which can result in binding constraints. *First*, asset price collapses can deplete collateral. Mendoza (2010) relies on a "Fisherian debt-deflation mechanism" which is set into motion when adverse shocks make the collateral constraint binding: as in Aiyagari and Gertler (1999), asset prices fall when agents fire-sale their assets due to deleveraging; as a result, the collateral constraint is reinforced. Bianchi and Mendoza (2010), Mendoza and Smith (2006) and Jeanne and Korinek (2019) build on these mechanisms to study credit booms and busts. A similar mechanism is at work in Perri and Quadrini (2018) where the liquidation value of capital and thus the ability to borrow decreases when expectations about future production are lowered following an adverse shock.

Second, contractionary devaluations in the presence of liability dollarization can lead to a deterioration of net worth. Bianchi (2011) and Benigno *et al.* (2011) show that agents do not internalize pecuniary externalities when deciding about the amounts they borrow. As a result, credit constraints bind after prolonged periods of overborrowing when an exogenous shock leads to a depreciation of the exchange rate. This, in turn, triggers adverse balance sheet effects (Céspedes *et al.*, 2004) and the exchange rate depreciates even more due to lower demand prospects.

Many of these models which highlight the asymmetric effects of collateral constraints emphasize the increase in leverage prior to the point when adverse shocks hit and constraints suddenly bind. Financial frictions are not only the source of unwinding pecuniary externalities, they also incentivize the built-up of vulnerabilities. Among various examples, the role of financial frictions for credit booms and collapses is highlighted by Tornell and Westermann (2002). Banking crisis can often be self-fulfilling as they are the corollary and logical consequence of credit booms. As noted by Fostel and Geanakoplos (2008), instead of representing an undervaluation of assets, abrupt asset price collapses are "corrective" adjustments of asset prices that were overvalued during booms.

The rest of this paper is organized as follows. Section 2 discusses the importance of non-linearities and to what extent the empirical strategy serves as a corollary to theoretical models of endogenous collateral constraints. Section 3 illustrates the empirical strategy as an example for the case of Argentina. The Argentine economy is often used as an example of an emerging market economy and many DSGE models in the EME literature are calibrated using Argentine data. Amplification coefficients for different types of shocks and different initial conditions are computed for each country. Since substantial heterogeneity across the 16 economies used in the sample characterizes these results, section 4 investigates the underlying drivers of these differences. Section 5 concludes.

2 Non-linearities

2.1 Dimensions of non-linearity: disproportionality, asymmetry and initial conditions

Non-linearities can be expected to characterize various economic relationships. For instance, business cycle dynamics are subject to the asymmetric amplification and propagation of shocks as shown by Kocherlakota (2000).[4] Theoretical work in macroeconomics relies increasingly on specifying highly non-linear models. A very prominent example is the financial accelerator mechanism: The pricing function for external financing with regards to entrepreneurs' net worth is convex as in Bernanke *et al.* (1998). Yet, most of applied time series research relies on linear models which might not capture certain effects if these only materialize under particular circumstances. Indeed, non-linearitiy could be one of the reasons why empirical linear models sometimes fail to capture effects which have been demonstrated convincingly in theoretical papers. The corollary applies to theoretical models where higher-order approximations and global solutions are increasingly used in order to account for important dynamics which would be neglected in linear, first-order approximations. This issue is particularly relevant for the literature on financial accelerators which is a standard ingredient in business cycle models, but whose empirical validation falls short of its wide-spread use.

Capturing non-linearities in empirical research is thus a challenging exercise as the development of the theoretical literature has shown. Two main questions can be addressed by non-linear time series models. First, one can study *differences in the characteristics of shocks.* In this respect, different shocks can affect macroeconomic variables disproportionally. Shocks can differ with regards to their direction (positive vs. negative shocks) as well as to their size (small vs. large shocks). Second, non-linearities can arise due to *differences in initial conditions* (regime-dependencies).

It is not hard to imagine that certain shocks trigger effects that change depending on whether the shock is large or small. If the system is characterized by non-linearity, one would then expect *disproportionate effects* as a response to shocks of different *magnitudes.* Equivalently, the *direction* of a shock will lead to *asymmetric effects* if non-linearity is present. Most importantly, these mechanisms can operate to a different extent depending

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[4]See also Acemoglu and Scott (1997); Neftci (1984).

on whether the economy is very vulnerable or not at the time when the shock hits. Thus, in contrast to linear models, *initial conditions* can lead to a heterogeneous propagation of shocks. Initial conditions, or regime-dependencies, describe the point of the business cycle at which the economy is situated when a shock hits. For example, adverse shocks can have a more detrimental impact when the economy is vulnerable. Conversely, one could imagine that reactions of macroeconomic variables are muted when these are away from their steady state due to a higher degree of inertia of those variables that have already taken on extreme values.

2.2   VARs and the true DGP: Why non-linearities matter

Impulse response analysis is a widely used tool in macroeconometrics to assess the dynamic responses to exogenous shocks. Impulse response functions (IRFs) are essentially counter-factual forecasts whose accuracy can be expected to decline as the forecast horizon increases. Most VARs assume the structure of the multivariate system to be linear. As pointed out by Jordà (2005), a crucial question thus is if a (linear) VAR is an appropriate approximation of the true data-generating process (DGP). If this is not the case, misspecification will produce incorrect forecasts and results from IRF analysis will consequently be of little use.

When interpreting a standard VAR as a (linear) approximation of the true DGP, one can define the IRF accordingly. Dropping time indices to simplify notation and defining Y to be a vector of all the lags of the variables in the VAR system as well as the respective parameters, a one-step-ahead forecast following a shock to y_i is defined as:

$$f\left(Y, y_i^o + \varepsilon_i\right) = f\left(Y, y_i^o\right) + \varepsilon_i \, \frac{\partial}{\partial y_i} f\left(Y, y_i^o\right)$$

The last term can be interpreted as the first horizon forecast of the corresponding IRF. The size and direction of the shock ε_i feeds with strict proportionality into the system if the functional form of f in y_i is linear. It is for this reason that no difference between various magnitudes and directions of shocks can be investigated in standard linear VARs; they are, by definition, symmetric. More importantly, whenever f is non-linear, its first derivative depends on the value of y_i^o. This shows that initial conditions are key to the propagation of shocks whenever the underlying dynamics are non-linear. The impact of the nature of shocks, be it their magnitude, direction or both, materializes conditional on the regime, y_i^o, in which the shocks hits. Put differently, initial conditions serve as an amplification (or attenuation) mechanism of shocks. As such, regime-dependencies matter *at the first order*. As a consequence, initial conditions govern to what extent shocks of different size or direction generate non-linearities. It will be shown below that non-linear VARs are appropriate tools to capture this conditionality.

2.3   Econometric approach: Threshold VARs

The present paper uses a non-linear VAR approach to investigate asymmetric reactions to country premium shocks. A threshold VAR is a non-linear multivariate system of equations that models non-linearity additively and can consequently be estimated by OLS. TVARs condition on the initial environment and approximate the non-linear DGP by

several regime-dependent DGPs which are by themselves linear. Each regime is defined by boundaries (equal to certain values of the threshold variable) and coefficients of the VAR system are specific to each regime. The system of equations to be estimated for the reduced-form VAR with one threshold is

$$Y_t = C_1 + \phi_1(L)Y_t + (C_2 + \phi_2(L)Y_t)\,\mathcal{I}(y^*_{t-d} > \gamma) + \varepsilon_t$$

where $\mathcal{I}$ is an indicator function which equals one if the threshold variable y^* at lag order d (the delay parameter) is greater than the threshold γ and zero otherwise. Equivalently, heteroskedasticity can be assumed across the two regimes:[5]

$$Y_t = C_1 + \phi_1(L)Y_t + \varepsilon_{1,t} + (C_2 + \phi_2(L)Y_t + \varepsilon_{2,t})\,\mathcal{I}(y^*_{t-d} > \gamma)$$

The above TVAR models can be classified as a special case of more general regime-switching models such as Markov-switching VARs (MSVARs). Regime-switching models often impose *exogenous switches*. Following the work of Hamilton (1989), Markov-switching models have been applied mostly to distinguish between periods of recessions and booms. However, the assumption that the latent state is exogenous is quite unrealistic in a business cycle context where endogenous movements can be expected to lead to regime-switches. Models with endogenous switching might thus be more appropriate to capture non-linear dynamics if, as often is the case, regimes are associated with recurring dichotomous, i.e. "good" and "bad", states of the economy.

In MSVARs, the state variable is generally not observed. MSVARs thus suffer from a lack of tractability of the underlying regime-switching process as the variable(s) which cause a regime-switch cannot be identified. In contrast, TVARs explicitly model the endogenous regime-switching process which is why they are also described as "self-exciting". TVARs can therefore be viewed as a type of MSVARs *with endogenous switching* where the probability structure is modelled simplistically. TVARs have the advantage that the regime-switching is tractable, but also require the choice of a threshold variable in order to endogenize the regime-switching. Since the focus of this paper is on the effect of pecuniary externalities for generating non-linearities, asset prices and exchange rates are respectively considered as switching variables. As mentioned above, I choose these variables as they constitute the channels via which non-linearities materialize. Prices matter for the valuation of debt and collateral and any shock that generates movements in these variables can potentially generate feedback mechanisms such as debt-deflation (asset price collapses) and exchange rate depreciation.

Regime-switching models have been largely applied to financial markets which are subject to sudden changes in the underlying statistical features of the respective time series (for an overview see Ang and Timmermann, 2012). TVARs in particular have been employed in the literature on asymmetric effects of monetary policy. This literature has

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[5]If the variance-covariance matrix of the overall residuals is used to draw shocks (either from the reduced form errors or from structural errors), homoskedasticity is assumed. This is however a huge assumption in the case of regime-switching models and complicates the identification of the structural interrelation of shocks in multivariate systems. A threshold model would - more or less by construction - imply that errors and thus the contemporaneous relations of shocks behave differently in the two regimes. It is thus possible to estimate the contemporaneous relations for each regime separately.

mainly focused on the domestic context and has been applied to the US. This question was first addressed empirically by Cover (1992) and has been investigated by the use of non-linear VARs by several authors (Weise, 1999; Balke, 2000; Atanasova, 2003). Lately, TVARs or smooth transition VARs have been used to assess the effect of fiscal policy in times of crisis in contrast to tranquil times (see for example Auerbach and Gorodnichenko, 2012; Fazzari *et al.*, 2015; Baum *et al.*, 2012). In contrast to existing studies that use threshold (or smooth transition) VARs, I specifically investigate to what extent different shocks lead to regime-switching and how the switches can be associated with specific economic mechanisms that are referred to in the theoretical literature. I also provide a simple algorithm for the computation of confidence bands for non-linear IRFs which, to my knowledge, are absent in other papers using non-linear IRFs.

