THE CREDIT CHANNEL AND MONETARY TRANSMISSION IN BRAZIL AND CHILE: A STRUCTURED VAR APPROACH

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The widespread adoption of inflation-targeting regimes by emerging market 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 is a paucity of empirical research for emerging markets relative to the large literature on advanced countries, partly reflecting shorter time series and other problems not typically faced in studies of the latter.[1] A few recent studies fit standard dynamic stochastic general equilibrium (DSGE) models to emerging market data (for example, Furlani, Portugal, and Laurini, 2008; da Silveira, 2008; Del Negro

and Schorfheide, 2008), but they largely ignore some key structural features of emerging markets in the chosen specification. Moreover, the Bayesian methods used for estimation in these studies often impose strong priors, so that the empirical investigation is less about discovery than about quantifying the parameters of some prescribed model. This is not to deny that DSGE models are useful for thinking about interrelationships in the macroeconomy. Nevertheless, they are often best used as a source of structural information that provides a skeleton with which investigators can organize the data, rather than imposing the model on the data, at least until one is sure that it is a good representation of the data. Often the only way DSGE models are judged is by comparing the results to a vector autoregression (VAR), but this is unlikely to be a very powerful test. Simple checks, such as whether the model's assumptions about expectations and shocks are consistent with the data, are far more likely to reveal deficiencies in the specification.

Our objective in this paper is to develop a model that uses a particular DSGE model (namely, the New Keynesian model) as a skeleton and then to expand it so as to resemble a structural VAR (SVAR). Unlike existing SVARs that either force the system to be recursive (or ordered) or 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 the model, we thereby produce a nonrecursive SVAR, which forms the basis of our 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 show in previous work (Catão, Laxton, and Pagan, 2008), the inclusion of an external debt accumulation equation in the structured VAR model not only is of interest in its own right—as it permits the tracking of the effects of monetary shocks through key external aggregates—but also imposes some stock-flow dynamics on the model that allow 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 what can be 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 emerging market macroeconomies. 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), which is particularly relevant in emerging markets (Edwards and Végh, 1997; Catão and Chang, 2010). The effect of shocks to banks' lending capacity—arising from, say, exogenous changes in reserve requirements or banking intermediation technology—on output, inflation, and other aggregates can also be traced out in the model. Our structural VAR model is also designed to retain one of the important features emphasized in the DSGE perspective, namely, the integration of stocks and flows. This is rarely addressed in standard SVARs.

We empirically implement the modeling strategy 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). Chile actually started targeting inflation in the early 1990s, but it operated a system of exchange rate bands through 1998, so targeting inflation was not the overriding goal of monetary policy. Moreover, an advantage of restricting the estimation period to 1999–2009 is that we are able to use the same sample for both countries, which facilitates 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 achieving a relatively parsimonious specification, so as 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, Laxton, and Pagan (2008), in that we replace the recursive SVAR used there with a nonrecursive SVAR. Substantial differences emerge between the conclusions reached with a traditional recursive VAR and those from this paper's approach. Despite the relatively short time span available for estimation, our structured VAR estimates do not generate price puzzles, exchange rate puzzles, or other anomalies that abound in the literature, which would be found for both countries under a standard recursive SVAR.

The main results are as follows. First, the transmission mechanism works faster in Brazil and Chile than in the United States and other advanced countries, 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. This is often found in small open economies such as Australia.

Second, the bank credit channel plays a nontrivial 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 intertemporal effect of monetary policy changes on absorption. The other is an intratemporal 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 or when the nontradables sector of the economy is more bank-dependent than its tradables counterpart, implying that the overall demand for bank credit will tend to increase as relative prices shift toward nontradable goods producers. This combination of the balance sheet effects of currency mismatches and the greater bank dependence of domestic firms implies that monetary policy will have nontrivial effects on bank lending and hence on absorption. Our estimates indicate that while the intertemporal channel eventually wins out, so that monetary tightening (loosening) depresses (boosts) bank credit, the intratemporal channel appears to play an offsetting role.

