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# **THE CREDIT CHANNEL AND MONETARY TRANSMISSION IN BRAZIL AND CHILE: A STRUCTURAL VAR APPROACH**

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## **Resumen**

Utilizamos una representación SVAR de expectativa aumentada para un modelo nuevo keynesiano de economía abierta para estudiar la transmisión monetaria en Brasil y Chile. El modelo estructural subyacente incorpora rasgos estructurales claves de economías emergentes, en particular el rol del canal del crédito bancario. Encontramos que un cambio en la tasa de interés tiene efectos más directos sobre actividad e inflación en ambos países en comparación con economías avanzadas, y que la dinámica del tipo de cambio juega un papel importante en la transmisión monetaria, ya que los movimientos monetarios tienen una alta respuesta frente a modificaciones a la tasa de interés de política. También encontramos que un shock crediticio de tamaño típico tiene grandes efectos en la actividad y la inflación en los dos países, pero más en Chile cuya penetración bancaria es mayor.

## **Abstract**

We use an expectation-augmented SVAR representation of an open economy New Keynesian model to study monetary transmission in Brazil and Chile. The underlying structural model incorporates key structural features of Emerging Market economies, notably the role of a bank-credit channel. We find that interest rate changes have swifter effects on output and inflation in both countries compared to advanced economies and that exchange rate dynamics plays an important role in monetary transmission, as currency movements are highly responsive to changes in policy-controlled interest rates. We also find the typical size of credit shocks to have large effects on output and inflation in the two economies, being stronger in Chile where bank penetration is higher.

# 1 INTRODUCTION

The widespread adoption of inflation-targeting (IT) regimes by emerging-market (EM) economies has generated considerable interest in the channels through which monetary policy shocks affect output, inflation and other relevant aggregates in such economies. Yet, there has been a paucity of empirical research for EMs relative to the large literature on advanced countries, partly reflecting shorter time series and other problems not typically faced in studies of the latter<sup>2</sup>. While there have recently been a few studies fitting DSGE models of a standard variety to emerging market data e.g. Furlani et al (2008), da Silveira (2008) and del Negro and Schorfheide (2008), one limitation with these studies is that some key structural features of EMs are largely ignored in the chosen specification. Moreover, Bayesian methods utilized for estimation in these studies often impose strong priors, so that the empirical investigation is less about "discovery" than quantifying the parameters of some prescribed model. This is not to deny that DSGE models are useful for thinking about inter-relationships in the macro economy, but our view is that they are often best used as a source of structural information that provides a skeleton to allow investigators to organize the data, rather than to be imposed upon it, at least until one is sure that the model is a good representation of the data. Often the only way the latter has been judged is by a comparison of the results to a VAR, but this is unlikely to be a very powerful test. Often simple checks, such as whether the assumptions of the model made about expectations and shocks are consistent with the data, are far more likely to show up deficiencies in the specification.

Our objective in this paper is to develop a model that uses a particular DSGE model, the New Keynesian model, as a skeleton, and to then expand this so as to resemble a Structural VAR (SVAR). Unlike existing SVARs that either force the system to be recursive ("ordered") or which impose restrictions based on the signs or long-run properties of impulse responses, we propose that the VAR be *structured* by reference to some skeletal model that has a theoretical base. After eliminating the expectations in such a model we thereby produce a non-recursive SVAR, and it is this which forms the basis of our structured VAR. By choosing the skeletal model appropriately we can make an allowance for the role of external debt accumulation, exogenous fluctuations in the terms of trade, and endogenous determinants of the external trade balance through variation in domestic absorption. As we have shown in previous work (Catão et al., 2008), the inclusion of an external debt accumulation equation in the structured VAR model is not only of interest in its own right---as it permits the tracking of

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<sup>2</sup> Even though the literature on the relative performance of IT regimes in Ems is now sizeable (see, e.g., Loayza and Soto, 2002; Fraga, Goldfajn, and Minella, 2004; Mishkin and Schmidt-Hebbel, 2007), model-based studies on the monetary transmission in these economies remain scarce. A notable exception has been the case of Chile, as discussed below.

the effects of monetary shocks through key external aggregates -- but it also imposes some stock-flow dynamics on the model that allows it to have an invertible VAR representation.

All linearized DSGE models imply that the data can be represented as a structured VAR. The shocks in the structure are identified in the DSGE approach by a combination of exclusion restrictions and the presence of some common parameters in the structural equations of the system. These serve to reduce the number of parameters to be estimated to what can be (hopefully) estimated from the data. The structured VAR we adopt retains some of the exclusion restrictions of DSGE models but attempts to be less restrictive in relation to the specification of the underlying structural equations. It also aims to eclectically introduce some of the features of EM macro economies. In particular we augment the canonical model to include a bank-dependent domestic private sector. This allows us to capture additional effects of monetary policy shocks through the bank intermediation channel emphasized by Bernanke and Blinder (1988), a channel which has been argued to be particularly relevant in emerging markets (Edwards and Vegh, 1997; Catão and Chang, 2010). The effect of shocks to banks' lending capacity---arising, from (say) exogenous changes in reserve requirements and/or banking intermediation technology---on output, inflation, and other aggregates can also be traced out in the model. Our structured VAR model is also designed to retain one of the important features emphasized in the DSGE perspective, viz. the integration of stocks and flows. This is rarely dealt with in standard SVARs.

The modeling strategy is empirically implemented on data for Brazil and Chile over 1999/1-2009/1. This sample period was chosen because the countries formally adopted an inflation target in 1999 (Brazil) and 2000 (Chile). While Chile did start targeting inflation from the early 1990s, it operated a system of exchange rate bands through 1998, so targeting inflation was not the overriding goal of monetary policy. Moreover, an advantage in restricting the estimation period from 1999 to 2009 is to use the same sample for both countries, so as to more readily facilitate comparison. Achieving a balance between retaining a large number of parameters, so as to capture the quite general dynamics that might be in the data, and the need for a relatively parsimonious specification to aid interpretation, is often more an art than a science, particularly when the sample sizes available for estimation are very short.

Given the sample size, some restriction on the VAR is needed. The model we apply to both countries represents an expansion of the methodology used in Catão et al (2008), in that we replace the recursive SVAR used there with a non-recursive SVAR. Substantial differences emerge between the conclusions reached with a traditional recursive VAR and that from this paper's approach. Despite the relatively short time span available for estimation (1999/1--2009/1), our structured VAR estimates do not generate "price puzzles", "exchange rate puzzles"

and other anomalies that abound in the literature, and which would be found for both countries if one applied a standard recursive SVAR.

The main results are as follows. First, compared with the United States and other advanced countries, the transmission mechanism works faster, with the bulk of the effects on output and inflation taking place within a year. The magnitude of monetary policy effects on inflation and output growth are much the same as in advanced economies, but the mechanism is different for inflation, with exchange rate rather than output gap effects dominating - something often found in small open economies such as Australia.

Second, the bank credit channel is found to play a non-trivial role in monetary transmission. Our results are consistent with the existence of two channels through which monetary policy affects credit and then output. One is via changes in the lending-deposit spread following shocks to the policy interest rate, amplifying the standard inter-temporal effect of monetary policy changes on absorption. Second, there is an intra-temporal effect: monetary tightening tends to appreciate the exchange rate in the short-run and this has expansionary effects on bank credit. The latter occurs when the domestic business sector tends to have a sizeable stock of foreign-currency denominated debt and/or the non-tradable sector of the economy is more "bank-dependent" than its tradable counterpart, implying that the overall demand for bank credit will tend to increase as relative prices shift towards non-tradable producers. This combination of the balance sheet effects of currency mismatches and the greater bank dependence of domestic firms imply that monetary policy will have non-trivial effects on bank lending and hence on absorption. Our estimates indicate that, while the inter-temporal channel eventually wins out, so that monetary tightening (loosening) depresses (boosts) bank credit, the intra-temporal channel appears to play an offsetting role.

Regarding the quantitative impact of credit shocks these tend to be larger for Chile. While neither is large in response to a 1% change in credit growth, the question is whether this is the right scenario given the typical size of credit shocks in emerging markets. Over the period 1999/1-2009/1 there has been very strong growth in credit in both countries with standard deviations in credit shocks of around 9% and 5% respectively. So, although the impact of a 1% change in credit upon inflation and output is relatively small, such large variations in actual credit growth might suggest that these developments have been important for macro-economic outcomes. Indeed, for Brazil the impact of a 9% (positive) shock to credit growth on inflation is roughly equivalent to a decrease of 80 basis points in the interest rate, all else constant. In the case of Chile, the inflationary impact of a positive 5% shock to credit growth is equivalent to a decrease of some 100 basis points in interest rates.

The remainder of the paper is divided into five sections. Section II reviews the existing evidence on the monetary transmission mechanism in Brazil and Chile and provides a

motivation for the model and results. Section III lays out the methodology, firstly in a general way, and then in the context of the structural model that is used as the skeleton for our SVAR. Section IV provides a discussion of the data including the construction of output and absorption gaps. Section V presents the estimation results for the structural VAR equations as well as the resulting impulse responses for money and credit shocks. The paper concludes with a brief summary and discussion of the main findings.

## **2 EXISTING EVIDENCE**

### **2.1 Brazil**

The introduction of the IT framework in Brazil in 1999 has generated significant interest in understanding the monetary transmission mechanism and, as a result, there has been a growing subsequent literature seeking to identify and measure the channels through which the central bank's policy interest rate (SELIC) affects output and inflation. A description of some channels and a discussion of the central bank of Brazil's model is provided in Bogdansky, Tombini, and Werlang (2000), and this basic framework has formed the basis for the empirical studies reviewed below.

Minella (2003) estimates a recursive four-variable VAR using the overnight interest rate, inflation, output and M1 over the period 1975--2000, breaking the estimation into three sub-samples: the "moderate" inflation period of 1975--85; the high inflation period (1985--1994); and the post-1994 low-inflation regime. He finds that inflation inertia declines in the post-1994 period and that there is only weak evidence that monetary policy affects inflation in this post-stabilization period, even though his estimates point to significant effects of monetary policy shocks on output. Minella notes that this may well be because of an identification problem arising from the fact that the 1994--2000 period was dominated by interest rate responses to financial crises and defense of the exchange rate peg, rather than by the overriding objective of anchoring inflation expectations. A possible reason for this "anomalous" result is that the exchange rate was not included in the VAR.