2.4 Threshold VARs as methodological counterparts to theoretical models

TVARs can be understood as the empirical counterpart to many non-linearities found in theoretical models which are subject to threshold effects. *First* and probably most relevant to the literature this paper is referring to, thresholds can capture the concept of occasionally binding constraints such as margin or collateral constraints. For example, similar to Aiyagari and Gertler (1999), leverage is constrained by the fraction κ: $b_{t+1} \leq \kappa q_t a_t$. If asset prices q_t fall sufficiently, the market value of assets $q_t a_t$ is lowered, the constraint binds and borrowing b_{t+1} is restricted. In this context, a TVAR with one threshold would correspond to a system where the two regimes respond to episodes of binding and non-binding constraints and the threshold marks the maximum leverage ratio κ.

Second, models often include non-linear functional forms which are characterized by threshold effects. For instance, Schmitt-Grohé and Uribe (2003) specify the functional form for a debt-elastic risk premium as follows: $R_{t+1}^h - R_{t+1} = \beta(e^{d-\bar{d}} - 1)$. If d is above a certain value $\bar{d}$, the risk premium rises (exponentially). Towards the lower end, it is however bounded by $-\beta$. The value $\bar{d}$ thus lends itself naturally as threshold above and below which a multivariate system can be approximated by two distinct linear systems.

As a *third* example, TVARs could be interpreted as the methodological pendant to models where the impact of certain variables on others is ambiguous. The literature on balance sheet effects is just one among many examples. The effect of exchange rate devaluations can be either expansionary or contractionary depending on initial conditions (Céspedes *et al.*, 2002). The threshold thus marks the point where the detrimental effect outweighs the beneficial one.

All these theoretical models and their non-linear functional forms have in common that initial conditions matter for the propagation of shocks. By definition, TVARs condition on initial settings and can thus capture these regime-dependencies. The effect of asymmetries (due to shocks of different directions) or disproportionality (due to shocks of different magnitude) can only materialize if the respective shock induces regime-switching. Initial conditions are therefore key to generating non-linearities.

2.5 Country risk shocks and endogenous regime-switching

Shocks to the country risk premium of EMEs are to a large degree driven by external factors due to systemic effects and financial contagion. An increase of the country risk premium obviously has adverse effects. This paper tackles to what extent the effects of these country risk shocks are characterized by non-linearities which are triggered by asset price or depreciation spirals (or both). The theoretical motivation of this question is grounded in models which rely on occasionally binding collateral constraints to generate asymmetric business cycle responses to various exogenous shocks.

The basic idea behind these non-linearities can be captured by a credit constraint that is both a function of asset prices as well as the level of the exchange rate. The credit constraint takes the following form:

$$\frac{r_t^* d_{t-1}(s_t)}{q_t a_t} \leq \kappa$$

where r_t^* denotes the interest rate to be paid on foreign currency denominated debt d_t which is valued in domestic terms by the exchange rate s_t. In addition, liability dollarization implies that d_t depends negatively on the exchange rate s_t (an increase denotes appreciation). Debt is issued in order to finance investment and working capital. The leverage ratio is limited from above by the margin requirement κ.

If the collateral constraint is binding, its multiplier μ_t is positive; thus:

$$\mu_t = \begin{cases} 0 & \text{if } r_t^* d_t(s_t) \leq \kappa q_t a_t \\ > 0 & \text{if } r_t^* d_t(s_t) > \kappa q_t a_t \end{cases}$$

The methodology of this paper consists in approximating these two regimes by two separate linear systems. As shown by Mendoza (2010), a binding collateral constraint has feedback effects on the external finance premium which rises with the value of the multiplier. Several channels can lead to a binding constraint and thus regime-switching. Most obviously, a rise in the foreign interest rate increases the leverage ratio. More interestingly though, asset price collapses lower the market value of collateral and can thus tighten the constraint. On the other hand, the real value of debt increases as a function of exchange rate depreciation: exchange rate movements can thus contribute to an increase of the debt burden.

If one can identify two regimes which correspond roughly to regimes of binding and non-binding constraints, non-linear effects should be detectable via impulse response analysis in TVARs. Respectively, I denote these regime with R1 and R2 where R1 is the regime in which asset prices are low or exchange rates weak (depreciation regime) and R2 corresponds to the state of the economy where asset prices are high or, respectively, exchange rates are in a state of appreciation with respect to their long-term average.

Regime-dependencies.   A positive (detrimental) country risk premium shock should have a larger impact when the economy is initially in a regime where credit constraints are non-binding and when the shock induces switching. This is the case when asset prices are high and the exchange rate is strong vis-à-vis other currencies (R2). If a (large) detrimental shock leads to lower asset prices or exchange rate depreciation, regime-switching can occur

and the credit constraint suddenly binds. This leads to a stronger reaction of real economic activity compared to a situation where asset prices are already low or exchange rates are already depreciated (R1).

Size of shocks. A similar logic applies when considering the probability of regime-switching of differently sized shocks. A priori, a large positive shock to te country risk premium which hits in the high asset price or exchange rate regime (R2) should be more likely to lead to regime-switching than a small shock.

Direction of shocks. Even more evident than the size of shocks, one can predict the probability of regime-switching when shocks are negative or positive. For instance, in R2, a positive (detrimental) shock increases the probability that a credit constraint binds whereas a negative (beneficial) shock only relaxes the constraint even more.

One can summarize the above discussion in the following hypotheses which will be analyzed in the empirical analysis below:

1. Positive shocks have a larger impact when the economy is initially in *R2* rather than in *R1*.

2. When the economy is initially in R2, *large* positive shocks should have a disproportionally larger impact than *small* positive shocks.

3. When the economy is initially in R2, *positive* shocks should have a greater impact than *negative* shocks.

3 Analysis and results

The dataset used for this paper comprises 16 emerging market economies. Table A1 lists the countries and the coverage of the data. The selection of the respective countries is solely based on data availability. Displaying impulse responses for every country would be too exhaustive. This section therefore illustrates the analysis that was conducted for each of the countries for the case of Argentina. Results across countries differ to a large degree in terms of magnitudes, but the case of Argentina is quite representative of overall qualitative patterns.

3.1 Set-up and standard VAR

The following variables are included in the VAR: interest rate spread (country risk premium), production, asset prices and the real exchange rate. I use the EMBI Global as a proxy for interest rate spreads in the private sector. This is an empirical standard in the literature on external interest rate shocks. Mendoza and Yue (2012) show that there is a high correlation of sovereign interest rates and firms' borrowing costs. The VAR system also includes the industrial production index (or respectively the manufacturing production index) as a measure of output and the MSCI Emerging Market indices in order to proxy asset price developments. Finally, the real effective exchange rate index is included in the VAR (an increase denotes appreciation).

Regarding the identification of shocks, I employ a Cholesky decomposition. Interest rate spreads are ordered first as their behaviour is partly exogenous:[6] in times of market distress in industrialized countries, EME spreads move simultaneously despite it being improbable that all EMEs' fundamentals deteriorate at the same time. Calvo *et al.* (2006) show that EME crises are characterized by sudden stops that are temporally clustered which suggests that external factors in international capital markets are important crisis triggers. Systemic effects such as contagion are at the root of *shocks* to EME spreads. However, this does not imply that EME spreads are exclusively determined by external factors. Purely exogenous shocks to EME spreads can be amplified via the feedback reaction of spreads to deteriorating fundamentals.

Data are at monthly frequency (for the time span 1994M1 to 2011M3 where available) and in log-deviations from a linear trend. A TVAR with trending variables cannot be estimated in levels as stationary variables are needed in order to identify a threshold. Table A2 in the appendix describes the data sources.

Figure 1 shows the results of a shock to the country spread for a standard VAR model: Asset prices and production fall and the real exchange rate depreciates. The channels for these responses are well-known. An increase in the country spread leads to higher borrowing costs and perceived lower future profits are reflected in falling asset prices. In addition, agents discount the future more when borrowing costs rise which also leads to decreasing asset prices. If firms are dependent on external finance (due to working capital constraints for example), output contracts as a reaction to higher borrowing costs. A depreciating exchange rate mirrors these developments.

3.2   Threshold estimation and non-linearity tests

The threshold of the switching variable of a TVAR can either be fixed by the econometrician beforehand or estimated from the model. In this paper, the threshold is determined endogenously by a grid search over possible values of the threshold variable where the grid is trimmed at a lower and upper bound in order to ensure a sufficient number of data points for the estimation in both regimes.[7] For the present analysis, the data were trimmed in order to assure a minimum of 70 data points in each regime. From the grid, the estimated threshold value corresponds to the model with the smallest determinant of the variance-covariance matrix of the estimated residuals (which is the multivariate equivalent to a sum of squared residuals criterion in a univariate model):

$$\gamma^* = \underset{\gamma}{\mathrm{argmin}} \log |\Omega_\varepsilon (\gamma)|$$

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[6]However, ordering the interest rate spread second after the production index does not change the results. The latter identification strategy assumes that only fast-moving variables like asset prices and the exchange rate are reacting to external interest rate shocks on impact. In fact, the results in figure 1 show that such an assumption is not as improbable as production does not react on impact though it is not restricted to do so.

[7]The level of trimming is chosen arbitrarily by the econometrician. No general guideline exists though 15% is very often used in the literature. The higher the number of data points, the less one is restricted in choosing extreme trimming values.

I test whether the chosen thresholds are actually meaningful by employing non-linearity tests for each equation of the VAR system. The null hypothesis that the coefficients of C_2 and $\phi_2(L)$ equal zero can be implemented by a Wald test. However, standard inference cannot be applied as the unknown threshold is not identified under the null (Hansen, 1996). One therefore has to apply a sup-Wald to be able to evaluate the statistical relevance of the endogenously chosen thresholds. Let W^* be the sup-Wald statistic of all possible statistics over the grid:

$$W^* = \sup_{\gamma} W(\gamma)$$

The distribution of this statistic does not follow a χ^2 distribution since γ is not identified. The bootstrap procedure of Hansen (1996, 1997) is therefore used to generate an empirical distribution for the sup-Wald statistic from which asymptotic p-values can be derived.

Non-linearity is tested for each equation of the VAR system. Table 1 displays the results from these non-linearity tests for the two threshold variables: asset prices and the real exchange rate. The results show that all countries except Poland display threshold effects in at least one of the equations of the reduced form VAR system. The non-linearity tests thus show that both asset prices and the real exchange rate are appropriate threshold variables and that these thresholds are meaningfully separating two regimes with statistically different dynamics. I interpret the results in table 1 as a first evidence of the presence of non-linearities, in particular regime dependencies.

3.3   Endogenous regime-switching

Computing IRFs for each regime in a standard fashion assumes that the economy stays within the respective regime which was in place when the shock initially hit. However, assuming such "intra-regime" IRFs rules out endogenous regime-switching which might take place following a shock. These IRFs can thus not be interpreted as being representative forecasts as they rely on the artificial assumption that shocks cannot cause regime changes.

It is for this reason that I analyze if endogenous regime-switching actually exists in the data. If a shock to the system generates movements in the threshold variable, its reaction then induces regime-switching over the forecast horizon. The probability of such a regime can be estimated from the model by specifying initial conditions and the characteristics of the shock. I specifically want to investigate which type of reduced-form innovation leads to regime-switching. For the non-linear analysis, 0.5 and 1.5 standard deviation innovations are considered, thus assuming a relatively small shock and a fairly large shock. The effects of both positive and negative shocks are analyzed.