Third, the quantitative impact of credit shocks tends to be larger for Chile than Brazil. While neither is large in response to a 1 percent 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, credit grew strongly in both countries, with standard deviations in credit shocks of around 9 percent in Brazil and 5 percent in Chile. So, although the impact of a 1 percent change in credit on inflation and output is relatively small, such large variations in actual credit growth might suggest that these developments have been important for macroeconomic outcomes. For Brazil, the impact of a positive 9 percent 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 percent shock to credit growth is equivalent to a decrease of around 100 basis points in interest rates.

The remainder of the paper is divided into five sections. Section 1 reviews the existing evidence on the monetary transmission mechanism in Brazil and Chile and provides a motivation for the model and results. Section 2 lays out the methodology, first in a general way and then in the context of the structural model that is used as the skeleton for our SVAR. Section 3 provides a discussion of the data, including the construction of output and absorption gaps. Section 4 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.

1. EXISTING EVIDENCE

The introduction of the inflation-targeting framework in Brazil in 1999 generated significant interest in understanding the monetary transmission mechanism. As a result, a growing literature seeks to identify and measure the channels through which the central bank's policy interest rate (SELIC) affects output and inflation. Bogdanski, Tombini, and Werlang (2000) describe some of the channels and discuss the central bank's model, and their framework forms 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 subsamples: the "moderate" inflation period (1975–85); the high inflation period (1985–94); and the low inflation regime (after 1994). 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 the 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.

Other studies acknowledge the role of the exchange rate as a determinant of Brazilian inflation. Bevilaqua, Mesquita, and Minella (2007) find that the large appreciation of the real since 2005 has contributed significantly to the fall in inflation. Favero and Giavazzi (2004) conclude that exchange rate movements affect inflation expectations and, through this channel, the central bank interest rate setting. This suggests not only that the exchange rate may affect current inflation by changing the cost of imported goods, but also that there may be an important expectational channel at work.

Some attention has also been devoted to the interest rate reaction function. Using a Hodrick-Prescott (HP) filter to produce the output gap measure, Minella (2003) finds that the parameter on the output gap in the monetary policy reaction function has the wrong sign and is not statistically significant from zero. He argues that this could arise because of simultaneity bias caused by supply shocks that depress the output gap and raise inflation. The same study also finds that exchange rate volatility has been an important source of inflation variability in Brazil, based on a smaller VAR estimated on monthly data but with a sample that includes the pre-inflation-targeting period (1994–2002).

Da Silveira (2008) and Furlani, Portugal, and Laurini (2008) reexamine some of these issues from the perspective of a New Keynesian open economy DSGE model derived from Galí and Monacelli (2005), who use Bayesian techniques to estimate their parameters. Given their open economy set-ups, both studies have the exchange rate playing a key role in the transmission of monetary shocks via uncovered interest parity, though neither of them contemplates a similar role for the country risk premium, as we do below. Furthermore, because all goods in the Galí-Monacelli set-up are tradables, changes in the real exchange rate are proportional to changes in the terms of trade. Da Silveira (2008), in particular, finds that monetary policy lowers inflation via a strong nominal exchange rate appreciation, but the effects are not particularly strong and they are reinforced through the effect of monetary policy on the output gap. Furlani, Portugal, and Laurini (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 rather mostly reflects inflation developments and, to some extent, the output gap. Both studies also find, as we do, that shock accommodation is relatively swift,

though their models do not allow for a bank credit channel, which our estimates identify as important.

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 DSGE modeling and Bayesian estimation. Early work in the VAR tradition includes Morandé and Schmidt-Hebbel (1997), Valdés (1998), Calvo and Mendoza (1999), and Cabrera and Lagos (2002). While some of these studies impose structural restrictions, they tend to rely strongly on atheoretic identifying assumptions and build a weak link between the estimated VAR and a theoretically based structural model. Not surprisingly, a number of puzzles emerged in this literature, including price, exchange rate, and liquidity puzzles (see Chumacero, 2003, for further discussion).

Much of the recent work is based on the small open economy model with Keynesian features set out in Galí 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. When the authors compute 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 (that is, the lower the degree of the central bank's credibility), the slower is the fall in inflation to a given monetary tightening, along with a higher real exchange rate appreciation and sacrifice ratio. They complement this calibration exercise with generalized method of moments (GMM) estimates of the monetary policy rule over the pre-1999 and post-1999 periods. They find that in the full-fledged inflation-targeting regime, monetary policy has become more forward-looking (that is, more responsive to expected future inflation than current inflation) and the coefficient on the deviations of inflation from target in the monetary policy rule has risen (rather than fallen).