The role of the exchange rate as a determinant of Brazilian inflation has been acknowledged in other studies. Bevilaqua, Mesquita and Minella (2007) find that the large appreciation of the real since 2005 has contributed significantly to the fall in inflation. In a similar vein, Favero and Giavazzi (2004) find that exchange rate movements affect inflation expectations, and, through this channel, the central bank interest rate setting. This suggests

that the exchange rate may not only affect current inflation by changing the cost of imported goods, but, as well, there may be an important expectational channel at work.

Some attention has also been devoted to the interest rate reaction function. Using an HP-filtered measure of the output gap, Minella and others (2003) find that the parameter on the output gap in the monetary policy reaction function has the wrong sign and is not statistically significant from zero. They argue that this could arise because of simultaneity bias caused by supply shocks that depress the output gap and raise inflation. In theory supply shocks should be netted out of measures of the output gap, and any failure to do so will induce misspecification errors. The same study also finds that exchange rate volatility has been an important source of inflation variability in Brazil when using a smaller VAR estimated on monthly data, but with a sample that includes the pre-IT period (1994--2002).

More recent work (da Silveira, 2008; and Furlani et al., 2008) has re-examined some of these issues from the perspective of a New Keynesian open economy DGSE model based on Gali and Monacelli (2005) whose parameters have been estimated with Bayesian techniques. Given their open economy set-ups, both studies have the exchange rate playing a key role in the transmission of monetary shocks via UIP, though none of them contemplate a similar role for the country risk premium as we do below. Furthermore, because all goods in the Gali-Monacelli set-up are tradables, real exchange rate changes are proportional to terms of trade changes. Da Silveira (2008) finds in particular that monetary policy lowers inflation via a strong nominal exchange rate appreciation, but such effects are not particularly strong and they are reinforced through the effect of monetary policy on the output gap. Further, as in the case of other studies for Chile (see below), Furlani et al. (2008) examine whether the monetary policy reaction function should respond to exchange rates and output as well as to inflation. They find that the Brazilian central bank does not respond much to exchange rate movements in setting domestic interest rates, but mostly to inflation developments and also, to some extent, the output gap. Both studies also find, as we do, that shock accommodation is relatively swift, though none of their models have an allowance for a bank credit channel, which our estimates below show as playing a role.

## **2.2 Chile**

As with Brazil, existing work on monetary transmission in Chile has moved from an earlier literature using VARs and semi-structural VARs to more recent work using DGSE modeling and Bayesian estimation. Early work in the VAR tradition includes Morande and Schmidt-Hebbel (1997), Valdes (1998), Calvo and Mendoza (1999), and Cabrera and Lagos (2002). While structural restrictions have been imposed in some of these studies, strong reliance was placed



on *a-theoretic* identifying assumptions and the link between the estimated VAR and a theoretically based structural model was weak. Not surprisingly in this literature, a number of puzzles ("price", "exchange rate", and liquidity puzzles) emerged (see Chumacero, 2003 for further discussion).

Much of the recent work has as its basis the small open economy model with Keynesian features set out in Gali and Monacelli (2005). This features monopolistic competition with Calvo pricing, differentiated output varieties, and complete asset markets. Céspedes and Soto (2005) present a variant of this model in which there is uncertainty about the monetary policy rule implemented by the central bank, implying that agents simultaneously optimize and solve a signal extraction problem about the nature of the monetary policy shock. Computing impulse responses under standard calibrations of the model for the case of a disinflation shock (a shock to the inflation target), they find that the higher this uncertainty (i.e., the lower degree of the central bank credibility), the slower is the fall in inflation to a given monetary tightening along with a higher real exchange rate appreciation and sacrifice ratio. Complementing this calibration exercise with GMM estimates of the monetary policy rule over the pre-1999 and the post-1999 period, they find that, in the full-fledged inflation targeting regime, monetary policy has become more forward-looking (i.e., more responsive to expected future inflation rather than current inflation), and that the coefficient on the deviations of inflation from target in the monetary policy rule has gone up (rather than come down).

Caputo et al. (2007) extend the basic open economy New-Keynesian model to incorporate nominal wage rigidity, habit persistence and a risk premium on external borrowing (rather than complete international asset markets). They then estimate this model with Bayesian techniques. They find that wage rigidity is typically more important than price rigidity for the Chilean economy, which complicates the trade-off between stabilizing inflation and output. Specifically, wage indexation generates a more persistent response of inflation to shocks and makes inflation fluctuations (and monetary policy responses to it) more costly in terms of output and employment. Estimates of the monetary policy response embodied in their model indicates that the policy response to inflation during the full-fledged IT period is stronger than that to output and the exchange rate. Furthermore, they also find that this period has witnessed greater interest rate smoothness with the responses to inflation (relative to output) becoming less aggressive. In fact, their estimates of the central bank reaction function perhaps indicate too mild a response to inflation developments as the estimated parameters fail to meet the standard stability condition for a Taylor rule in a closed economy.

Del Negro and Schorfheide (2008) have a similar strategy to ours in that they estimate a DSGE model (a version of the Gali-Monacelli model) to derive predictions of what the VAR coefficients ( $\Pi$ ) would be,  $\Pi^*$ . A prior distribution for  $\Pi$  is then constructed by centering on

$\Pi^*$  and having a covariance matrix that is proportional (through the inverse of a hyperparameter  $\lambda$ ) to the form of the covariance matrix of  $\hat{\Pi}_{OLS}$ .  $\lambda$  determines the extent to which the VAR coefficients are preferred to those from the DSGE model. When  $\lambda = 0$  one would adopt the unrestricted VAR values for  $\Pi$  while, as  $\lambda$  becomes large, one would prefer the values implied by the estimated DSGE model (of course, one also needs to determine the covariance matrix of the VAR errors as well). The parameter  $\lambda$  basically enables one to explore how sensitive the conclusions will be to whether one uses the DSGE model or the VAR. One might choose  $\lambda$  by reference to predictive success and they use the highest posterior probability as a criterion. There are some difficulties in moving back to structural shocks, simply because these are defined by the DSGE model, and so it really needs to be correctly specified, but for monetary policy shocks the difficulties are much less as the structural equation defining this is *a-theoretic*.

A first question addressed is the extent to which the central bank responds to terms of trade and exchange rate fluctuations relative to inflation. Similarly to Caputo et al (2007) they find that Chile's central bank responds mostly to inflation rather than output. They also find evidence of very low pass-through from the terms of trade and nominal exchange rate shocks to CPI inflation - implying that the shocks which affect inflation are mostly domestic, rather than external. A second part of their investigation is to compare the impulse responses of the DSGE model and the combined DSGE-VAR model i.e. with an estimated  $\lambda$ . These are not very different, except for the exchange rate responses, and show little persistence. Just as we find in this paper, Del Negro and Schorfheide (2008) report strong effects of interest rate shocks on the output gap and inflation, though those on inflation are not as strong as our estimates suggest. As with the literature on Brazil, and, in contrast with the model we develop below, none of these studies contemplates a separate role for the credit channel in monetary transmission. Nor do they consider (with the partial exception of Caputo et al., 2007) an integration of external debt stock, current account flows and exchange rate dynamics, as we do below.

### 3 METHODOLOGY

#### 3.1 Types of Structural VARs (SVARs)

Let  $z_t$  be an  $n \times 1$  vector of variables in the macro-economic system. A typical structural equation in a macro model (normalized on the first variable  $z_{1t}$ ) has the form:

$$z_{1t} = z_t' \alpha_1 + z_{t-1}' \delta_1 + E_t(z_{t+1})' \gamma_1 + \varepsilon_{1t}$$

where  $\varepsilon_{1t}$  is a structural shock and  $z_t^-$  is  $z_t$  less  $z_{1t}$ . One approach to modeling these systems is structural VARs. The classic version of these are data-oriented in that their role is to fit the data as closely as possible but still provide a structural interpretation *in terms of the impulse responses*. The latter is usually the focus of the analysis to the point that it is extremely rare to see the fitted SVAR equations ever presented and their consistency with the underlying theoretical model discussed. This means that they might well fail to be consistent with theoretical ideas. One instance in which this has been the case is found in the literature employing SVAR models with long-run restrictions on the effects of money growth. Pagan and Robertson (1998) found that the structural equation meant to be identified as a supply curve was influenced positively by nominal money growth.

In standard SVARs  $\gamma_1 = 0$  and various restrictions are placed on  $\alpha_1$  (mainly exclusion restrictions) in order to identify the shocks  $\varepsilon_{1t}$ . In particular, the system is assumed recursive ("ordered") and is sometimes referred to as a "just-identified VAR". An alternative way of generating the structural equations of an SVAR comes from theory-oriented models such as DSGE models. These impose restrictions (mainly exclusion) upon  $\alpha_1$ ,  $\delta_1$  and  $\gamma_1$  in order to identify the shocks. Our approach is an intermediate one to these two polar cases in that we use a theory-oriented model to provide a skeletal structure which is then augmented (if necessary) to get a better match to the data. In some ways our approach resembles Del Negro and Schorfheide's, except that their focus is on the VAR that is the reduced form of the structural equations while ours focuses directly on modifying the structural equations.

Perhaps the simplest way to move from a theory-oriented SVAR to a data-oriented one is to substitute out  $E_t(z_{t+1})$  in the structural equations of the former. An approach that does not utilize any particular theory-oriented model, and is more robust to model specification error, is to make the expectation a function of some model variables  $\xi_t$ . Then if one regresses  $z_{t+1}$  against  $\xi_t$  to get coefficient estimates  $\hat{A}_1$ ,  $E_t(z_{t+1})$  could be measured as the combination  $w_t = \hat{A}_1 \xi_t$ .<sup>3</sup> This approach was used in the FRB-US model - see Brayton et al (1996)- although the variables  $\xi_t$  were only a few of those in the FRB-US system. Notice that, since  $\xi_t$  will generally involve both contemporaneous and lagged values of the model variables, the resulting SVAR will no longer be recursive. For example if one had a consumption Euler equation of the form

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<sup>3</sup> If direct measures of expectations were available these could be used.

$$n_t = E_t(n_{t+1}) + \delta(i_{t-1} - \pi_{t-1}) + \varepsilon_t^n$$

where  $n_t$  is consumption expenditure,  $i_t$  is a nominal interest rate,  $\pi_t$  is an inflation rate and  $\varepsilon_t^n$  is a preference shock, then substituting  $\xi_t' \phi_1 + \xi_{t-1}' \phi_2$  for  $E_t n_{t+1}$  will produce the following SVAR equation:

$$n_t = \xi_t' \phi_1 + \xi_{t-1}' \phi_2 + \delta(i_{t-1} - \pi_{t-1}).$$

Hence the original variables appearing in the structural equation have been augmented by  $\xi_t$  and  $\xi_{t-1}$ . What is critical though is that  $\phi_1$  and  $\phi_2$  can be estimated without reference to any structural model, so that the presence of these extra variables does not create any substantial estimation problems. Even if there was a coefficient attached to  $E(n_{t+1})$  the estimation issues are not major, since only a single variable  $\xi_t' \phi_1 + \xi_{t-1}' \phi_2$  needs to be instrumented.