Figure 4 displays the "empirical probabilities" of a regime-switch for both threshold variables. Following a reduced-form innovation to each of the variables, IRFs of the threshold variable are calculated for each data point in the initial regime (see the below discussion on generalized impulse responses for calculation details). The probability of regime-switching is derived from the number of times the switching variable crossed its threshold and is plotted against the forecast horizon at which the regime-switching occurs.

In particular, figure 4 plots the mean probabilities across the countries in the sample.[8] Though, technically speaking, the calculation of these probabilities is only based on reduced-form innovations instead of structurally identified shocks, the results can nevertheless reveal the dynamics that govern a TVAR and endogenous regime-switching. Figure 4 (a) displays the probabilities of regime-switching when asset prices are used as the threshold variable. The upper panel corresponds to a regime-switch from the low asset price regime (R1) to the high asset price regime (R2) and the lower panel shows the probabilities for switching to R1 when initially being in R2. Obviously, reduced-form innovations of asset prices lead to regime-switches. More importantly though, innovations to the country spread can cause regime-switching whereas this is not the case for reduced-form innovations of the exchange rate or production.

A similar pattern is displayed in figure 4 (b). Besides the trivial case of exchange rate shocks leading to regime-switching when the exchange rate is used as threshold variable, shocks to the interest rate spread can entail endogenous regime-switching, but this is less the case for shocks to the other variables in the TVAR. The empirical probabilities displayed in figure 4 thus show that country premium shocks are likely to lead to non-linear effects by endogenously generating movements in the respective threshold variable. Comparing the empirical probabilities of regime-switching following an interest rate shock for the two threshold variable specifications, namely asset prices and exchange rates, figure 4 shows that shocks are more likely to lead to regime switches when asset prices define the regimes. I interpret this as a first indication of a more powerful non-linearity channel operating via asset price collapses rather than depreciation spirals. The results also imply that "intra-regime" IRFs, i.e. IRF analysis that assumes that the threshold variable does not cross the threshold in reaction to a shock, are inappropriate for the analysis of country risk premium shocks.

3.4 Generalized IRFs

In order to allow for the possibility of endogenous regime-switching, non-linear impulse responses have to be calculated. The computation of impulse responses in non-linear VAR systems is more complicated (and considerably more time-consuming) than in standard linear VARs. In the linear case, the response to a shock is computed under the assumption that a shock only hits the economy at a particular point in time but neither before nor during the forecasting horizon. Linear VARs are thus history-independent and reactions to shocks are strictly proportional to the shock itself. In contrast, threshold VARs rely on the system being in one of the two regimes. Impulse responses for threshold VARs are thus history-dependent.

This history-dependence necessitates the computation of generalized impulse response functions (GIRFs) as developed by Koop *et al.* (1996). The approach relies on the simulation of data depending on which regime the system is in at the time the shock hits the economy (the history Ω_{t-1} up to point t). The advantage of GIRFs is not only that it allows for the analysis of regime-dependent responses, but also that effects of shocks of different sizes and directions can be analyzed. Due to this history- and shock-

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[8]Probabilities were calculated for each country and subsequently averaged.

dependence, GIRFs lend themselves as an appropriate framework to analyze the above mentioned dimensions of non-linearity such as regime-dependencies, asymmetries (positive vs. negative shocks) and shock non-linearity (small vs. large shocks).

A GIRF - similar to the notation of Koop *et al.* (1996) - is defined as:

$$GIRF_y\left(h, \Omega_{t-1}, u_t\right) = E\left[y_{t+h} \mid \Omega_{t-1}, u_t\right] - E\left[y_{t+h} \mid \Omega_{t-1}\right]$$

The response of variable y at horizon h can thus be calculated by simulating the evolution of the VAR system conditional on a certain history Ω_{t-1} following the shock u_t and subtracting the evolution of y conditional on the same history Ω_{t-1} without having imposed u_t at time t.[9] In order to infer on the reaction of the system conditional on being above or below the threshold, the simulations are repeated for several histories which correspond to the respective regime. The appendix provides details on the algorithm used for the simulations.

Figure 5 shows the non-linear IRFs with confidence bands. For the computation of confidence bands, I adapt a bootstrap procedure to the case of GIRFs. I describe the algorithm which I employ for this procedure in the appendix. Since non-linear IRFs allow for endogenous regime-switching, the samples corresponding to the two regimes are not independent from one another. Therefore, overlapping confidence bands cannot be interpreted as an indication that the two IRFs in question are not statistically different from each other. To evaluate the statistical difference of two non-linear IRFs, one would have to know the covariance of the two responses. The results of the non-linearity tests discussed above (table 1), however, implied that many equations of the VAR system are characterized by non-linearity.

In order to make results comparable across the different types of shocks, the IRFs were all scaled to correspond to equally sized, positive spread shocks.[10] Figure 5 (a) shows the non-linear impulse responses for the TVAR system where asset prices are used as threshold variable. Most notably, responses are more pronounced when the shocks hits the economy when it is initially in a high asset price regime (R2). Similarly, as expected from the hypotheses that were described in section 3.5, larger shocks have a disproportionally greater impact on production than small shocks. The same holds for the comparison of positive and negative shocks though differences here are smaller. Figure 5 (b) describes the responses of the variables where the exchange rate serves as threshold variable. Here, differences between the responses of production are not very pronounced.

Visual analysis of the impulse responses in figure 5 might imply that the differences between the IRFs are not very pronounced. However, this is due to the fact that I also plot the confidence bands around the response functions. When only plotting the IRFs, differences are much more evident. (As mentioned above, overlapping confidence bands

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[9]This does however not mean that other random shocks do not hit the VAR system before and after the shock. On the contrary, the GIRF approach relies on the simulation of the system under various sequences of shocks. Averaging the conditional means of the generated GIRFs equals out the shocks with which the simulations were generated. The result is thus the response of the system with history Ω_{t-1} conditional on the shock u_t only.

[10]This means that IRFs from 0.5 standard deviation shocks were multiplied by 2 and the ones from 1.5 standard deviation shocks were divided by 1.5 to yield comparable IRFs corresponding to 1 standard deviation shocks. In addition, IRFs from negative shocks were multiplied by -1.

do not imply that non-linearities are not present.) The differences across IRFs therefore materialize in the amplification coefficients that I compute in the next section.

3.5 Amplification coefficients

Visual analysis of non-linear impulse responses is cumbersome with a total of eight IRFs for each of the 16 countries and each of the four variables (this amounts to 512 IRFs in total). I therefore compute amplification coefficients in order to be able to compare two IRFs which differ with regards to magnitude, direction and/or initial regime. Amplification coefficients are calculated as follows. Let

$$\phi_k = \sum_m^h \sum_{j=1}^m IRF_{t+j,k}/(2m+1)$$

be the cumulative sum up to horizon h of the IRF of type k. A "type" corresponds to a certain choice of initial conditions, size of the shock as well as the direction of the shock. Cumulative sums instead of point-wise values are used in order to make the analysis less sensitive to outliers. The value p is chosen to be 12, but results are robust to different horizons. Using moving averages instead of cumulative sums also yields similar results. The amplification coefficient between the IRF of type k and the one of type l is then defined as the relative change between the "impact coefficients" ϕ_k and ϕ_l:

$$\Phi_{k,l} = \frac{\phi_k - \phi_l}{|\phi_l|} \times 100$$

These amplification coefficients are calculated for each country and various combinations of IRFs of different types k and l. Since some of the amplification coefficients turned out to be very large (for example whenever ϕ_l is very small), I use the power transformation $\Phi_{k,l}^* = \Phi_{k,l}^{(1/3)}$ for the cross-country analysis. With eight different types of IRFs, a total of 28 comparisons could be investigated, but this would definitely go beyond the scope of what is feasible and interesting. In contrast, I focus on specific comparisons in order to investigate the different types of non-linearity: (1) regime differences, (2) asymmetry (positive vs. negative shocks) and (3) non-proportionality (large vs. small shocks).

The results are listed in table 2. Before computing ϕ_k and ϕ_l, all non-linear IRFs were rescaled to positive, one standard deviation shocks and normalized by the country-specific size of the shock in order to compare reactions across countries. An amplification coefficient is negative if the impact coefficient ϕ_k has a stronger negative value than ϕ_l. Since I simulate the IRFs in response to a positive shock, the expected responses are negative. Amplification coefficients are therefore also expected to be negative. I compute amplification coefficients for the reaction of production (real activity) as well as the respective threshold variable.

Table 2 displays the amplification coefficients for the response of production. In order to visualize the results, the very same amplification coefficients are also displayed in figure 6 as bar charts. At first sight, there is obviously a considerable amount of heterogeneity that characterizes the responses. For most countries, amplification coefficients are negative implying effects of shocks to the interest rate spread are more contractionary in the high asset price regime (R2) which is consistent with the hypotheses described in section 2.5.

Such a mechanism is consistent with the finding of Perri and Quadrini (2018) who show in a theoretical model that output collapses are stronger the more initial conditions can be associated with long periods of credit expansion. In a similar vein, Gourinchas and Obstfeld (2012) show empirically that the degree of recession in a financial crisis can be associated with the degree of credit expansion and leverage in pre-crisis periods. With regards to the analysis of asymmetries (direction of shocks) and non-linearities (magnitude of shocks), most of the amplification coefficients are negative as expected though the results show as well a large degree of heterogeneity across countries.

Which of the two channels of non-linearity produces higher amplification coefficients (in absolute terms)? As the most important aspect of non-linearity is the one concerning regimes differences, I am concentrating on discussing the amplification coefficients which correspond to the comparisons of R2 and R1.[11] Despite the large amount of heterogeneity, the amplification coefficients show that, on average, there are higher contractions of output following shocks that originate in R2 rather than R1 when asset prices define the threshold of the TVAR system. One could therefore conclude that a "debt-deflation mechanism" is a more important trigger of regime switches than are pecuniary externalities stemming from depreciation spirals. This finding is in line with the empirical probabilities displayed in figure 4.

The main result (or puzzle) of this section remains that the amplification coefficients display a substantial amount of heterogeneity across countries. Since asymmetry and non-linearity is generated in TVARs due to regime-switches, the patterns found in the *regime comparisons* carry over to the ones applying to comparisons of shocks of different magnitudes and directions: Table 2 shows that the sign of the amplification coefficients found for regime differences is often the same as the one calculated for asymmetries and non-linearities. The heterogeneity of the amplification coefficients across countries and their potential will be discussed in the following.

4   Investigation of cross-country heterogeneity

In order to investigate the heterogeneity of responses and the validity of the underlying framework that motivates the present analysis, I further explore the presence of the mechanisms behind occasionally binding constraints and asymmetries. As described above, the literature has analyzed various mechanisms that are characterized by occasionally binding constraints and financial frictions (Mendoza, 2010; Bianchi, 2011; Benigno *et al.*, 2011). Pecuniary externalities such as exchange rate depreciation lead to the asymmetric enforcement of credit constraints when debt is denominated in foreign currency while it is collateral constraints that bind due to deleveraging when adverse shocks lead to asset price collapses. Whereas the first asymmetry is driven by liability dollarization, the second arises in the presence of high leverage. One should thus expect higher degrees of non-linearity in economies that are vulnerable to exchange rate depreciations and/or are characterized by a high degree of leverage.