Caputo, Liendo, and Medina (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 inflation-targeting 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) implement a similar strategy to ours in that they estimate a DSGE model (a version of the Galí-Monacelli model) to derive predictions of what the VAR coefficients (Π) would be (Π^*). A prior distribution for Π is then constructed by centering on Π^* and having a covariance matrix that is proportional (through the inverse of a hyperparameter, λ) to the form of the covariance matrix of $\hat{\Pi}_{OLS}$. The value of λ 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 Π. As λ 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 λ basically enables the analyst to explore how sensitive the conclusions will be to the choice of using the DSGE model versus the VAR. One might choose λ by reference to predictive success and then 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, so it really needs to be correctly specified. The difficulties are less serious for monetary policy shocks, however, as the structural equation defining this is atheoretic.

A first question addressed by Del Negro and Schorfheide (2008) is the extent to which the central bank responds to the terms of trade and exchange rate fluctuations relative to inflation. Similarly to Caputo, Liendo, and Medina (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 consumer price index (CPI) inflation, which

implies that the shocks that 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 (that is, with an estimated λ). 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, although the impact on inflation is 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, Liendo, and Medina, 2007) an integration of external debt stock, current account flows and exchange rate dynamics, as we do below.

2. METHODOLOGY

Let $\mathbf{z}_t$ be an $n \times 1$ vector of variables in the macroeconomic system. A typical structural equation in a macroeconomic model (normalized on the first variable, z_{1t}) has the following form:

$$z_{1t} = \mathbf{z}_t^{-'}\boldsymbol{\alpha}_1 + \mathbf{z}_{t-1}'\boldsymbol{\delta}_1 + E_t(\mathbf{z}_{t+1})'\boldsymbol{\gamma}_1 + \varepsilon_{1t},$$

where ε_{1t} is a structural shock and $\mathbf{z}_t^-$ is $\mathbf{z}_t$ less z_{1t}. One approach to modeling these systems is to use 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) find that the structural equation meant to be identified as a supply curve was influenced positively by nominal money growth.

In standard SVARs, $\boldsymbol{\gamma}_1 = 0$ and various restrictions are placed on $\boldsymbol{\alpha}_1$ (mainly exclusion restrictions) in order to identify the shocks, ε_{1t}. In particular, the system is assumed to be recursive (or 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) on α_1, δ_1 and γ_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 that is then augmented (if necessary) to yield 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 $E_t(\mathbf{z}_{t+1})$ in the structural equations of the former. An approach that does not use any particular theory-oriented model, and is more robust to model specification error, is to make the expectation a function of some model variables, ξ_t. Then if one regresses $\mathbf{z}_{t+1}$ against ξ_t to get coefficient estimates $\hat{\mathbf{A}}_1$, $E_t(\mathbf{z}_{t+1})$ could be measured as the combination $\mathbf{w}_t = \hat{\mathbf{A}}_1 \xi_t$.[2] This approach was used in the FRB/US model (see Brayton and others, 1997), although the variables ξ_t were only a few of those in the FRB/US system. Since ξ_t generally involves 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

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

where n_t is consumption expenditure, i_t is a nominal interest rate, π_t is an inflation rate and ε_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 in the structural equation have been augmented by ξ_t and ξ_{t-1}. What is critical, though, is that ϕ_1 and ϕ_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.

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2. If direct measures of expectations were available, they could be used.

Our strategy is thus 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). An important part of our strategy is the skeletal model that forms the core of our SVAR, for which we use a relatively standard New Keynesian model set out in the next subsection.