Our strategy then will be to construct an SVAR by first setting out a small theory-consistent model and then replacing the expectations appearing in it by what would be implied by an unrestricted VAR. Thereafter, we ask whether the resulting structural equations need to be augmented with further information (largely lagged values of the system variables). Thus an important part of our strategy is the skeletal model that is to be the core of our SVAR. Whilst many choices of this might be made we will use a relatively standard New Keynesian model set out in the next sub-section.

### 3.2 The Skeletal Structure for our Structural VAR

Our starting point for structuring the VAR is a canonical small macro model that has been used quite extensively in the macroeconomics literature, and has often been deployed for analysis at both the IMF and various central banks (cf. Berg, Karam, and Laxton, 2006). It is implicitly derived from optimizing (Euler) equations for consumption and investment (which we aggregate to domestic absorption), a Phillips curve type of equation for inflation, an exchange rate equation driven by uncovered interest parity (UIP) and a Taylor type rule relating the policy-controlled interest rate to expected inflation and the output gap. Our variant of it distinguishes between absorption ( $n_t$ ) and output ( $y_t$ ). Later we will use a convention that a coefficient  $\alpha_{xy}$  shows a contemporaneous effect between  $x_t$  and  $y_t$ ,  $\beta_{xy}$  shows the effect between  $x_t$  and  $y_{t-1}$  and  $\gamma_{xy}$  is between  $x_t$  and  $y_{t-2}$ . Some license is taken

when expectations are involved. Thus  $\alpha_{ne}$  is the coefficient in the absorption equation that connects  $n_t$  and the expected value  $E_t(\tilde{n}_{t+1})$ . Thus, the model can be written as<sup>4</sup>:

$$\begin{aligned}
\tilde{n}_t &= \alpha_{m^e} E_t(\tilde{n}_{t+1}) + (1 - \alpha_{m^e}) \tilde{n}_{t-1} + \beta_{nr} \hat{r}_{t-1} + \varepsilon_t^n \\
\tilde{y}_t - \tilde{n}_t &= \beta_{yz} \tilde{z}_t + \beta_{yy^*} \tilde{y}_t^* + \delta_{yn} \tilde{n}_t + \varepsilon_t^y \\
\hat{\pi}_t &= \alpha_{\pi^e} E_t(\hat{\pi}_{t+1}) + (1 - \alpha_{\pi^e}) \hat{\pi}_{t-1} + \alpha_{\pi y} \tilde{y}_t - \alpha_{\pi z} \Delta \tilde{z}_t + \varepsilon_t^\pi \\
\hat{i}_t &= \beta_{ii} \hat{i}_{t-1} + \alpha_{i\pi^e} E_t \hat{\pi}_{t+1} + \alpha_{iy} \tilde{y}_t + \varepsilon_t^i \\
\tilde{z}_t - E_t \tilde{z}_{t+1} - (\hat{r}_t - \hat{r}_t^*) &= \zeta_t + \varepsilon_t^z \\
\Delta \hat{d}_t &= \bar{d} [(1 - \bar{\psi}) \hat{i}_t - (1 + \bar{i}) \hat{\pi}_t] + \omega_n \tilde{n}_t - \tilde{y}_t + (\omega_m - \omega_x) \tilde{z}_t - \omega_x \hat{t}_t \\
\bar{\psi} = \bar{\tau} &= \bar{\pi} + \Delta \ln \bar{Y}_t \\
\hat{r}_t &= \hat{i}_t - E_t(\pi_{t+1} - \bar{\pi}_{t+1}).
\end{aligned}$$

The first equation provides a specification for the log of domestic absorption ( $n_t$ ) - absorption being GDP minus net exports. It is measured as a log deviation from some "equilibrium" value and so should be regarded as a "gap" variable. Models that emphasize "gaps" are a convenient way of organizing policy and forecast discussions, allowing one to concentrate separately upon where one sees the system heading and the path of adjustment to that point. Most modern macroeconomic models can be written as "gap" models and so the approach is fairly flexible. The equilibrium value may be a constant or time varying.<sup>5</sup> In this case the absorption gap depends upon the real rate of interest  $r_t$ . The definition of the real rate will involve an expected inflation rate. In steady state this would be the target rate of inflation ( $\bar{\pi}_t$ ) and so we will work with the real interest rate adjusted for the inflation target,  $\hat{r}_t$ , given in (\ref{last}). In most empirical work with the NK model the target is taken as a constant, but this cannot be the case for Brazil or Chile over the whole period of inflation targeting. When the target is varying it may be reacting to the past inflation rate. Indeed a simple regression of the target on lagged observed inflation in Brazil does suggest such a relation, although it is rather weak so we have chosen to treat the target as exogenous. Notice that no other variable

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<sup>4</sup> A  $\bar{\cdot}$  indicates a log deviation from equilibrium values and a  $\hat{\cdot}$  indicates a levels deviation.

<sup>5</sup> The need for the latter often reflects the fact that there are permanent stochastic components that need to be removed to induce stationarity in the measured gap. Later we will provide a discussion of such transformations in detail but, even if there were no permanent components in the data, it is often the case for emerging-market economies that the equilibrium values to which the system will be adjusting are shifting over time in response to structural changes in the economy. Consequently, care should be taken when constructing these gaps and any assessments of the resulting measures should rely heavily on institutional knowledge of the economy being studied. This knowledge can be quite informative, not only of the presence of structural changes (such as those in the transition from high to low inflation regimes and across monetary policy frameworks), but also of how sensible one's estimation results appear to be. A striking example in the empirical macro literature of the problems arising from ignoring country-specific features in broad cross-country regressions pertains to the identification of the long-run effects of fiscal deficits on inflation; although solidly backed by theory, these effects are not easily discernable without properly taking into account country--group specific features in the estimation strategy. See Catão and Terrones (2005).

in this model determines the level of absorption, which is consistent with the standard Euler equation for consumption in DSGE models.<sup>6</sup> Implied in such a specification is that the other variables making up absorption, principally investment, are also functions of a real interest rate. While it might be worth considering augmenting this equation with some expressions for the rate of return on investment and other measures of the actual relative price or cost of capital (including tax wedges for instance), such measures are not readily available for emerging markets. Notice that the proposed specification also implicitly captures (through the lagged terms on absorption) accelerator effects, which are often found to have significant explanatory power in investment equations.

The second equation is meant to determine output and links the real GDP gap ( $\tilde{y}_t$ ), the domestic absorption gap ( $\tilde{n}$ ), import and export gaps. For Brazil the import and export shares are largely the same so that  $\tilde{y}_t - \tilde{n}_t$  can be regarded for that country as the log deviation of the current account from zero. Since imports are determined by total expenditure and the real exchange rate, while exports are related to the real exchange rate and foreign expenditure, we simply eliminate imports and exports to produce a relation linking the output and domestic absorption gaps, the log of the real exchange  $z_t$  (measured as a deviation from steady state) and the foreign output gap  $\tilde{y}_t^*$ .

The specification assumes that there is no lag between trade flows and their determinants but, since this is not something derived from any theoretical framework, it will need to be investigated. In fact, it might be thought that there will be lags because of the delays between orders and deliveries.

Although this equation is fundamentally an identity, because exports and imports have been replaced by some functional form, it will no longer be so. Hence we add a shock to it to allow for this. Such a shock is useful since in many emerging-market countries the opening up of the economy produces an import surge that is much larger than expected from the price and output elasticities for import demand. Although much of this movement can be accounted for through a time varying equilibrium value for the import share, one will probably want some of the changes observed to be captured by a shock that is persistent.

The third equation provides a specification for the "inflation gap",  $\hat{\pi}_t$ , where  $\hat{\pi}_t$  is the deviation of inflation from the target rate of inflation. It includes the output gap and the

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<sup>6</sup> As discussed in the next section, we will augment this canonical specification to include the role of domestic interest spreads - an interest wedge which arises in models with deposit- and credit-in-advance constraints (see, e.g., Edwards and Vegh, 1997). Since the domestic interest spread is itself a function of the policy interest rate as well as of a measure of the supply side of bank credit, this baseline specification for absorption will remain unchanged except for the addition of an extra term on the "excess credit" measure.

"exchange rate gap".<sup>7</sup> As suggested by previous studies the exchange rate has a very significant role to play in influencing the price of tradeables and CPI inflation.

The next equation is a monetary policy reaction function, where  $\hat{i}_t$  is defined as the nominal interest rate less the target inflation rate. The parameter  $\beta_{ii}$  seeks to capture the degree of interest rate smoothing in central bank policy, which is usually highly significant in policy reaction functions and Brazil and Chile have been no exceptions. In light of evidence from existing studies reviewed in section II, we do not include the exchange rate in the monetary policy rule. In our background empirical work, we did test if the exchange rate should be present and any dependence seemed indeed very weak - this is shown explicitly later for Brazil.

The exchange rate equation is risk-adjusted UIP. Note that the exchange rate is defined such that a rise in it represents an appreciation. There are two shocks in this equation. One,  $\zeta_t$ , is a "risk premium" that can be considered as relating to model variables, while the other,  $\varepsilon_t^z$ , is a function of non-model variables, and is treated as white noise and uncorrelated with  $\zeta_t$ . As standard in many real business cycles models of small open economies (see, e.g., Schmitt-Grohe and Uribe, 2003), this risk premium shock is generally made a function of the level of net foreign debt relative to GDP (again measured relative to a steady state value), but other factors may play a role, such as the level of domestic interest rates. This is because much of domestically issued debt in these countries is held by foreigners. These external debt servicing obligations tend to increase country risk. Inclusion of debt obligations as a gauge of country risk and a wedge in UIP equations is not only appealing from a theoretical perspective, but also consistent with recent external developments in many emerging markets, where a decline in net external debt has been accompanied by a decline in standard measures of country risk such as the EMBI spread.