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[11]Technically speaking, the results of the comparison of large and small shocks as well as positive and negative shocks depend not only on the initial conditions, but also on the choice of magnitudes and sizes and can therefore change accordingly.

In simple scatter plots, I analyze various variables with regards to their capacity in explaining the differences in the magnitudes of amplification coefficients across countries: (1) financial integration, (2) external leverage, (3) dollarization and (4) the efficiency of debt enforcement. Financial integration is calculated as the sum of total asset and liabilities over GDP. I use this variable to analyze to what extent non-linearities could be driven by external forces, i.e. the built-up of vulnerabilities by high inflows of foreign capital that trigger credit booms. Similar effects should also be captured by the second variable that is analyzed. External leverage is defined as in Gourinchas and Obstfeld (2012) and is a measure of the country's external indebtedness vis-à-vis its assets (comparable to private agent's balance sheets). External leverage is an indicator of a country's sensitivity to asset price collapses as a result of extended credit expansions (see, among others, Perri and Quadrini, 2018). For both financial integration and external leverage, the expected relationship between these indicators and the amplification coefficients is negative implying that higher vulnerability is associated with more disproportionate reactions, i.e. the contraction of output in response to a country risk premium shock is larger in the more vulnerable regime or for a more detrimental shock.

Total liabilities in foreign currency as % of GDP is used as an indicator of liability dollarization. As mentioned above several times, higher liability dollarization should lead to more vulnerability to exchange rate depreciation spirals. The expected relationship between their indicator and the amplification coefficients is therefore also negative. The efficiency of debt enforcement is an indicator collected by Djankov *et al.* (2008) and is a measure of institutional quality and investor protection. Inefficient debt enforcement can be a cause for phenomena such as debt intolerance (Reinhart *et al.*, 2003) and the degree of capital outflows during sudden stops as investors fear a loss of their assets. Therefore, the non-linear reaction to adverse shocks can be expected to be more pronounced in an environment of low institutional quality. As this measure is increasing in investor protection, the expected correlation between the efficiency of debt enforcement and the amplification coefficients is positive: the better the institutional quality of a country, the less non-linearities should be present. Table A3 in the appendix describes the datasets used for these variables which capture the different degree to which economies are characterized by financial frictions.

4.1 Threshold variable: Asset prices

Figures 7-8 show the results for the correlations of amplification coefficients and the above structural variables respectively for the response of output and the threshold variable which is in this case asset prices. It emerges from figure 7 that the greater amplification resulting from initial conditions corresponding to the high asset price regime rather than the low asset price regime is related to structural variables. If an economy is characterized by high degrees of external leverage and dollarization and when it is financially integrated with world markets, the effect of interest rate shocks is higher in R2. Also, the effect is higher the less a country is characterized by efficient debt enforcement.

A stronger reaction in the high asset price regime is driven by the fact that highly valued assets are associated with periods of high leverage and credit expansion. Leverage determines if a collateral constraint is binding or not in response to an adverse shock.

Adrian and Shin (2010) show that traders' leverage is procyclical due to asset price changes directly translating into changes in net worth which justify the adjustment of balance sheets. High asset prices incentivize credit expansion and thus contribute to the built-up of vulnerability. The fact that amplification coefficients are correlated with the degree of financial frictions such as dollarization, external leverage and institutional quality provides evidence for the assumption that regime-dependencies are caused by differing degrees of binding credit constraints.

It is logical to not only look at the responses of production, but to also investigate the amplification coefficients of the responses of the threshold variable. Figure 8 relates amplification coefficients of asset price responses to structural variables. The plots show that there is a high correlation between the degree of financial frictions and the collapse of asset prices due to the economy being in the high asset price regime which is associated with high leverage and credit expansion. Figure 8 displays the main mechanisms that are at play: Following a shock to the interest rate spread, asset prices decrease and this reaction is amplified by a binding collateral constraint. The amplification is higher the more a country is characterized by financial frictions.

The amplification coefficients which measure disproportionality (size of shocks) and asymmetry (direction of shocks) can also be related to structural variables which is a corollary to the results of the regime differences. In the second row of figures 7 and 8, the amplification coefficient from a large positive shock in R2 vis-à-vis a small positive shock in R2 is plotted against structural variables. A negative amplification coefficient implies that the adverse impact of a country premium shock is stronger when the shock is large. The plots show that the amplification from a large shock is rising with the degree of financial frictions. A similar result holds for the analysis of asymmetries (third row of scatter plots): adverse shocks have a larger impact than "beneficial" shocks, the more a country is characterized by high financial integration and leverage, considerable liability dollarization and low investor protection. As a general result, the comparison of regime differences suggest a distinct propagation of shocks which differ in size and direction. This is in line with the above described concept that regime-dependencies matter at the first order when VARs are interpreted as approximations to the true DGP.

In order to assess the strength of the correlations that are displayed in the scatter plots, I run outlier-robust regressions and report their t-statistics on top of each scatter plot. With such a small sample size, these t-statistics should be evaluated with care. Nonetheless, most of the discussed correlations between structural variables and the computed amplification coefficients are statistically meaningful for the TVAR system where asset prices serve as a threshold variable.

4.2 Threshold variable: Exchange rates

Figure 9 displays the scatter plots for the amplification coefficients of the response of production in the TVAR system where the exchange rate is used as a threshold variable. The relation with structural variables is far less pronounced than in the case where asset prices serve as the threshold variable. Apart from the correlation of amplification due to regime differences and financial integration, most plots do not show a clear pattern. The fact that the results differ from the ones where another threshold variable is used is not

surprising as the distribution of regimes do not coincide. High asset price regimes are not necessarily associated with periods of exchange rate appreciation and vice versa as the plots of the regime distributions in figures 2 and 3 show.

Despite a lack of clear pattern for the response of production, the dynamics of the exchange rate remain interesting. As already discussed above, pecuniary externalities can lead to "depreciation spirals". Depreciating exchange rates increase the debt burden in the presence of liability dollarization and binding credit constraints. The upper row of figure 10, which displays the correlation between the amplification coefficients for exchange rates and the structural variables, shows that the higher amplification due to shocks hitting in an appreciation regime rather than in a depreciation regime can be associated with a higher degree of external leverage and dollarization. This result captures the idea that agents do not internalize the effects of sudden exchange rate depreciations after prolonged periods of real appreciation when their debt is denominated in foreign currency. The depreciation spiral is thus more pronounced the more a country is characterized by these financial frictions. However, the results on the reaction of production show that exchange rate spirals do not translate to a large extent into a contraction of production as the real effects of depreciating exchange rates are not as detrimental as in the case of asset price spirals. Despite the fact that the reaction of the exchange rate is characterized by non-linear effects, this does not map into non-linear reactions of output. The financial frictions associated with the non-linear reaction of exchange rates do not imply that the reaction of output should also be related to these financial frictions.

The above described amplification channel where credit constraints bind as a result of currency depreciation is closely linked to the flexibility of the exchange rate. The choice of the exchange rate regime matters with regards to the degree of "repudiation" of externalities arising from exchange rate movements. Non-linearities can be expected to be larger the less flexible exchange rates are since agents tend to not internalize the increasing probability of collapsing exchange rate pegs when adverse shocks hit the economy. The plots of the amplification coefficients of the amplification coefficients of exchange rates against exchange rate flexibility (results not shown) indicate the presence of such a mechanism: the more rigid the exchange rate regime, the stronger is the amplification effect from appreciation vs. depreciation regimes.

5   Conclusion

This paper provides an empirical validation of asymmetric responses of EME business cycles to external shocks. These non-linearities are often found in theoretical models characterized by financial frictions such as credit constraints. The methodology of TVAR estimation and non-linear impulse response analysis is therefore strongly linked to the idea that the impact of certain shocks depends on initial conditions and is amplified (or attenuated) whenever there is regime-switching. This switching is thought of as the empirical corollary to a collateral constraint that binds when the threshold variable crosses from a favourable to a non-favourable regime. Special emphasis is put on the importance of *endogenous* regime-switching induced by the reactions of variables to an external country risk premium shock.

The literature on EME business cycle patterns has identified foreign interest rate shocks as important drivers of business cycles, but has also shown that these shocks matter only due to their effect in times of crisis. In a similar vein, this paper validates this result empirically and finds that substantial non-linearity and dependence on initial conditions characterizes the business cycle responses to country risk premium shocks.

Individual country analyses reveal a large degree of heterogeneity across EMEs. However, heterogeneity can be traced back to individual country characteristics. The externalities that lead to asymmetry in models with occasionally binding constraints, "debt-deflation" and "depreciation spirals", are analyzed with respect to their capacity in explaining the large heterogeneity of degrees of non-linearity that is found across countries. The results show that these externalities are stronger the more a country is characterized by financial and structural frictions such as liability dollarization, leverage or institutional deficiencies.

As a more general result, I find that non-linearity is mainly driven by regime-dependency. The effect of shocks of different magnitude and/or direction depends largely on the initial conditions which define the vulnerability of an economy. This calls for the further use and development of non-linear time series techniques as opposed to linear techniques which evaluate dynamic systems at their means. This is especially relevant in the context of EME business cycles which are characterized by high macroeconomic volatility and are therefore very likely to be away from their steady state at most times.

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Appendix

A.   Algorithm for computation of GIRFs

1. A history Ω_{t-1}^r for all the variables is chosen depending on which regime is assumed; this means that a particular realization ω_{t-1}^r of the threshold variable is drawn randomly based upon the regime criterion. This history comprises all the lags up to order p of the variables in the VAR.

2. Shocks are drawn based on the variance-covariance matrix of the residuals. As a joint distribution of these shocks is assumed, a k-dimensional vector u_{t+h}^b is drawn at each horizon (where k denotes the number of endogenous variables in the VAR).

3. The future evolution of all variables is simulated using the estimated coefficients for both regimes as well as the shock process for $h + 1$ periods. Hence, the model is allowed to switch regimes over the forecast horizon. The resulting sequence is denoted by $Y_{t+h}(\omega_{t-1}^r, u_{t+h}^b)$.

4. Step 3 is repeated but the shock sequence at $t = 0$ is replaced by a shock of size δ_j for the variable j and the corresponding contemporaneous shocks for the other variables. This $k \times 1$ vector is denoted by u_j^*. The resulting sequence is denoted by $Y_{t+h}(\omega_{t-1}^r, u_{t+h}^b, u_j^*)$.

5. Steps 2 to 4 are repeated R times (here: $R = 300$) in order for the shocks to average out.

6. Steps 1 to 4 are repeated B times (here: $B = 300$) to obtain an average over the respective regime history and - once again - to iterate over a large number of draws of shock sequences which are expected to average out.