2.1 The Skeletal Structure for our Structural VAR

Our starting point for structuring the VAR is a canonical small macroeconomic model that has been used quite extensively in the macroeconomics literature and has often been deployed for analysis at the International Monetary Fund (IMF) and various central banks (see 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 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 distinguishes between absorption (n_t) and output (y_t). Later we use a convention that a coefficient α_{xy} shows a contemporaneous effect between x_t and y_t, β_{xy} shows the effect between x_t and y_{t-1}, and γ_{xy} is between x_t and y_{t-2}. Some license is taken when expectations are involved. Thus α_{ne} is the coefficient in the absorption equation that connects n_t and the expected value $E_t(n_{t+1})$. Thus, the model can be written as follows:[3]

$$\tilde{n}_t = \alpha_{nn^e} E_t(\tilde{n}_{t+1}) + (1 - \alpha_{nn^e})\tilde{n}_{t-1} + \beta_{nr}\hat{r}_{t-1} + \varepsilon_t^n; \tag{1}$$

$$\tilde{y}_t - \tilde{n}_t = \alpha_{yz}\tilde{z}_t + \alpha_{yy^*}\tilde{y}_t^* + \delta_{yn}\tilde{n}_t + \varepsilon_t^y; \tag{2}$$

$$\hat{\pi}_t = \alpha_{\pi\pi^e} E_t(\hat{\pi}_{t+1}) + (1 - \alpha_{\pi\pi^e})\hat{\pi}_{t-1} + \alpha_{\pi y}\tilde{y}_t - \alpha_{\pi z}\Delta\tilde{z}_t + \varepsilon_t^\pi; \tag{3}$$

$$\hat{\imath}_t = \beta_{ii}\hat{\imath}_{t-1} + \alpha_{i\pi^e} E_t\hat{\pi}_{t+1} + \alpha_{iy}\tilde{y}_t + \varepsilon_t^i; \tag{4}$$

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3. A tilde (~) indicates a log deviation from equilibrium values; a hat (^) indicates a deviation in levels.

$$\tilde{z}_t - E_t \tilde{z}_{t+1} - (\hat{r}_t - \hat{r}_t^*) = \zeta_t + \varepsilon_t^z; \tag{5}$$

$$\Delta \hat{d}_t = \overline{d}\left[(1 - \overline{\psi})\,\hat{\imath}_t - (1 + \overline{\imath})\hat{\pi}_t\right] + \omega_n \tilde{n}_t - \tilde{y}_t + (\omega_m - \omega_x)\tilde{z}_t - \omega_x \hat{tt}_t; \tag{6}$$

$$\overline{\psi} = \hat{\imath}_t = \overline{\pi} + \Delta \ln \overline{Y}_t \tag{7}$$

$$\hat{r}_t = \hat{\imath}_t - E_t(\pi_{t+1} - \overline{\pi}_{t+1}). \tag{8}$$

The first equation provides a specification for the log of domestic absorption $(\tilde{n}_t)$, 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 on where one sees the system heading and the path of adjustment to that point. Most modern macroeconomic models can be written as gap models, so the approach is fairly flexible. The equilibrium value may be constant or time varying.[4] In this case, the absorption gap depends on 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 $\overline{\pi}_t$, so we work with the real interest rate adjusted for the inflation target, $\hat{r}_t$, given in equation (8). Most empirical work with the New Keynesian model incorporates the target 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

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4. 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. We discuss these transformations in detail below, 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 in question. 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 reasonable one's estimation results appear to be. A striking example in the empirical macroeconomic 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 specific country-group features in the estimation strategy. See Catão and Terrones (2005).

it is rather weak. We have therefore chosen to treat the target as exogenous. No other variable in this model determines the level of absorption, which is consistent with the standard Euler equation for consumption in DSGE models.[5] 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 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. The proposed specification also implicitly captures accelerator effects through the lagged terms on absorption, 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 ($\tilde{n}_t$), and import and export gaps. For Brazil the import and export shares are largely the same, so $\tilde{y}_t - \tilde{n}_t$ can be regarded as the log deviation of the current account from zero. Imports are determined by total expenditure and the real exchange rate, while exports are related to the real exchange rate and foreign expenditure. We thus simply eliminate imports and exports to produce a relation linking the output and domestic absorption gaps, the log of the real exchange rate 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 this needs to be investigated further since it is not derived from any theoretical framework. In fact, delays between orders and deliveries may cause lags.

Although this equation is fundamentally an identity, it ceases to be so when we replace exports and imports by a functional form. Hence, we add a shock to allow for this. 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

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5. As discussed in the next section, we augment this canonical specification to include the role of domestic interest spreads (through an interest wedge that arises in models with deposit) and credit-in-advance constraints (see, for example, Edwards and Végh, 1997). Since the domestic interest spread is itself a function of the policy interest rate and 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.

import share, one will probably want some of the observed changes 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. It includes the output gap and the exchange rate gap.[6] As suggested by previous studies, the exchange rate plays a very significant role in influencing the price of tradeables and CPI inflation.