The debt equation is a linearized version of the identity describing how the level of external debt to GDP changes over time, the derivation of which is fleshed out in Catão et al (2008).<sup>8</sup> In this equation, the terms of trade ( $\hat{t}$ ) are taken as exogenously given - a reasonable assumption for a small open economy.  $\omega_m$ ,  $\omega_x$  are  $\omega_n$  are the import, export shares and

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<sup>7</sup> We have constrained the coefficients in this equation so that they add to one. In many models they add to the discount factor. Since in a quarterly model this will often be around .99, we follow a common practice that enforces a restriction that the coefficients sum to unity. It is known that this aids identification.

<sup>8</sup> In its application to the present paper, we have chosen ignore the impact of current exchange rate variations on current debt revaluations. This arises from the fact that we did not have evidence on the frequency of revaluations of actual debt to exchange rate movements, but it seems unlikely that it would react strongly to contemporaneous movements in exchange rates.

absorption shares in GDP, respectively,  $\bar{d}$  is the steady-state ratio of net foreign debt to potential GDP, and  $\psi_t$  is the nominal potential GDP growth ( $\Delta \ln \bar{Y}_t$ ). Note that, since the evolution of the terms of trade and the trade balance shape the path of net external debt through this identity, and the level of debt would affect the exchange rate through the risk premium term, those variables will also be potentially important determinants of the exchange rate, output and inflation.

### 3.3 Augmenting the Skeletal Structure

As mentioned in the previous section our first step in extending the skeletal structure is to replace the expectations  $E_t n_{t+1}$ ,  $E_t \pi_{t+1}$  and  $E_t z_{t+1}$  by functions of the model variables. In general it will be the case that every variable affects the conditional expectations, although simulations of theory-oriented models such as that in Berg, Karam and Laxton (2006) ( using calibrated values for the parameters) suggest that many of the variables are of little importance. Now, the skeletal model will have a  $VAR(2)$  solution for the variables if the model shocks followed a  $VAR(1)$ .<sup>9</sup> But a  $VAR(2)$  would also be consistent with a range of other models. Thus a reasonable strategy is to begin by assuming that a  $VAR(2)$  should capture expectations quite well i.e. a  $VAR(2)$  in the skeletal model variables is fitted and used to produce  $E_t n_{t+1}$ ,  $E_t \pi_{t+1}$  and  $E_t z_{t+1}$ . Because there will be sixteen parameters in this regression, and our data sets are only 40 observations long, it seems sensible to delete some of the variables from the regressions if they did not seem important to the explanation of the variables expectations are being formed about. Generally we retained a variable in the expectations generating equation if its t ratio exceeded (or was close to) two.

The canonical model also needs to be expanded to capture previous research outlined in section 2. We have already augmented the standard NK model to reflect open-economy considerations more precisely. This involved separating out absorption and output effects as well as introducing an external liability equation. A second extension was to incorporate an equation for private bank credit growth. In the Brazilian and Chilean financial systems, many firms (particularly smaller and medium-sized) are still largely dependent on domestic banks for

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<sup>9</sup> In fact the solution is a VARX model with two lags in the endogenous model variables and no lags in the exogenous external variables ( $y_t^*$ ,  $r_t^*$  and  $tt$ ). If the exogenous variables are represented as a VAR(1) then the VARX model can alternatively be written as a VAR(2) in all endogenous and exogenous variables, but with the special structure that the exogenous variables depend only upon their own past history.



funding, and have limited access to international capital markets. Consequently, this channel might be expected to play a role. At least it needs to be examined.

Ideally we wish to capture a credit channel effect involving the amount of credit being granted by banks that is in excess of some “normal” level. Microfounded models of the credit channel featuring deposit- and credit-in-advance constraints as well as costly banking (Edwards and Végh, 1997; Goodfriend and McCallum, 2007) imply that a wedge appears in the Euler equation governing absorption. Similar emphasis on the amplification mechanism associated with bank interest rate spreads is found in the earlier literature on the “credit view” (Bernanke and Blinder, 1988; Kashyap and Stein, 1994), where it is suggested that lending-deposit spreads ought to feature in an absorption equation like (\ref{one}). As shown in Edwards and Végh (1997) for a typical emerging market context, such spreads are a direct positive function of the policy interest rate itself plus a term related to the credit-to-expenditure ratio.<sup>10</sup> Therefore it would seem logical to use such a ratio as credit to either GNE or GDP. But this is a difficult series to work with for both countries due to a sharp rise in it over the sample period e.g. in the case of Brazil, after being fairly stable around .24 from 1999 to 2004, it rose sharply from 2004 to be .44 at the end of the period. Hence it behaves more like an integrated series over our data period.<sup>11</sup>

Because of the problems with this data set we use credit growth relative to a constant level as a proxy. As we observe in the next section, many gap measures are in fact constructed from the growth rate of a series. So the gap between the log of the credit to GDP ratio and its normal level would be a function of the growth in credit less the potential growth rate in GDP. Since the latter is reasonably constant this suggests that our measure captures the ideas about excess credit in a reasonable way, but one that is more tractable statistically, since it is a stationary process.<sup>12</sup>

Since at any point in time domestic bank credit is endogenously determined within a system, it is necessary to decide how to account for this within an SVAR. In order to allow for contemporary effects of credit expansion on expenditure, we place the excess credit variable before any other variable of the system i.e. we assume that it is determined only by the past values of any model variables. To be consistent with some theoretical ideas, we restrict the

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<sup>10</sup> The specific way in which Edwards and Vegh (1997) model bank technology yields a relationship between the lending and deposit spreads (measured relative to the base interest rate) and the credit to deposit ratio. But since deposit-in-advance constraints imply that deposits are proportional to expenditure, this directly translates into a functional relationship between bank spreads and the credit to expenditure ratio.

<sup>11</sup> The estimated AR(1) coefficient is 1.02.

<sup>12</sup> Allowing for the presence of an autonomous component in "excess credit" variable which is not directly related to the interest rate seems particularly appropriate in the case of Brazil, where a large development bank (BNDES) accounts for up to one-fourth of domestic credit and whose lending policies and rates arguably respond to other incentives and whose lending rates are typically below market rates.

explanatory variables entering the excess credit equation to be absorption, the real interest rate (separated into its nominal and inflation components), and the real exchange rate. The latter can enter the equation for two reasons. First, there may be sizeable balance-sheet effects of the type documented in Calvo, Izquierdo and Mejia (2004).<sup>13</sup> Secondly, it is possible that the non-tradable sector may be more “bank-dependent” than the tradable sector (cf. Catão and Chang, 2010). Both structural features imply that a real exchange rate (REER) appreciation will increase real credit demand in the non-tradeable sector, leading to higher aggregate credit. Although neither Brazil nor Chile are as dollarized as many other emerging markets, significant balance-sheet effects may also be present, and they could further strengthen the positive impact of a REER on domestic credit. Section 5 provides supportive econometric evidence that real exchange rate appreciations tend to foster domestic credit growth for Brazil but not for Chile.

We now need to make some specific comments about how the canonical model is to be augmented following the elimination of expectations. An obvious extension was to add higher order lags in the structural equations. The need for such higher order lags was often suggested by the fact that there was serial correlation left in the individual equations of the skeletal SVAR. A second lag in the dependent variable was sometimes found to be needed. Because it is often the case that the coefficients of the two lags appear with opposite signs it is worth noting that a term such as  $az_{t-1} + bz_{t-2}$  (with  $a > 0, b < 0, |b| < a$ ) can be written as  $(a + b)z_{t-1} - b\Delta z_{t-1}$  i.e. there is both a level and a growth rate effect.

Estimation of each of the equations was done with instrumental variables. This was done partly to avoid the fact that systems estimation requires that the specification of all the equations of the system be correct to yield the expected efficiency gains. Otherwise partial systems methods involving the use of instrumental variables should be preferable, and this is the route we take here. In doing so, some rules governing the selection of instruments need to be given. First, any exogenous or lagged variable appearing in an equation will be taken as an instrument for itself. Second,  $E_{t-1}\tilde{n}_t, E_{t-1}\tilde{\pi}_t$  and  $E_{t-1}\tilde{z}_t$  from the VAR(2) were used as instruments for  $\tilde{n}_t, \tilde{\pi}_t$  and  $\tilde{z}_t$ . Finally, residuals from structural equations further up the system were taken to be suitable instruments. Thus the residual from the credit equation is used as an instrument in the absorption equation. This can be justified if the assumption used in many DSGE models that the shocks in the structural equations are mutually uncorrelated

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<sup>13</sup> Even though the “dollarization” of private sector liabilities in Brazil is not nearly as extensive as in many other EMs, it is far from negligible. Starting from negligible amounts in the early 1990s, foreign currency denominated debt rose to 36% of total corporate debt in 1999, reaching 40% in 2002 (Bonomo, Martins, and Pinto, 2004, Table A.2). Using a large panel of firm-level data, Bonomo and others (2004) also find that balance sheet effects of currency movements have significant effects on credit demand and investment.

with one another is valid. If the number of instruments equaled the number of variables in each equation, and residuals were among the former, then we would be enforcing this restriction. This is not strictly true however if we have an excess of instruments, but using the residuals as instruments does tend to enforce it. In some cases the residuals can be good instruments e.g. the correlation of  $\varepsilon_t^n$  with  $\tilde{n}_t$  is .41 (Brazil) and .88 (Chile), but in other instances we might expect that the conditional expectation would be a more powerful instrument. After estimation we checked if the shocks were mutually uncorrelated and the correlations were not significantly different from zero. It is desirable to have uncorrelated shocks for well defined policy experiments.