7. The GIRF is the difference between the simulated forecast assuming the shock u_j^* and the forecast assuming no particular shock:

$$GIRF(h, \Omega_{t-1}, u_j^*) = \left[Y_{t+h}(\omega_{t-1}^r, u_{t+h}^b, u_j^*) - Y_{t+h}(\omega_{t-1}^r, u_{t+h}^b) \right] / (B \times R)$$

B.   Confidence bands for GIRFs

Most papers which work with GIRFs do not compute confidence bands. Besides the absence of methodological work on this issue, the reason lies in the fact that the simulations of the GIRFs are already computationally quite intensive. Setting $R = B = 1000$ can take several hours - even when resorting to parallel computing to alleviate computation time. Adding bootstrap confidence bands with another 1000 simulation runs would thus be equivalent to a simulation with 10^9 repetitions. In the present paper, the number of simulation runs B and R were thus reduced for the calculation of confidence bands. My method for computing bootstrap confidence bands employs the following simple algorithm:

1. Artificial data is generated recursively using the estimated coefficients and errors from the TVAR structure.

2. Using the *recursive datatset*, the regression coefficients C_1, $\phi_1(L)$, C_2 and $\phi_2(L)$ as well as error terms are calculated from a TVAR assuming the threshold corresponds to the estimated value γ.

3. Using the *original dataset*, but the coefficients and errors from step 2, GIRFs are calculated as described in the above algorithm for each particular combination of shocks and initial conditions.

4. Steps 1 to 3 are repeated S (here: $S = 300$) times to generate a sample distribution of the GIRFs from which confidence bands are drawn at the respective significance levels.

C.   Data used for the econometric analysis

Table A1: Data coverage

Country	Coverage		Country	Coverage	
Argentina (ARG)	1994M1	2011M3	Peru (PER)	1997M3	2011M3
Brazil (BRA)	1994M4	2011M3	Philippines (PHL)	1993M12	2011M3
Chile (CHL)	1999M5	2011M3	Poland (POL)	1994M10	2011M3
China (CHN)	1994M3	2011M3	Russia (RUS)	1997M12	2011M3
Colombia (COL)	1997M2	2011M3	Thailand (THA)	1997M5	2006M2
Hungary (HUN)	1999M1	2011M3	Turkey (TUR)	1996M6	2011M3
Mexico (MEX)	1993M12	2011M3	South Africa (ZAF)	1997M3	2011M3
Malaysia (MYS)	1996M10	2011M3			

Table A2: Data sources

Variable	Data used	Source	Countries
Interest rate spread	EMBI Global	Datastream (EMBI)	All
Production	Industrial production index, s.a.	Datastream (nat'l)	ARG, BRA, CHL, HUN, PER*
		Datastream (IFS)	MYS*, POL*, RUS*, TUR*
		IFS	CHN*, PHL*
	Manufacturing production index, s.a.	Datastream (nat'l)	ZAF
		DANE[1]	COL*
		Bank of Thailand	THA
	Global economic activity indicator, s.a.	INEGI[2]	MEX
Asset prices	MSCI in LCU	Datastream (MSCI)	All
Exchange rate	Real effective exchange rate index	IFS	BRA, CHL, CHN, COL, HUN, MEX, MYS, PHL, POL, RUS, ZAF
		BIS	ARG, PER, THA, TUR

Notes: * Seasonal adjustment is performed for those variables which are not already adjusted using the United States Census Bureau's X12-ARIMA seasonal adjustment application in EViews. The production series were adjusted using samples starting as early as 1970M1 until 2011M6 the latest depending on the length of the data series. (1) DANE: Departamento Administrativo Nacional de Estadística, Colombia. (2) INEGI: Instituto Nacional de Estadística, Geografía e Informática, Mexico.

Table A3: Data used to investigate cross-country heterogeneity

Variable	Dataset	Coverage	Authors
Financial integration	"External Wealth of Nations"	1970-2007	Lane and Milesi-Ferretti (2007)
External leverage	Calculated following Gourinchas and Obstfeld (2012)	1970-2007	
Dollarization	"Fin. Exchange Rates and Int'l Currency Exposures"	1990-2005	Lane and Shambaugh (2010)
Debt enforcement	Numbers taken from table 2	2006	Djankov *et al.* (2008)

Notes: For the scatter plots, the data in this table are averaged over the sample period (where available) for each country.

Tables

Table 1: Asymptotic p-values for sup-Wald statistics of non-linearity tests

Equation	ARG	BRA	CHL	CHN	COL	HUN	KOR	MEX
Threshold variable: Asset prices								
Interest rate spread	0.868	0.032*	0.532	0.326	0.060*	0.285	0.000*	0.760
Production	0.000*	0.174	0.038*	0.098*	0.186	0.265	0.290	0.006*
Asset prices	0.000*	0.738	0.007*	0.004*	0.114	0.070*	0.000*	0.000*
Exchange rate	0.016*	0.268	0.162	0.336	0.000*	0.030*	0.003*	0.114
Threshold variable: Real exchange rate								
Interest rate spread	0.094*	0.000*	0.050*	0.002*	0.200	0.089*	0.151	0.018*
Production	0.582	0.628	0.028*	0.044*	0.004*	0.278	0.692	0.146
Asset prices	0.000*	0.332	0.018*	0.044*	0.132	0.596	0.000*	0.000*
Exchange rate	0.164	0.000*	0.126	0.290	0.344	0.163	0.278	0.000*

Equation	MYS	PER	PHL	POL	RUS	THA[1]	TUR	ZAF
Threshold variable: Asset prices								
Interest rate spread	0.516	0.008*	0.050*	0.474	0.172	0.000*	0.440	0.000*
Production	0.004*	0.084*	0.124	0.536	0.000*	0.139	0.328	0.188
Asset prices	0.000*	0.146	0.278	0.610	0.002*	0.000*	0.044*	0.044*
Exchange rate	0.324	0.826	0.588	0.148	0.224	0.000*	0.000*	0.480
Threshold variable: Real exchange rate								
Interest rate spread	0.010*	0.086*	0.000*	0.160	0.028*	0.093*	0.402	0.004*
Production	0.014*	0.150	0.314	0.808	0.008*	0.001*	0.164	0.064*
Asset prices	0.056*	0.272	0.082*	0.230	0.000*	0.048*	0.388	0.138
Exchange rate	0.006*	0.842	0.018*	0.188	0.000*	0.010*	0.116	0.058*

Notes: Non-linearity tests follow the bootstrap procedure of Hansen (1996, 1997). Stars denote significance at the 10% significance level.

Table 2: Amplification coefficients for response of production

Type of shock	ARG	BRA	CHL	CHN	COL	HUN	KOR	MEX
Threshold variable: Asset prices								
R2 vs. R1	-3.54	-2.57	-3.65	2.71	1.04	-5.23	4.51	-0.63
Large vs. small	-5.57	-2.09	-1.12	2.18	1.72	-2.02	3.97	-1.55
Negative vs. positive	-7.09	1.54	2.40	3.85	3.09	-3.29	4.55	-2.01
Threshold variable: Exchange rate								
R2 vs. R1	3.17	-3.48	3.99	-1.98	2.14	-4.66	-2.34	-1.48
Large vs. small	1.48	-3.49	-5.07	-9.57	-9.13	-2.02	-3.27	-3.41
Negative vs. positive	1.25	-4.42	-5.53	-13.76	-12.65	-3.46	-3.82	-3.92
Type of shock	**MYS**	**PER**	**PHL**	**POL**	**RUS**	**THA**	**TUR**	**ZAF**
Threshold variable: Asset prices								
R2 vs. R1	-3.26	-2.00	4.65	-2.08	-2.90	3.15	-3.64	-2.38
Large vs. small	-1.42	1.15	-4.48	1.34	-1.59	1.89	1.55	-6.73
Negative vs. positive	1.72	2.66	-4.60	1.98	1.98	2.91	-2.31	-8.42
Threshold variable: Exchange rate								
R2 vs. R1	3.33	-3.80	-3.83	1.17	2.75	3.73	1.13	4.37
Large vs. small	-3.19	-1.87	-5.57	1.92	-4.54	3.19	-3.83	2.02
Negative vs. positive	-4.35	-2.36	-5.32	1.56	-6.56	4.35	-4.68	4.13

Notes: Amplification coefficients are computed as described in section 3.5. A negative coefficient implies a higher contraction for the regime that is expected to be more vulnerable or the type of shock that is expected to have a more detrimental impact.

Table 3: Amplification coefficients for response of threshold variable

Type of shock	ARG	BRA	CHL	CHN	COL	HUN	KOR	MEX
Threshold variable: Asset prices								
R2 vs. R1	-4.27	-4.83	2.76	3.18	1.95	-7.17	1.56	1.86
Large vs. small	-2.09	1.20	1.06	2.51	2.86	-2.20	3.44	1.92
Negative vs. positive	-2.74	2.71	1.91	3.82	3.54	-3.40	4.09	2.26
Threshold variable: Exchange rate								
R2 vs. R1	-6.31	-3.58	2.98	1.40	-2.42	-4.03	2.36	-3.20
Large vs. small	-6.86	-2.99	-4.45	4.23	3.03	-4.86	-2.41	-1.96
Negative vs. positive	-10.39	-3.64	-5.13	5.76	3.95	-7.73	-1.48	-1.38
Type of shock	**MYS**	**PER**	**PHL**	**POL**	**RUS**	**THA**	**TUR**	**ZAF**
Threshold variable: Asset prices								
R2 vs. R1	-1.95	-3.74	-2.88	-2.45	-1.70	-1.42	-3.68	2.07
Large vs. small	-2.00	2.51	-1.54	0.82	-2.18	-1.33	2.12	-4.87
Negative vs. positive	-1.25	3.52	-2.43	2.27	-1.69	-2.06	1.01	-6.47
Threshold variable: Exchange rate								
R2 vs. R1	-4.69	-6.65	-10.34	2.97	-4.84	-2.73	3.11	-1.22
Large vs. small	3.04	-13.48	0.69	1.72	-3.67	-5.95	2.22	-6.25
Negative vs. positive	4.59	-19.74	-2.89	2.71	-5.56	-7.77	3.18	-7.87