The fourth equation is a monetary policy reaction function, where $\hat{\imath}_t$ is defined as the nominal interest rate less the target inflation rate. The parameter β_{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 are no exceptions in this regard). In light of evidence from existing studies reviewed earlier, we do not include the exchange rate in the monetary policy rule. In our background empirical work, we tested whether the exchange rate should be present, and the results showed only a very weak dependence. We discuss this below for Brazil.

The exchange rate equation (equation 5) is risk-adjusted UIP, where $\hat{r}_t^*$ is the external real interest rate (proxied by the interest rate on U.S. three-month Treasury bills). The exchange rate is defined such that a rise represents an appreciation. There are two shocks in this equation. One, ζ_t, is a risk premium that can be considered as relating to model variables, while the other, ε_t^z, is a function of nonmodel variables and is treated as white noise and as uncorrelated with ζ_t. In many real business cycle models of small open economies, this risk premium shock is made a function of the level of net foreign debt relative to GDP (again measured relative to a steady-state value) (see, for example, Schmitt-Grohé and Uribe, 2003). Other factors may play a role, however, such as the level of domestic interest rates. Much of domestically issued debt in these countries is held by foreigners, and 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

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6. We have constrained the coefficients in this equation so that they add to one. In many models, they add to the discount factor. This is often around 0.99 in a quarterly model, so we follow a common practice that enforces a restriction that the coefficients sum to unity, which aids identification.

standard measures of country risk such as the J. P. Morgan Emerging Markets Bond Index.

The debt equation (equation 6) 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, Laxton, and Pagan (2008).[7] In this equation, the terms of trade ($\hat{tt}_t$) are taken as exogenously given—which is a reasonable assumption for a small open economy. The variables ω_m, ω_x, and ω_n are the import, export, and absorption shares in GDP, respectively, $\overline{d}$ is the steady-state ratio of net foreign debt to potential GDP, and ψ_t is the nominal potential GDP growth rate ($\Delta \ln \overline{Y}_t$). 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.

2.2 Augmenting the Skeletal Structure

As mentioned in the previous section, our first step in extending the skeletal structure is to replace the expectations $E_t \tilde{n}_{t+1}$, $E_t \tilde{\pi}_{t+1}$, and $E_t \tilde{z}_{t+1}$ 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).[8] 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, that is, a VAR(2) in the skeletal model variables is fitted and used to produce $E_t \tilde{n}_{t+1}$, $E_t \tilde{\pi}_{t+1}$, and $E_t \tilde{z}_{t+1}$. Because there are

7. In its application to the present paper, we have chosen to 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.

8. The solution is a VARX model with two lags in the endogenous model variables and no lags in the exogenous external variables ($\tilde{y}_t^*$, $\hat{r}_t^*$, and $\hat{tt}_t$). 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 on their own past history.

sixteen parameters in this regression, and our data sets are only forty observations long, we deleted 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 earlier. We have already augmented the standard New Keynesian 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 businesses) are still largely dependent on domestic banks for funding, and they have limited access to international capital markets. This channel might thus be expected to play a role, and 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 equation (1). 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.[9] 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 because of a sharp rise over the sample period. For example, in the case of Brazil, after being fairly stable around 0.24 from 1999 to 2004, it rose sharply to 0.44

9. The specific way in which Edwards and Végh (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. This directly translates into a functional relationship between bank spreads and the credit-to-expenditure ratio, since deposit-in-advance constraints imply that deposits are proportional to expenditure.

at the end of the period. Hence, it behaves more like an integrated series over our data period.[10]

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. Thus, the gap between the log of the credit-to-GDP ratio and its normal level would be a function of the growth of credit less the potential growth rate of GDP. Given that the latter is reasonably constant, this suggests that our measure captures the ideas about excess credit reasonably well, but it is statistically tractable since it is a stationary process.[11]

At any point in time, domestic bank credit is endogenously determined within a system, so we need to decide how to account for this within an SVAR. To allow for contemporary effects of credit expansion on expenditure, we place the excess credit variable before any other variable of the system, that is, 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 explanatory variables entering the excess credit equation to 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 Mejía (2004).[12] Second, the nontradables sector may be more bank dependent than the tradables sector (see Catão and Chang, 2010). Both structural features imply that a real exchange rate appreciation will increase real credit demand in the nontradables 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

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10. The estimated AR(1) coefficient is 1.02.