As noted previously, our convention is that a coefficient  $\alpha_{xy}$  shows a contemporaneous effect between  $x_t$  and  $y_t$ ,  $\beta_{xy}$  shows the effect between  $x_t$  and  $y_{t-1}$  and  $\gamma_{xy}$  is between  $x_t$  and  $y_{t-2}$ . Thus the credit growth equation for Brazil could be written in the form

$$pc_t = \beta_{cc} pc_{t-1} + \beta_{cy} \tilde{y}_{t-1} + \beta_{ci} \hat{i}_{t-1} + \beta_{cz} \tilde{z}_{t-1} + \gamma_{cz} \tilde{z}_{t-2} + \varepsilon_t^c$$

while the absorption equation might be written as

$$\tilde{n}_t = \alpha_{ne} (E_t \tilde{n}_{t+1}) + (1 - \alpha_{ne}) \tilde{n}_{t-1} + \beta_{nr} (\hat{i}_{t-1} - E_{t-1} \hat{\pi}_t) + \alpha_{nc} pc_t + \beta_{nc} pc_{t-1} + \varepsilon_t^n$$

For the equations generating expectations we add a superscript “e” to the coefficients. Hence we might have:

$$E_t \hat{\pi}_{t+1} = \alpha_{\pi\pi}^e \hat{\pi}_t + \alpha_{\pi z}^e \tilde{z}_t + \alpha_{\pi y}^e \tilde{y}_t + \alpha_{\pi n}^e \tilde{n}_t + \beta_{\pi z}^e \tilde{z}_{t-1}.$$

## 4 DATA

### 4.1 Brazil

We restricted our sample to the IT period as a response to evidence that there have been large structural changes around the point of its introduction. Limiting the estimation to the post-1998 period is advisable in light of evidence from Tobini and Lago-Alves (2006) which suggests that there were significant structural changes in inflation dynamics and exchange rate pass-through in Brazil before and after 1999, as well as far-reaching changes in the price indexation system and inflation dynamics since 1995 (see Minella, 2002).

Seasonally adjusted national income account data was taken from IMF's International Financial Statistics (IFS) and the Brazilian Planning Ministry Research Institute (IPEA).

Domestic bank credit to the private sector was taken from the same sources and seasonally adjusted using the X11 routine in AREMOS. The real exchange rate series is from the IMF and is computed as a weighted average among nearly all trading partner using CPI deflators and 2000 weights. The indicator of world real income has been computed as the trade weighted average of real GDP of its main trading partners, who account for over 80 percent of Brazil's trade.

The inflation variable is seasonally adjusted CPI including all items -- administered and “free market” prices. While it has been customary to separate the two on account of the belief that “administered” prices have a stronger backward-looking adjustment component (largely due to the nature of the multi-year contracts between the government and the new incumbents in the utility industries privatized during the 1990s), we see this distinction as somewhat artificial. For a number of reasons it can be potentially misleading for the purpose of setting the monetary policy stance and perhaps irrelevant if the task at hand is indeed to model aggregate inflation. A first reason is that administered prices still respond to demand pressures, even if with a longer lag, because of backward-looking indexation clauses in the underlying concessional contracts. Second, although utility prices are typically key inputs to “free market” prices, the interaction between the two is certainly complex and, even though one might add both series into a VAR, this would be unlikely to address such complexity. Thirdly, the extent to which wage earners make such a distinction between the types of inflation is unclear. Indeed, if they only care about overall inflation, second-round effects will stem from this, and that reduces the advantage of decoupling the two inflation rates. For these reasons, the estimation results reported below refer to the “all item” CPI.

To parameterize (\ref{debt}) we need  $\omega_m$ ,  $\omega_x$  and  $\bar{d}$ . These were replaced by the average import, export and net debt ratios to GDP over 1999-2009. Because  $\omega_m$  and  $\omega_x$  were virtually the same over this period we fix them both at .11. The average growth rate in real GDP over the period and target inflation (4.5% in recent years) have been used to compute  $\bar{\psi}$ . This makes it 1.9% per quarter.

## 4.2 Chile

As with Brazil, we restrict our sample to the post-1999 period and use quarterly data throughout. Seasonally adjusted national income data was taken from the IMF's International Financial Statistics and the Central Bank of Chile. The real exchange rate series is from the IMF and is computed as a weighted average among nearly all trading partners using CPI deflators and 2000 weights. As for Brazil, the indicator of world real income has been computed

as the trade weighted average of real GDP of its main trading partners, although in this case accounting for over 90 percent of Chile's foreign trade.

To parameterize (\ref{debt}) we need  $\omega_m$ ,  $\omega_x$  and  $\bar{d}$ . The first two were replaced by taking the simple average of  $\omega_m$  and  $\omega_x$  over 1999-2009. The debt ratio was the historical average over this period. Likewise, the average growth rate in real GDP (around 4%) and target inflation (3%) have been used to compute  $\bar{\psi}$ . This makes it 1.95% per quarter.

### 4.3 Producing Gap Measures for both Countries

Gap measures are present in the skeletal model we use and so estimates of them must be made. In many cases the permanent component is extracted with the Hodrick-Prescott (HP) filter. To explain why we do not do this we note that Harvey and Jaeger (1993) pointed out that the HP filter can be regarded as extracting a permanent component  $P_t$  from a series  $z_t$  by applying the Kalman Smoother to the state space model

$$\begin{aligned} z_t &= P_t + T_t \\ \Delta^2 P_t &= v_t \\ T_t &= u_t \\ \lambda &= \frac{\text{var}(u_t)}{\text{var}(v_t)}. \end{aligned}$$

The model clearly implies that:

$$\begin{aligned} \Delta^2 z_t &= \Delta^2 P_t + \Delta^2 T_t \\ &= v_t + \Delta^2 u_t \\ &= e_t n \\ &= e_t + \alpha_1 e_{t-1} + \alpha_2 e_{t-2}, \end{aligned}$$

where  $e_t$  is an uncorrelated process. Setting  $\lambda = 1600$  we find that  $\alpha_1 = -1.77$ ,  $\alpha_2 = .8$ .

Fitting this model to Brazilian GDP data over 1999/1-2009/1 we get  $\alpha_1 = -.95$ ,  $\alpha_2 = -.05$ . Of course this process has a common unit root to the MA and AR parts which cancels, implying that the log of Brazilian GDP is an  $I(1)$  process, which contrasts with the implied  $I(2)$  model when using a HP filter.

This suggests that we want to utilize a measure of the permanent component of a series that is extracted under the assumption that data is  $I(1)$ . One such filter that does this is the Beveridge-Nelson (BN) filter. The logic of this is that the permanent value of  $z_t$  is

$$\begin{aligned} P_t &= E_t(z_\infty) \\ &= E_t\left\{z_t + \sum_{j=1}^{\infty} \Delta z_{t+j}\right\} \\ &= z_t + E_t \sum_{j=1}^{\infty} \Delta z_{t+j}, \end{aligned}$$

so that the transitory component is  $z_t - P_t = -E_t \sum_{j=1}^{\infty} \Delta z_{t+j}$ , and this is the “gap”. Thus it is necessary to prescribe a model for  $\Delta z_t$  in order to be able to compute the transitory component.

In the situation when  $\Delta z_t$  is an  $AR(p)$ ,  $E_t \sum_{j=1}^{\infty} \Delta z_{t+j}$  will be a linear function of  $\Delta z_t, \Delta z_{t-1}, \dots, \Delta z_{t-p+1}$ . Hence the BN measure of the output gap in that case is constructed as the negative of an average of growth rates. Notice that this means that one will see a *negative* relation between the output gap and growth, so that regressing inflation against the growth in output should produce a negative coefficient on the latter.

It is sometimes said that the problem with the BN estimate of an output gap is that it is not smooth enough in comparison with the HP filtered estimate. If one uses a low order AR process to approximate  $\Delta y_t$  this may well be true but, when a higher order AR is adopted, it is often much smoother (see, e.g. Morley, 2007). The intuition is that the gap is constructed by averaging growth rates, and that will generally result in some persistence in the output gap measure. However, the greater smoothness seen with the HP filter comes from two sources. One is the assumption that the permanent component evolves very smoothly i.e. it is  $I(2)$ , and the other is that it is a *two-sided filter*, and so is using weighted averages of growth rates into both the past and the future. To see this we HP filtered the Chilean log of GDP series and then regressed this against three lagged and forward values of GDP growth. That produced a regression of the form

$$\tilde{y}_t^{HP} = .59\Delta y_t + .19\Delta y_{t-1} + .19\Delta y_{t-1} + .11\Delta y_{t-1} - .54\Delta y_{t+1} - .09\Delta y_{t+2} - .26\Delta y_{t+3}.$$

While 75 percent of the variation in  $\tilde{y}_t^{HP}$  is explained by these variables, only 33% is due to that of the lagged and current values of  $\Delta y_t$ . Hence the BN filter is unlikely to approximate the HP one too closely while it remains a one-sided filter. The relation between

the HP filtered gap and growth rates seen above shows that there are clear econometric issues with using the former as a regressor, since future values of the growth rates are involved. Indeed in Laxton et al. (1992) it was found in a simulation experiment in which the potential level of output actually followed an  $I(1)$  process, that using the output gap coming from a HP filter produced an estimate for the parameter on the output gap in a Phillips curve that was well below the true value used in producing the simulated data.

In each case the BN filtered output and absorption gaps were found by fitting an  $AR(4)$  to growth rates over the sample period. When it comes to measuring the foreign output gap a difficulty arose in that the appropriate series (the trade-weighted GDP of Brazil's and Chile's trading partners) have growth rates that are extremely persistent - over 1999/1-2009/1 fitting an  $AR(1)$  to Brazil data gave a point estimate of the  $AR(1)$  coefficient of .99. Hence, in this case, the BN filter is not the appropriate one, and the HP filter is much closer to what is needed. Since we have no model of trading partners GDP (this is treated as exogenous), there seems to be no reason not to use the HP filter on that data to produce a foreign output gap.

## 5 MODEL ESTIMATES

We will present the estimated structured VAR models for Brazil and Chile along with the impulse responses to a 100% basis point rise in the annualized interest rate and a 1% point rise in the growth rate in credit. It should be borne in mind that these rises are relative to the steady state levels and that all variables in the equations of the model are intended to be measured that way. Thus what is being explained is the nature of the adjustment process back to equilibrium. There may be forces here that are not present in equilibrium e.g. as noted earlier nominal interest rates may affect disequilibrium expenditure but in equilibrium we expect that expenditure will be governed by the real rate of interest.

As we observed in the introduction it is rare for those using SVAR models to present the structure. Rather, only impulse responses are given. One reason for this is that it is quite possible to have “sensible” impulse responses resulting from what might appear to be “odd-looking” structural equations. Indeed it has been our experience that this is very common. But there are three compelling reasons to present the structural equations of our model.

1. It enables one to see how the skeletal model needs to be modified to fit the data. This has some information for those who wish to just work with a theoretical model that is

close to our skeletal one. There are in fact quite a few papers that utilize variants of our skeletal model for policy analysis in emerging economies.