Notes: Amplification coefficients are computed as described in section 3.5. A negative coefficient implies a higher contraction for the regime that is expected to be more vulnerable or the type of shock that is expected to have a more detrimental impact.
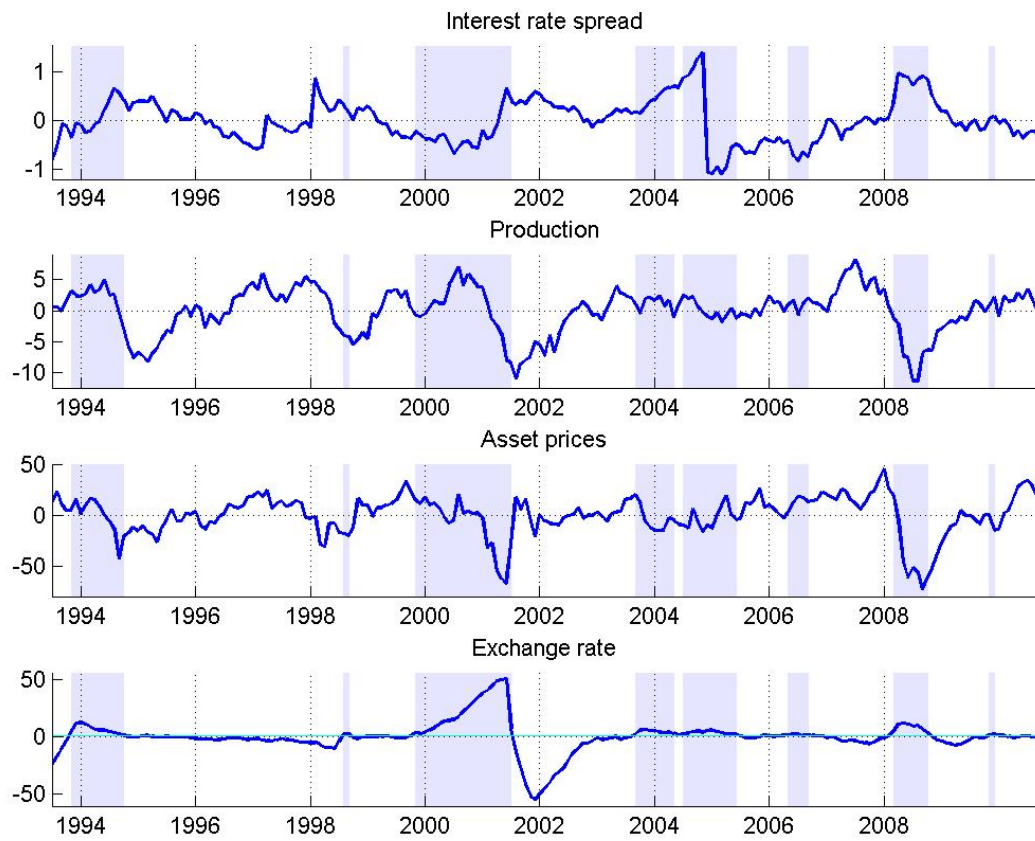
Figures

Figure 1: Standard VAR: Argentina
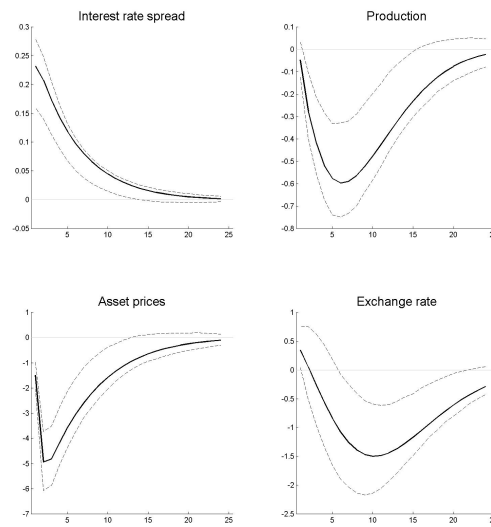
Notes: Responses to a one standard deviation shock to the EMBI spread. Dashed lines represent 90% confidence intervals.

Figure 2: Time series plots for Argentina - Threshold: Asset prices
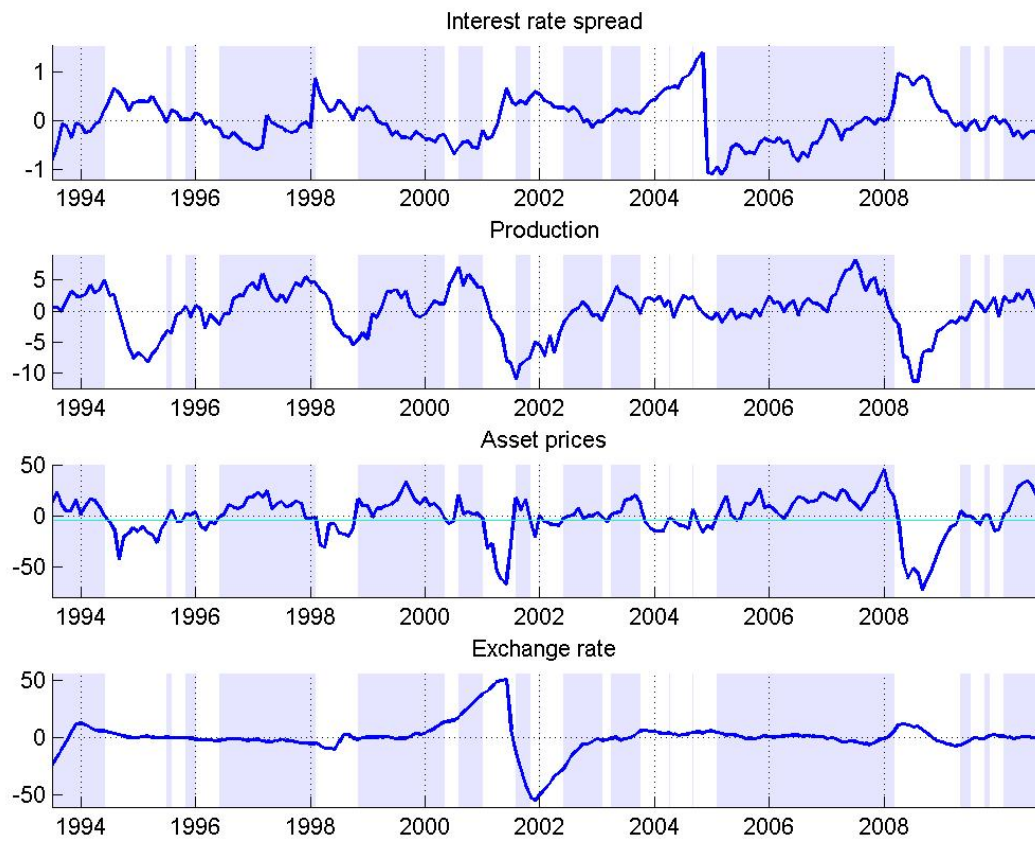
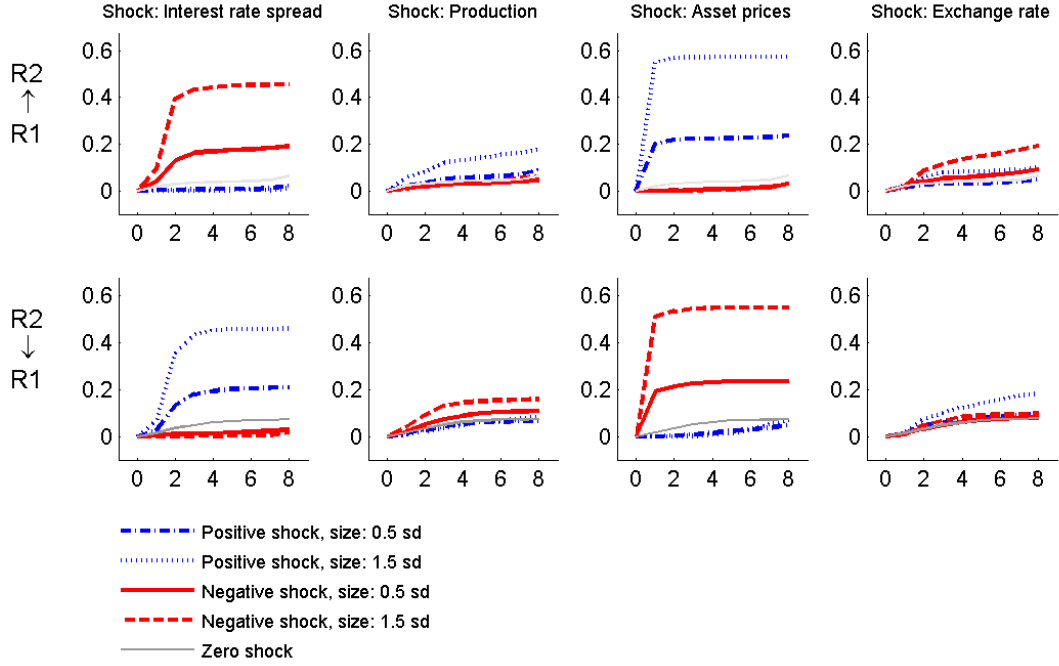*Notes*: Shaded regions correspond to R2, i.e. the high asset price regime.

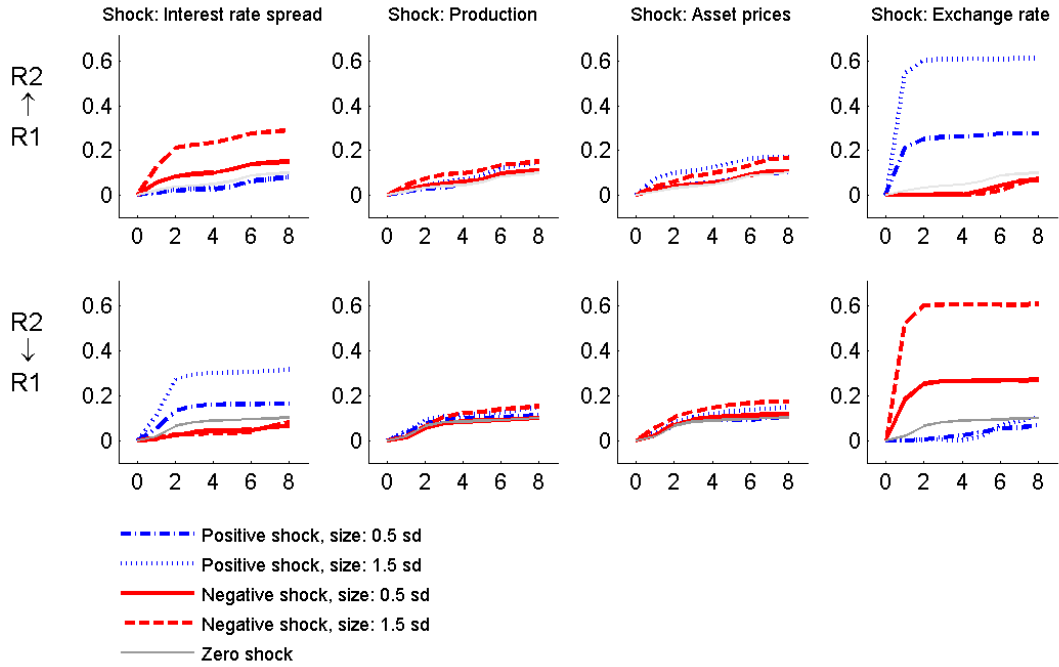Figure 3: Time series plots for Argentina - Threshold: Exchange rates

Notes: Shaded regions correspond to R2, i.e. the appreciation regime.

Figure 4: Empirical probabilities of regime-switching

(a) Threshold variable: Asset prices

(b) Threshold variable: Exchange rate

Notes: The figures display the "probability" of a regime-switch, for regimes corresponding to asset prices in panel (a) and for regimes corresponding to exchange rates in panel (b). The probability of regime-switching is derived from the number of times the switching variable crossed its threshold and is plotted against the forecast horizon at which the regime-switching occurs. In particular, following a shock in $t = 1$, the graph plots the share of occurrences of a regime switch over the forecasting horizon (in months which are displayed on the x-axis). The first row describes the probability of switching from R1 to R2 following the respective shock as shown in the column headers. The second panel shows the probabilities of switching from R2 to R1 following a shock in $t = 1$.