11. Allowing for the presence of an autonomous component in the excess credit variable that 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. BNDES's lending policies and rates arguably respond to other incentives, and its lending rates are typically below market rates.

12. While the dollarization of private sector liabilities was not nearly as extensive in Brazil as in many other emerging markets, it is far from negligible. Foreign-currency-denominated debt rose from very low levels in the early 1990s to 40 percent of total corporate debt in 2002 (Bonomo, Martins, and Pinto, 2003, table A2). Using a large panel of firm-level data, Bonomo, Martins, and Pinto (2003) also find that the balance sheet effects of currency movements have a significant impact on credit demand and investment.

real exchange rate on domestic credit. Section 4 provides supportive econometric evidence that real exchange rate appreciations tend to foster domestic credit growth in Brazil but not 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 higher-order lags was 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 the coefficients of the two lags often appear with opposite signs, a term such as $az_{t-1} + bz_{t-2}$ (with $a > 0$, $b < 0$, and $|b| < a$) can be written as $(a+b)z_{t-1} - b\Delta z_{t-1}$, that is, there is both a level and a growth rate effect.

We estimated all the equations with instrumental variables. This was done partly to avoid the fact that systems estimation requires that all the equations in the system be correctly specified 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. The following rules governed the selection of instruments. First, any exogenous or lagged variable appearing in an equation is 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, we used the residual from the credit equation 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 with one another is valid. If the number of instruments equaled the number of variables in each equation, and if 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—for example, the correlation of ε_t^n with $\tilde{n}_t$ is 0.41 (Brazil) and 0.88 (Chile). In other instances, we might expect that the conditional expectation would be a more powerful instrument. After estimation, we checked whether 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 α_{xy} shows a contemporaneous effect between x_t and y_t, β_{xy} shows the effect

between x_t and y_{t-1}, and γ_{xy} is between x_t and y_{t-2}. Thus, the credit growth equation for Brazil could be written in the form

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

while the absorption equation might be written as

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

For the equations generating expectations, we add a superscript *e* to the coefficients. Hence, we have

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

3. DATA

Our estimation data come from readily available official statistics for both countries. After describing these data sources and discussing some punctual issues regarding the choice of indicators for their theoretical counterparts, this section lays out and discusses some of the underlying methodological issues underpinning our estimates of output and absorption gap measures for each country.

3.1 Brazil

We restricted our sample to the inflation-targeting period as a response to evidence that large structural changes occurred around the point of its introduction. In particular, Tombini and Lago Alves (2006) find that inflation dynamics and exchange rate pass-through in Brazil recorded significant structural changes before and after 1999, while Minella (2003) reports far-reaching changes in the price indexation system and inflation dynamics after 1995.

Seasonally adjusted national income account data were taken from the IMF's *International Financial Statistics* (IFS) and the Brazilian Institute of Applied Economic Research (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 partners using CPI deflators and 2000 weights. The indicator of real world income was computed as

the trade-weighted average of real GDP of the country's main trading partners, which account for over 80 percent of Brazil's trade.

The inflation variable is seasonally adjusted CPI including all items—with both administered and free market prices. While it is 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 in 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 is perhaps irrelevant if the task at hand is indeed to model aggregate inflation. First, administered prices still respond to demand pressures, albeit with a longer lag, because of backward-looking indexation clauses in the underlying concession contracts. Second, although utility prices are typically key inputs to free market prices, the interaction between the two is certainly complex, and simply including both series in a VAR is unlikely to address such complexity. Third, 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 full CPI.

To parameterize equation (6), we need ω_m, ω_x, and $\bar{d}$. These were replaced by the average ratios of imports, exports, and net debt to GDP over 1999–2009. Because ω_m and ω_x were virtually the same over this period, we fix them both at 0.11. We used the average growth rate of real GDP over the period and target inflation (4.5 percent in recent years) to compute $\bar{\psi}$. This makes it 1.9 percent per quarter.