2. It is sometimes useful to be able to refer back to structural equations when seeking to determine an explanation for either the pattern or the magnitude of impulse responses. Indeed, one can conduct sensitivity analysis by varying the estimated structural parameters to see what the effect would be of adopting alternative parameter values. Given the small sample sizes it seems unlikely that we can precisely determine the values of these parameters, and so it is useful to be able to assess how sensitive the conclusions are to the point estimates of the structural equations used in constructing the impulse responses - a central theme in Del Negro and Schorfheide (1998).
3. A "full disclosure" principle seems a desirable attitude to have in empirical work. This would seem to demand the provision of information on the structural equations, even though this rarely done.

In the model we have annualized inflation and interest rates. This means that the UIP equation has to be changed accordingly. Gaps are converted to annual percentage values. The debt ratio is measured as net debt to annualized GDP.

## **5.1 Brazil**

### **5.1.1 The Model**

The structured VAR that was fitted is given below with  $t$  ratios in brackets:



$$\begin{aligned}
pc_t &= -0.35 pc_{t-1} + 2.67 \tilde{y}_{t-1} - 1.26 \hat{i}_{t-1} + 0.603 \tilde{z}_{t-1} - 0.326 \tilde{z}_{t-2} + \varepsilon_t^c \\
&\quad (-2.37) \quad (2.18) \quad (-2.19) \quad (3.44) \quad (-1.71) \\
\tilde{n}_t - \tilde{n}_{t-1} &= 0.61 (E_t(\tilde{n}_{t+1}) - \tilde{n}_{t-1}) - 0.025 \hat{r}_{t-1} + 0.004 pc_t + 0.006 pc_{t-1} + \varepsilon_t^n \\
&\quad (12.86) \quad (-2.66) \quad (1.22) \quad (2.29) \\
\tilde{y}_t &= 0.61 \tilde{n}_t - 0.02 \tilde{z}_{t-1} + 0.42 \tilde{y}_{t-1} + \varepsilon_t^y \\
&\quad (1.92) \quad (-1.92) \quad (2.92) \\
\hat{\pi}_t - \hat{\pi}_{t-1} &= 0.58 (E_t(\hat{\pi}_{t+1}) - \hat{\pi}_{t-1}) + 0.79 \tilde{n}_t - 0.17 \tilde{z}_{t-1} + 0.15 \tilde{z}_{t-2} + \varepsilon_t^\pi \\
&\quad (4.86) \quad (0.84) \quad (-3.01) \quad (2.49) \\
\hat{i}_t &= 1.08 \hat{i}_{t-1} + 0.26 E_t \hat{\pi}_{t+1} - 0.32 \hat{i}_{t-2} - 0.02 \tilde{z}_{t-1} + \varepsilon_t^i \\
&\quad (10.48) \quad (3.20) \quad (-3.93) \quad (-1.54) \\
\zeta_t &= 0.27 (\tilde{z}_{t-1} - E_{t-1} \tilde{z}_t) - 0.43 \hat{i}_{t-1} + 0.56 \hat{r}_{t-1}^* + 0.28 \hat{d}_{t-1} + 0.17 \hat{z}_{t-1} + 0.08 E_{t-1} \hat{\pi}_t + \varepsilon_t^\zeta \\
&\quad (1.1) \quad (2.8) \quad (2.9) \quad (2.1) \quad (1.93) \quad (0.47) \\
\zeta_t &= \tilde{z}_t - E_t \tilde{z}_{t+1} - \left[ \frac{(\hat{r}_t - \hat{r}_t^*)}{4} \right].
\end{aligned}$$

The equations generating expectations were:

$$\begin{aligned}
E_t \hat{\pi}_{t+1} &= 0.38 \hat{\pi}_t - 0.28 \tilde{z}_t + 0.22 \tilde{z}_{t-1} + 1.29 \tilde{n}_t \\
&\quad (2.89) \quad (-4.58) \quad (3.72) \quad (1.74) \\
E_t \tilde{n}_{t+1} &= -0.07 \hat{i}_t + 0.016 \tilde{z}_t - 0.02 \tilde{z}_{t-1} + 1.02 \tilde{n}_t - 0.70 \tilde{n}_{t-1} + 0.01 pc_t \\
&\quad (-2.46) \quad (2.25) \quad (-2.67) \quad (6.90) \quad (-5.83) \quad (1.65) \\
E_t \tilde{z}_{t+1} &= 0.26 \tilde{z}_t - 0.91 \tilde{y}_t - 4.12 \tilde{n}_t + 6.35 \tilde{y}_t^* - 0.8 \hat{d}_t . \\
&\quad (1.79) \quad (-0.95) \quad (-2.07) \quad (3.76) \quad (-4.21)
\end{aligned}$$

In constructing these equations we were aware that a small number of observations were being used and this probably meant we should err on the side of caution. Consequently, we tended to use the basic rule that a variable was left in the structural equations if it had a t-ratio greater than (or close enough to) unity. This is the equivalent of applying AIC to decide on whether a regressor should be retained. In some instances we introduced a variable into the model although it was not significant. One reason was when it was supposed to be present in the skeletal model and so this shows that there is evidence in the data against that part of the model. But in other cases it was our attempt to respond to what were likely to be queries from readers.

Looking at the equations we start with the credit growth equation. It is negatively affected by interest rates (and it seemed to be nominal rather than real, although it should be remembered that these rates are measured relative to their equilibrium rate which includes target inflation) and positively by the output gap. The high magnitude of the coefficient on the output gap is a response to the very high growth in credit relative to GDP. This suggests that more attention has to be paid to getting a better “excess credit” variable in the face of major

changes in credit availability over the sample period. The positive exchange rate effect that might have been expected from our earlier discussion is also observed.

Credit growth then augments the absorption equation provided by the skeletal model. It seems to play a significant role in affecting absorption.<sup>14</sup> Apart from that introduced variable, another item of interest is that forward-looking expectations seem to dominate the backward component (0.61 versus 0.39). For many advanced economies this might be regarded as unusual, but it is consistent with the conclusion of Caputo et al (2007) for Chile. However, setting the coefficient to .5 would not be inconsistent with the data at standard levels of statistical significance.

Since the skeletal model would have an identity for GDP connecting absorption, imports and exports, one would need import and export price and income elasticities to complete it. Instead we have just fitted a regression that aims to capture these missing functions. Because the import and export shares are much the same in Brazil the dependent variable is basically the deviation of the current account from its steady state value. The terms are much like what we would expect, although the exchange rate effects are not particularly strong. No lags in absorption were significant in this equation.

The inflation equation is close to the skeletal one although the absorption gap (a better performing variable than the output gap in this case) is not significant. The prediction by the skeletal model that it is the change in the exchange rate which affects inflation is easily accepted, but we have chosen to leave the exchange rate in level rather than difference form.

The interest rate rule has neither an absorption nor an output gap present, which is in line with Furlani et al. (2008). Expected inflation produced a better fit than actual inflation. It is also the case that the exchange rate has a small impact on interest rate decisions and it was therefore left in the relation. The most striking difference to the skeletal model is the presence of a second lag in interest rates. As noted earlier this equation can be re-written in terms of a first order lag of the policy interest rate with a coefficient of 0.76 and a lagged change in the policy rate with a coefficient of 0.32 i.e.  $.76\hat{i}_{t-1} + .32\Delta\hat{i}_{t-1}$ .

The exchange rate equation is more complex than habitually found in the literature, particularly in the stylized DSGE models used to fit these countries, as reviewed in section 2. It is worth thinking about its form by generalizing the skeletal model through a replacement of  $E_t\tilde{z}_{t+1}$  by  $\{\phi\tilde{z}_{t-1} + (1-\phi)E_t\tilde{z}_{t+1}\}$ . This is done in a number of empirical implementations of theoretical models (see e.g., Berg et al., 2006) to reflect the well known failure of UIP in

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<sup>14</sup> It should be noted that the equation actually estimated involved  $\Delta n_t$  as dependent and  $E_t(n_{t+1} - n_{t-1})$  as a regressor. We used  $E_{t-1}n_t$  and  $n_{t-1}$  as separate instruments for the latter variable. This was also true of the inflation equation.

exchange rates in the short-run. Doing so would produce the first term on the right hand side of the equation which we estimated, except for a component  $\phi(E_t z_{t+1} - E_{t-1} z_t)$ . But when this latter term is added to the regression it had a zero coefficient. It might be thought that the lagged nominal interest rate is an odd regressor and, to show that it is not proxying for a real interest rate, we added in the expected inflation rate, whereupon it is apparent that the latter is not accepted by the data. Once again however we note that  $\hat{i}_t$  is measured as a deviation from an equilibrium rate and, since this includes the expected inflation rate,  $\hat{i}_t$  cannot be thought of as a purely nominal rate.

### 5.1.2 The Impulse Responses

Figure 1 gives the impulse responses for the 100 basis point rise in interest rates and Figure 2 contains those to an increase of 1% in credit growth. Also present in them are “confidence intervals”. These were found by simulating the model with the point estimates of the parameters and then choosing the .025%-.975% range for the simulated parameter estimates. In doing these simulations it should be noted that the estimated parameters might imply an unstable VAR, as the debt equation is always close to a unit root. When it was unstable the simulated values were discarded.<sup>15</sup>

First the interest rate rise. The strong exchange rate appreciation is probably what produces the inflation response rather than the output gap effect. There is no "price puzzle" or "exchange rate puzzle" in the results. The bulk of the effects take place within five quarters. This entails a much shorter lag than in traditional closed economy models, based on what existing estimates show for the US and the Euro area (see, e.g., Angeloni et al., 2003). The immediate contractionary impact of the rate rise on the absorption gap is stronger than on the real GDP gap, so the trade balance improves. At the same time, the higher onshore-offshore interest differential appreciates the real exchange rate (UIP-type effects) and boosts external debt (e.g. through the carry-trade). Consistent with the theories discussed above on the credit channel in emerging markets, the initial real exchange rate appreciation tends to boost bank credit growth (through both higher relative price of non-tradeables and a positive balance sheet effect), somewhat offsetting the negative effect of monetary tightening on absorption through the inter-temporal channel. Thus, the positive effects of the appreciation on credit growth and

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<sup>15</sup> It should be noted that the oscillations seen in the credit shock confidence intervals come from the fact that the autoregressive parameter in the credit growth equation is negative. Because the point estimate is small the effect dies out quickly in the estimated impulse responses. But in some simulations a large negative value can be found and then the oscillations persist for quite some time.

hence on absorption kick in and the trade balance deteriorates between the second and fifth quarter after the shock, while the pace of disinflation and credit contraction both slow down. Overall, though, the negative inter-temporal channel still dominates, leading ultimately to fall in inflation, real credit growth, and absorption on average over the entire post-shock period.