Figure 5: Non-linear impulse responses: Argentina

(a) Threshold variable: Asset prices
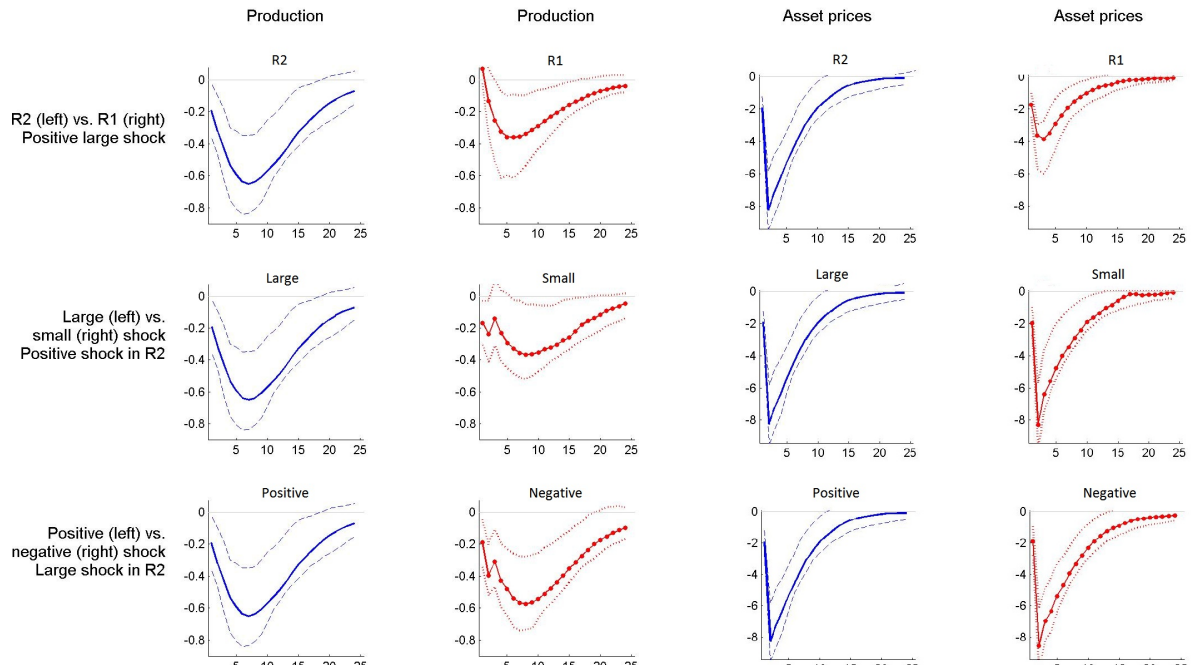
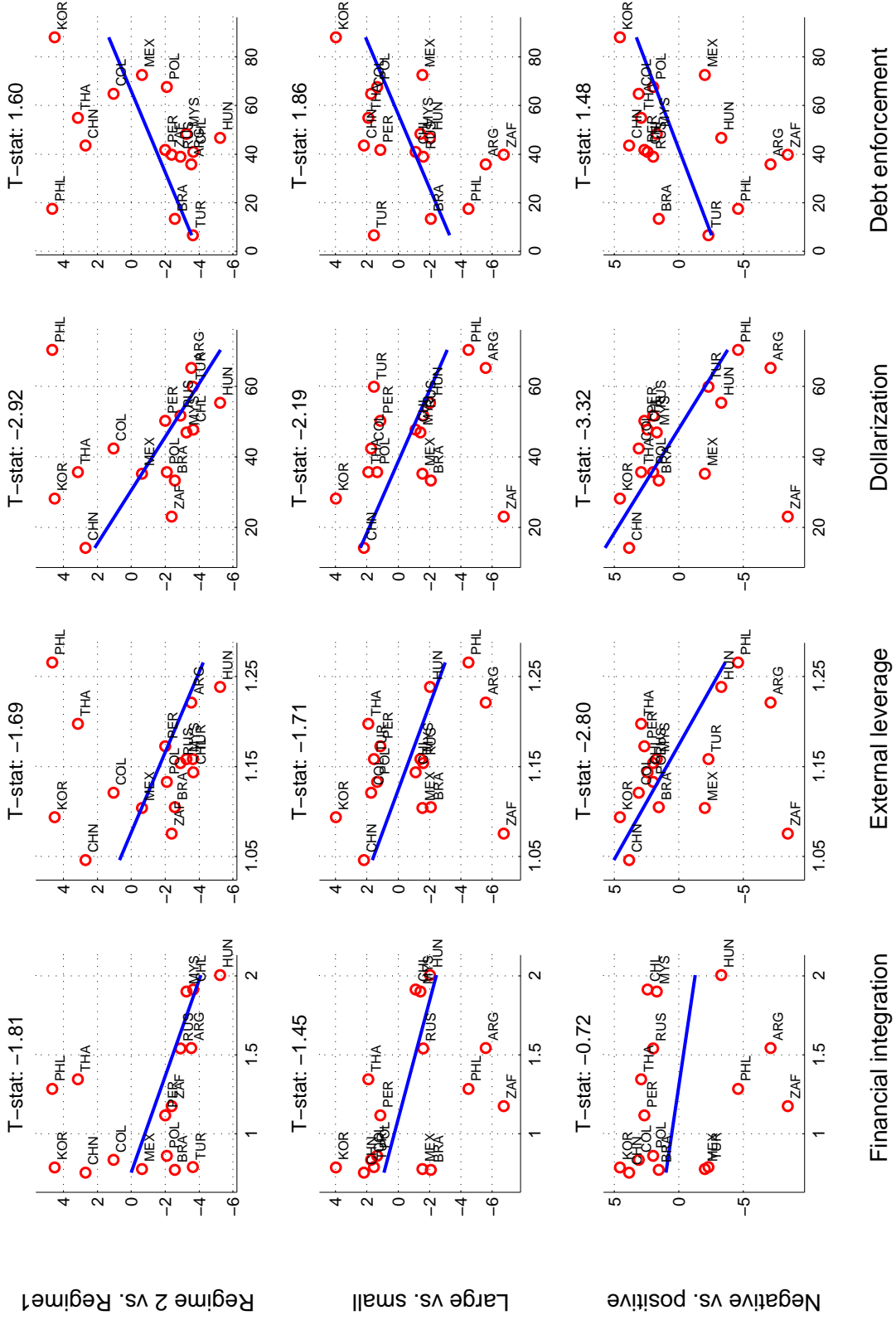## (b) Threshold variable: Exchange rate
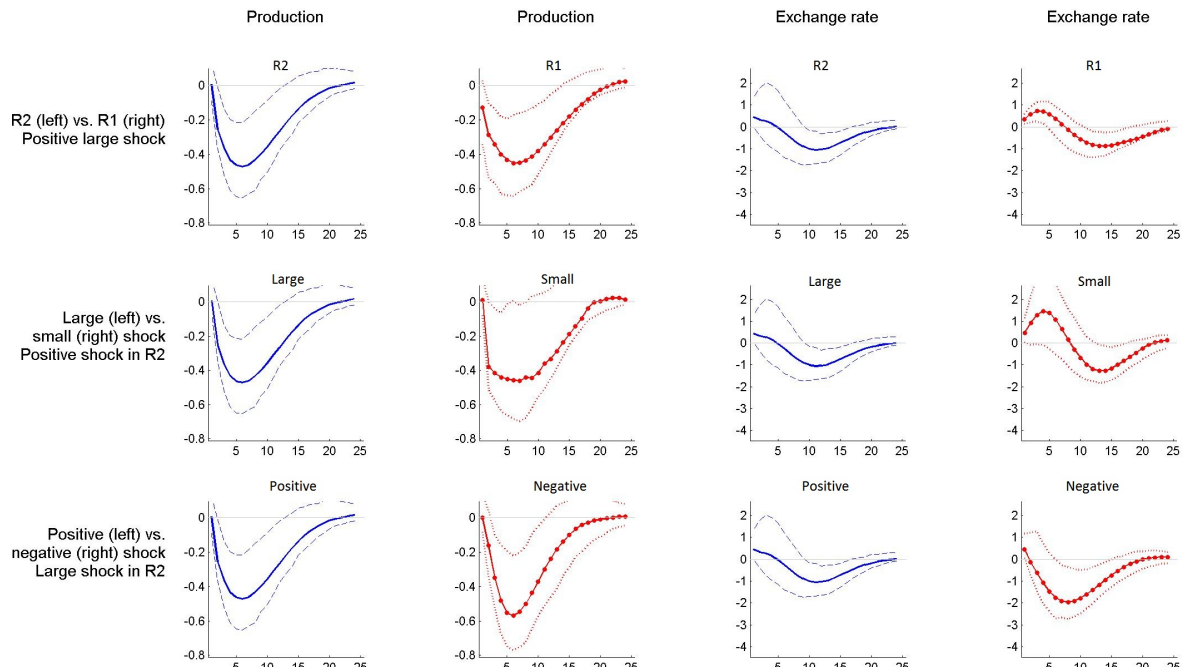
Notes: The first panel shows the impulse responses when asset prices are used as a threshold variable and the second panel shows the impulse responses when exchange rates are used as threshold variables. All IRFs are rescaled to one standard deviation, positive shocks to the country risk premium.
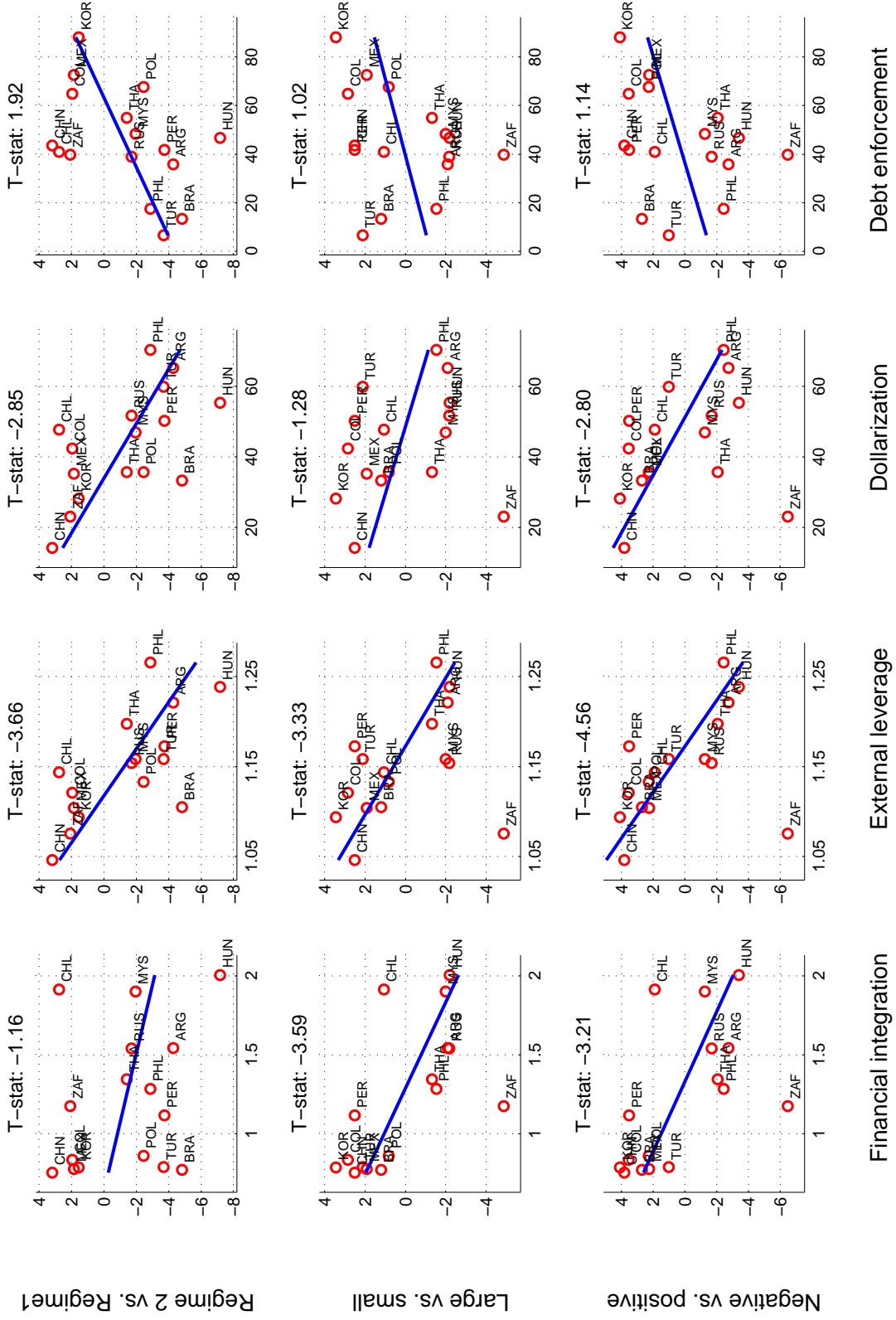
Figure 6: Amplification coefficients

(a) Regime differences

(b) Large vs. small shocks

(c) Negative vs. positive shocks
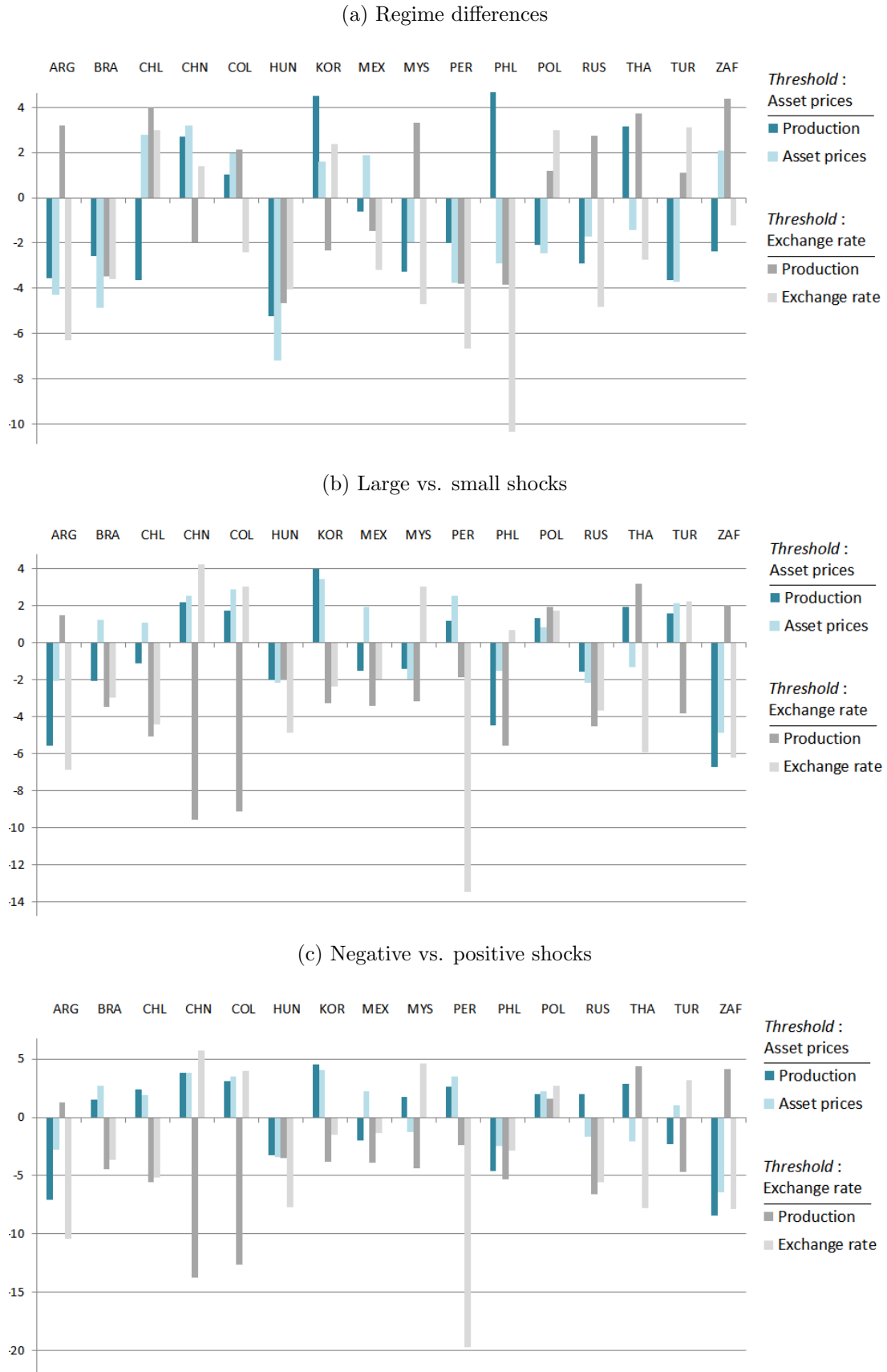
Notes: The amplification coefficients displayed in the bar charts correspond to the ones listed in tables 2 and 3. Amplification coefficients are computed as described in section 3.5. A negative coefficient implies a higher contraction for the regime that is expected to be more vulnerable or the type of shock that is expected to have a more detrimental impact.

Figure 7: Heterogeneity of amplification coefficients
Response variable: Production — Threshold variable: Asset prices

Notes: The scatter plots display the correlation patterns between structural variables (financial integration, external leverage, liability dollarization and debt enforcement) as described in section 4 and in table A3 as well as the computed amplification coefficients. The regression line in each plot and the t-statistics reported on top correspond to robust regressions of amplification coefficients on structural variables.

Figure 8: Heterogeneity of amplification coefficients

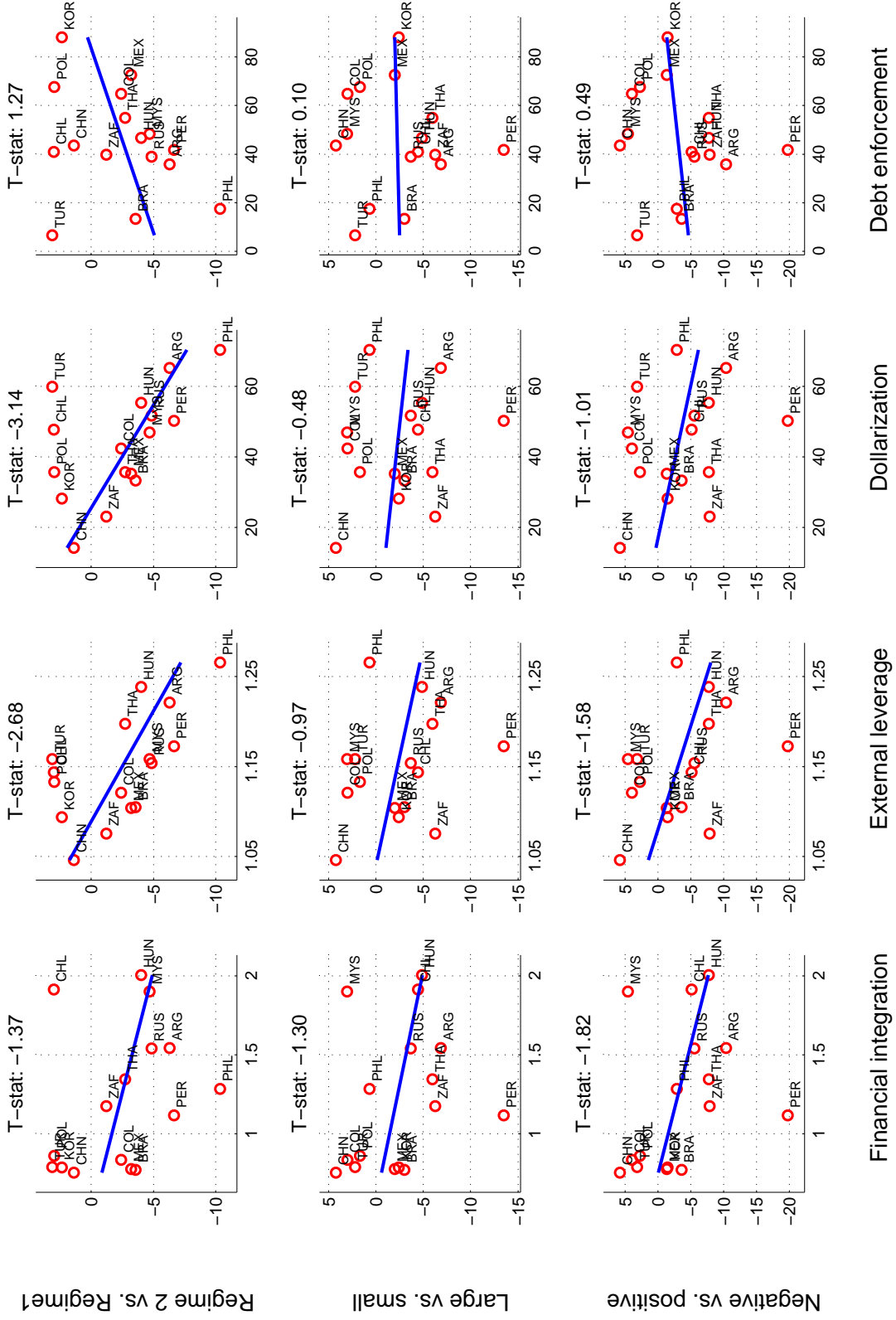Response variable: Asset prices — Threshold variable: Asset prices

Figure 9: Heterogeneity of amplification coefficients

Response variable: Production — Threshold variable: Exchange rate

Notes: The scatter plots display the correlation patterns between structural variables (financial integration, external leverage, liability dollarization and debt enforcement) as described in section 4 and in table A3 as well as the computed amplification coefficients. The regression line in each plot and the t-statistics reported on top correspond to robust regressions of amplification coefficients on structural variables.

Figure 10: Heterogeneity of amplification coefficients

Response variable: Exchange rate — Threshold variable: Exchange rate

Notes: The scatter plots display the correlation patterns between structural variables (financial integration, external leverage, liability dollarization and debt enforcement) as described in section 4 and in table A3 as well as the computed amplification coefficients. The regression line in each plot and the *t*-statistics reported on top correspond to robust regressions of amplification coefficients on structural variables.