Turning to the experiment involving credit growth, there is a rise in absorption and a rise in inflation. As expected, the absorption gap (which is equivalent to absorption growth in the immediate aftermath of the shock) increases (by around .02%). Yet, it rapidly disappears as interest rates rise and there is an exchange rate depreciation which chokes off credit growth and hence absorption. The effects are even more short lived than those associated with interest rate shocks, particularly regarding credit growth, virtually vanishing after four quarters. The same rise and fall is true for inflation, while lasting a quarter or two longer. Given there have been very large movements in real credit growth - with one standard deviation being the equivalent of 9% (annualized) growth - the impulses above understate the impact of credit over the period, since the "norm" is not 1% growth but variations which are some 9 times as high. This indicates that the macro effects of a standard deviation in credit growth appear to be of a higher magnitude, even if more short-lived, than in developed countries.<sup>16</sup> It is also worth noticing that such a strength of the credit channel is robust to dropping the other feature of our model which distinguishes it from more standard new-keynesian set-ups estimated in previous work - namely, the debt accumulation equation.<sup>17</sup> This indicates that the importance of the credit channel in monetary transmission stands on its own, quite separately from the open economy features of the skeleton model.

## 5.2 Chile

The structured VAR is fitted over 1999/1-2008/4 and its equations are given below. As for Brazil absolute t ratios are in brackets. The import, export and debt ratios were set to their averages over the period and the nominal growth in potential GDP was set to 1.95% per quarter, following from a potential growth rate of 4.0% p.a and target inflation of 3% p.a.. The answers are not sensitive to this choice.

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<sup>16</sup> See the discussion by Eichenbaum (1994) on the difficulties faced by empirical work in the identification of credit channel effects in the US, for which longer and better data and more disaggregated empirical evidence are available.

<sup>17</sup> As one might expect, the main effect of shutting off the debt accumulation equation is on the real exchange rate response. Estimated impulse-responses with the debt-accumulation equation shut-off are not reported to conserve on space but are readily available from the authors upon request.

$$\begin{aligned}
pc_t &= 0.5 pc_{t-1} - 0.57 \hat{r}_{t-1} + 0.15 \tilde{z}_{t-1} - 0.17 \tilde{z}_{t-2} + \varepsilon_t^c \\
&\quad (2.93) \quad (-.83) \quad (0.66) \quad (-0.77) \\
\tilde{n}_t - \tilde{n}_{t-1} &= 0.65 (E_t(\tilde{n}_{t+1}) - \tilde{n}_{t-1}) - 0.15 \hat{r}_{t-2} + 0.04 pc_{t-1} - 0.06 pc_{t-2} + \varepsilon_t^n \\
&\quad (5.53) \quad (-1.4) \quad (1.33) \quad (-2.25) \\
\tilde{y}_t &= 0.30 \tilde{n}_{t-1} - 0.05 \tilde{z}_{t-1} + 0.05 \tilde{z}_{t-2} + 0.43 \tilde{y}_{t-1} + \varepsilon_t^y \\
&\quad (2.43) \quad (-1.58) \quad (1.65) \quad (3.02) \\
\hat{\pi}_t - \hat{\pi}_{t-1} &= 0.49 (E_t(\hat{\pi}_{t+1}) - \hat{\pi}_{t-1}) + 0.21 \tilde{y}_t - 0.05 \tilde{z}_t + 0.04 \tilde{z}_{t-1} + 0.03 pc_t + \varepsilon_t^\pi \\
&\quad (10.35) \quad (2.86) \quad (-3.17) \quad (3.06) \quad (3.57) \\
\hat{i}_t &= 1.23 \hat{i}_{t-1} - 0.45 \hat{i}_{t-2} + 0.16 \tilde{y}_t + 0.14 E_t(\hat{\pi}_{t+1}) + \varepsilon_t^i \\
&\quad (8.73) \quad (-3.46) \quad (0.98) \quad 2.87 \\
\zeta_t &= 0.60 \tilde{z}_{t-1} - 0.22 \tilde{z}_{t-2} + 0.29 \hat{d}_t + \varepsilon_t^z \\
&\quad (7.38) \quad (-3.17) \quad (4.38) \\
\zeta_t &= \tilde{z}_t - E_t \tilde{z}_{t+1} - \left[ \frac{(\hat{r}_t - \hat{r}_t^*)}{4} \right].
\end{aligned}$$

The equations generating expectations were as follows.

$$\begin{aligned}
E_t \hat{\pi}_{t+1} &= 1.41 \hat{\pi}_t - 0.63 \hat{\pi}_{t-1} + 0.117 \tilde{n}_t + .025 pc_{t-1} + 0.325 \tilde{y}_t^* \\
&\quad (10.19) \quad (-4.34) \quad (1.16) \quad (1.59) \quad (2.65) \\
E_t \tilde{n}_{t+1} &= -0.13 \hat{i}_{t-1} + 0.4 \tilde{y}_t - 0.26 \tilde{y}_{t-1} - 0.513 \tilde{n}_{t-1} + 0.314 \tilde{y}_t^* \\
&\quad (-1.16) \quad (1.72) \quad (-1.39) \quad (-3.81) \quad (1.89) \\
E_t \tilde{z}_{t+1} &= 1.15 \hat{i}_t + 0.64 \tilde{z}_t - 0.32 \tilde{z}_{t-1} - 0.28 pc_{t-1} - 0.63 \hat{d}_t + 0.96 \tilde{y}_t^* \\
&\quad (2.38) \quad (3.49) \quad (-2.21) \quad (-2.55) \quad (-3.43) \quad (1.13)
\end{aligned}$$

There are some notable similarities and differences to the Brazilian case. First, the exchange rate effect on credit growth is much weaker than in Brazil - possibly reflecting more hedged private sector balance sheets which make them less sensitive to currency valuation effects - and it is also more imprecisely estimated. Second, like Brazil, credit plays a role in affecting absorption, although it is more like the growth in credit than the level, which is consistent with the fact that the dependent variable is the growth in absorption. Third, in terms of the output equation, a noticeable difference is that exchange rate effects are more than twice as strong for Chile and enter as rates of change (since the estimated coefficients on  $\tilde{z}_{t-1}$  and  $\tilde{z}_{t-2}$  are the same). This points to a higher elasticity of the trade balance to the real exchange rate in Chile. At the same time, the impact of domestic absorption on the real GDP gap is less strong, consistent with Chile being a much more open economy. There is also a direct effect of credit on inflation, and the coefficient on the output gap is now very significant statistically. As in Brazil, it is also clear that the exchange rate does play a non-trivial role in inflation, despite the general wisdom that exchange rate pass-through in Chile is lower. Indeed, the point estimate of -.05 on the real exchange rate change ( $\Delta \tilde{z}_t$ ) in the inflation equation suggests that a 10% nominal appreciation lowers CPI inflation by 50 basis points, all else

constant - and this is a non-negligible effect. Combining this estimate with the evidence (illustrated in the impulse-responses below) that the exchange rate is highly responsive to shocks to the domestic interest, it follows that interest rate shocks do have a sizeable effect on inflation, not just through the output gap effect but also via the exchange pass-through into domestic CPI. Turning to the interest rate equation, it resembles that for Brazil in having two lags of the interest rate as well as a significant positive response to expected inflation. The output gap has a stronger effect on interest rate settings than was true of Brazil, though neither is precisely estimated. It is notable that the coefficient on expected inflation is less than  $1 - \beta_{ii} - \gamma_{ii}$  and so the “Taylor principle” fails. There are two reasons why this does not lead to explosive inflation. One is that there is a separate exchange rate-induced effect on inflation in an open economy, and the other is that expectations are generated independently of the structural model. Finally, as in Brazil, deviations from UIP are significantly related to changes in the debt to GDP ratio  $\hat{d}_t$ , implying that fluctuations in net external debt have a significant impact on exchange rate dynamics, consistent with our earlier theoretical discussion.

### 5.2.1 Impulse Responses

It is first worth noting that if a recursive  $SVAR(2)$  was fitted to a standard ordering of variables  $\{\tilde{y}_t, \hat{\pi}_t, \hat{i}_t, \tilde{z}_t\}$  one would find that a rise in interest rates causes a rise in inflation and an initial depreciation in the exchange rate. This remains true if one expanded the system to the full set of variables  $\{pc_t, \tilde{n}_t, \tilde{y}_t, \hat{\pi}_t, \hat{i}_t, \tilde{z}_t, \tilde{d}_t\}$ . The structured VAR impulses given in Figure 3 are however very different and consistent with what one would get with standard New Keynesian models. One reason for the differences would seem to lie in the very strong exchange rate appreciation despite the lowish pass-through in Chile relative to other emerging market countries. At the heart of it is the very strong exchange rate response in Chile to the onshore-offshore interest rate differential: Figure 3 indicates that the real exchange rate response to monetary tightening is more than twice as high as that for Brazil, with the real exchange rate appreciating by over 6.7 percent at peak in response to a 100 basis points rise in the domestic policy rate, all else constant (compared with a 2% response in Brazil), despite the fact that the effect through the external debt term is similarly sized (0.28 vs. 0.29). So, while our single equation estimates shown above indicate the exchange rate pass-through into CPI in Chile is nearly three times as low as in Brazil, because the exchange rate is so responsive to onshore-offshore interest rate differentials, the effect of monetary policy on inflation working through the exchange rate channel is also quite strong in Chile.

A shock of a 1% increase in credit growth was applied to the model. The results are given in Figure 4. Again the results are similar to Brazil although exchange rate effects (a depreciation) are substantially stronger. Because the standard deviation of the estimated credit growth equation shock is around half of that for Brazil, performing one standard deviation shocks would make the results for the two countries comparable, though still stronger in Chile.<sup>18</sup> This seems consistent with evidence of a much greater banking sector penetration in Chile than in Brazil, as gauged by standard financial deepening indicators such as the ratio of bank credit to GDP (72% in Chile compared with 40% in Brazil by end-2008).

## 6 CONCLUSION

This paper has laid out a structural model of monetary transmission which incorporates key features of EMs in a manner parsimonious enough to be estimated with existing data and yet grounded on a DGSE theoretical skeleton. In particular, we have allowed for the role of a bank-dependent domestic sector and the impact of bank credit on aggregate demand and external aggregates which have not featured in previous studies. A SVAR representation of the model was derived and used to examine the Brazilian and Chilean experiences with full-fledged IT regimes since 1999. Differences in economic structure as well as in the track-record of economic policy making between the two countries - Brazil being far more closed to trade, less reliant on primary commodities, and with a more recent record of monetary and inflation stability - make such a comparative assessment of monetary transmission in the two countries particularly interesting. Such a diversity also provides a stricter test of the general validity of the skeleton model and our estimation approach.

Our SVAR estimates yield very sensible results for both countries. Indeed, no "price puzzles", "exchange rate puzzles" or any counter-intuitive results in impulses which are often found in previous VAR studies show up. This suggests that the proposal of a DGSE skeletal model as the basis of a structured VAR representation might provide a useful approach for examining monetary transmission in other EMs that are operating IT regimes.

A common finding is that, compared with evidence for advanced countries (notably the US), the transmission mechanism operates with shorter lags: the bulk of the effects on output and inflation take place within five quarters. This is arguably consistent with structural factors (e.g. shorter maturity of domestic credit) and the still considerable (albeit no longer as strong)

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<sup>18</sup> As with Brazil, the estimated strength of the credit channel in Chile is robust to dropping the debt accumulation equation from the model, so stands on its own relative to the open economy features of this model economy. The respective estimates are available from the authors upon request.

weight of the exchange rate and imported inflation in domestic currency pricing that is often mentioned in the literature.

As alluded to above, the exchange rate effects on disinflation are non-trivial. This is all the more interesting for Brazil as it is still a relatively closed economy to foreign trade with ratios of exports and imports to GDP below 15%. In both countries, there is a sizable effect on the exchange rate of domestic interest rate policy moves and net external debt accumulation has a significant bearing on deviations from UIP. This is consistent with structural models based on interest parity with an endogenous country risk premium. These have featured in the literature of other countries but have not been as prominent in previous work on Brazil and Chile. In this vein, the strong exchange rate response to such risk-adjusted interest rate differentials helps explain recent episodes of large real currency swings, as both net external debt and onshore-offshore interest rate differentials have varied widely in recent years.

Regarding to the role of bank credit in monetary transmission, our estimates indicate a non-trivial role for bank credit in monetary propagation. In both countries, there is evidence that policy interest rates changes affect credit growth and that the latter affects absorption. Moreover, at least in the case of Brazil, such a credit channel plays an intra-temporal role in moderating the impact of monetary policy shocks on absorption via exchange rate effects: while higher interest rates reduce absorption through the standard inter-temporal effect, it also boosts bank credit demand via a short-run exchange rate appreciation that monetary tightening typically entails. The attendant balance sheet and/or wealth effects arising from such currency appreciations (particularly for non-tradable producers which tend to be more dependent on bank credit) thus mitigates the otherwise standard contractionary effect that monetary tightening has on absorption. Even though the contractionary effect wins out in the aggregate, it appears that such an effect is somewhat mitigated by the intra-temporal exchange rate effect. We also find an independent role for credit shocks (which may just reflect changes in reserve requirements and other regulations, as well shocks to intermediation efficiency). Our estimates suggest that there are non-trivial effects on output and inflation in both countries, although these are reasonably short-lived, and particularly so in the case of Brazil.

An obvious practical implication is that policies that affect bank credit directly have an important bearing on output and inflation in both countries, at least in the short-run. This may incidentally help explain the relative shallowness of the recent financial crisis in both countries, despite the sheer size of the external adverse shock to these countries' terms of trade, trading partners income, and country risk. As counter-cyclical credit policies in both countries were far-reaching, this mitigated the fall in absorption and prevented a bank crisis which – given the significance of the estimated impact on absorption - would have greatly added to the contractionary impact of the external shock.



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Figure 1: Brazil: Impulse Responses to a 1 % Rise in The Annualized Interest Rate

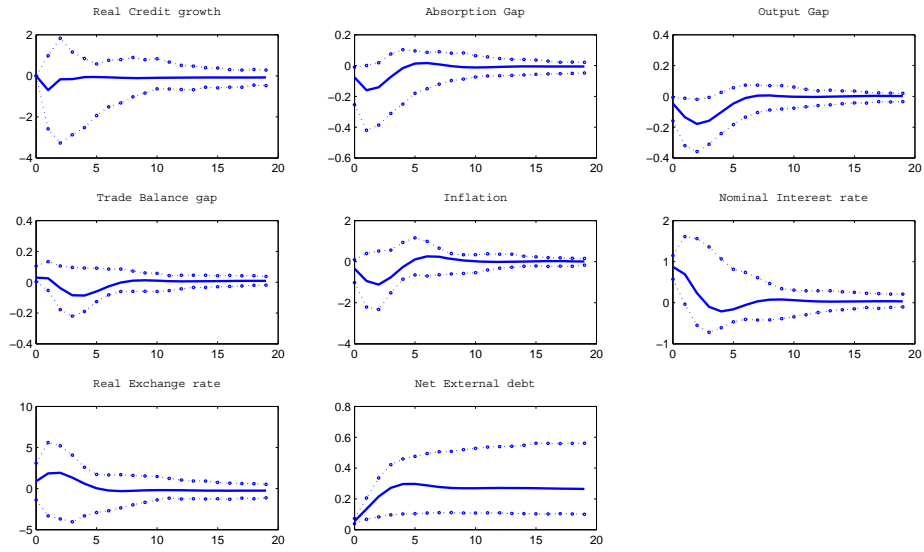


Figure 2: Brazil: Impulse Responses to a 1 % Rise in Credit Growth

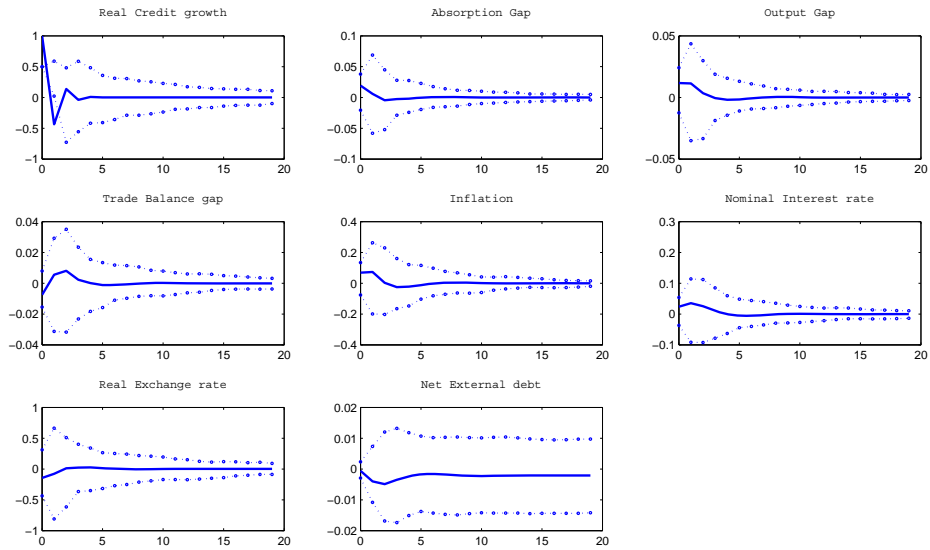


Figure 3: Chile: Impulse Responses to a 1 % Rise in The Annualized Interest Rate

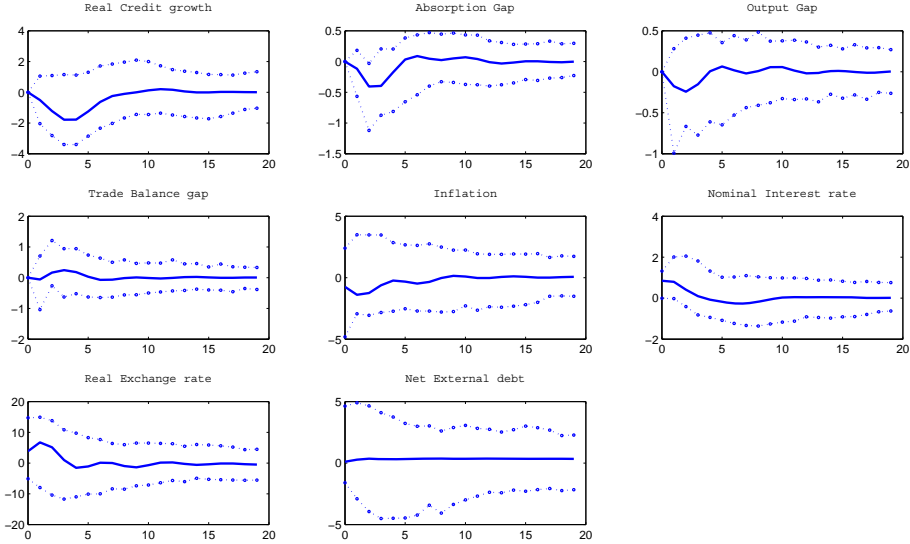
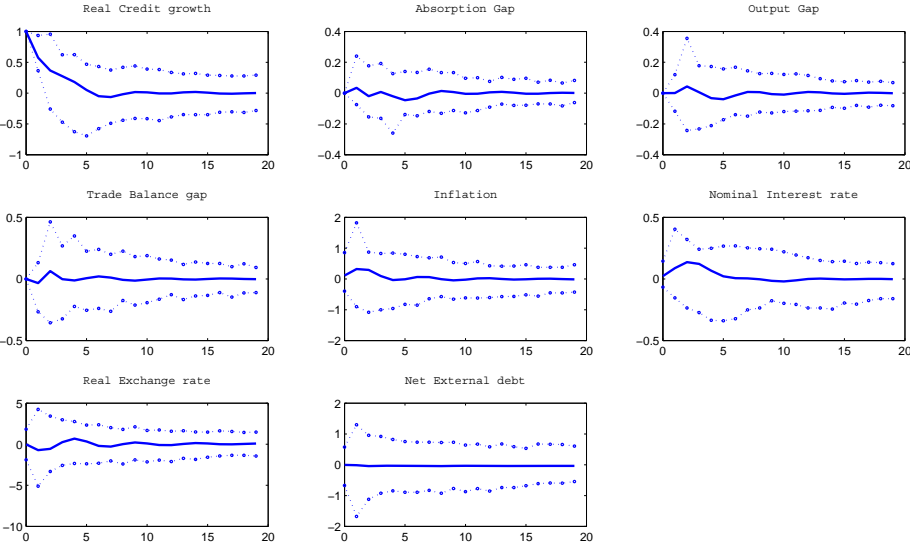


Figure 4: Chile: Impulse Responses to a 1 % Rise in Credit Growth



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