3.2 Chile

As with Brazil, we restrict our sample to the post-1999 period and use quarterly data throughout. Seasonally adjusted national income data were 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 in the case of Brazil, the indicator of real world income was computed as the trade-weighted average of real GDP of the

country's main trading partners, which account for over 90 percent of Chile's foreign trade.

To parameterize equation (6), we need ω_m, ω_x, and $\bar{d}$. The first two were replaced by taking the simple average of ω_m and ω_x over 1999–2009. The debt ratio was the historical average over this period. Likewise, we used the average growth rate of real GDP (around 4 percent) and target inflation (3 percent) to compute $\bar{\psi}$. This makes it 1.95 percent per quarter.

3.3 Producing Gap Measures for Both Countries

Measures of output and absorption gaps are present in the skeletal model we use, so we need to estimate them. In much of the existing literature, the permanent component is extracted with the Hodrick-Prescott (HP) filter. As Harvey and Jaeger (1993) point out, however, 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:

$$z_t = P_t + T_t;$$

$$\Delta^2 P_t = v_t;$$

$$T_t = u_t;$$

$$\lambda = \frac{\text{var}(u_t)}{\text{var}(v_t)}.$$

The model clearly implies that

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

where e_t is an uncorrelated process. Setting $\lambda = 1{,}600$, we find that $\alpha_1 = -1.77$, $\alpha_2 = 0.8$. Fitting this model to Brazilian GDP data over 1999:1–2009:1, we get $\alpha_1 = -0.95$, $\alpha_2 = -0.05$. Of course this process has a common unit root to the moving average and autoregressive (AR) parts which cancels, implying that the log of Brazilian GDP is an I(1) process, which contrasts with the I(2) model implied by the HP filter.

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

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

such that the transitory component is $z_t - P_t = -E_t\sum_{j=1}^{\infty}\Delta z_{t+j}$, and this is the gap. We thus need to prescribe a model for Δz_t to be able to compute the transitory component. When Δz_t is an AR(p), $E_t\sum_{j=1}^{\infty}\Delta z_{t+j}$ will be a linear function of $\Delta z_t, \Delta z_{t-1}, ..., \Delta z_{t-p+1}$. In that case, the BN measure of the output gap is constructed as the negative of an average of growth rates. This means that one will see a negative relation between the output gap and growth, so that regressing inflation against output growth should produce a negative coefficient on the latter.

One criticism of the BN filter is that the resulting output gap estimate is not as smooth as the HP-filtered estimate. This may well be true when a low-order AR process is used to approximate Δy_t, but a higher-order AR often produces much smoother results (see, for example, Morley, 2007). The intuition is that the gap is constructed by averaging growth rates, which generally results 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, that is, it is I(2), and the other is that it is a two-sided filter that uses weighted averages of growth rates in both the past and the future. To demonstrate this, we applied an HP filter to the Brazilian log of GDP series and then regressed this against three lagged and forward values of GDP growth. That produced a regression of the form

$$\tilde{y}_t^{HP} = 0.45\Delta y_t + 0.28\Delta y_{t-1} + 0.17\Delta y_{t-2} + 0.20\Delta y_{t-3}$$
$$- 0.39\Delta y_{t+1} - 0.28\Delta y_{t+2} - 0.30\Delta y_{t+3}.$$

While 75 percent of the variation in $\tilde{y}_t^{HP}$ is explained by these variables, only 33 percent is due to the lagged and current values of Δy_t. The BN filter is therefore unlikely to approximate the HP filter 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. Laxton, Shoom, and Tetlow (1992) perform a simulation experiment in which the potential level of output actually followed an I(1) process. They find that using the output gap from an HP filter produces an estimate for the parameter on the output gap in a Phillips curve that is well below the true value used in producing the simulated data.

In each case, the BN-filtered output and absorption gaps were determined by fitting an AR(4) to growth rates over the sample period. We ran into difficulties measuring the foreign output gap, however, in that the appropriate series (the trade-weighted GDP of Brazil's and Chile's trading partners) have extremely persistent growth rates: fitting an AR(1) to Brazilian data for 1999:1–2009:1 yielded a point estimate of the AR(1) coefficient of 0.99. Hence, the BN filter is not appropriate in this case, whereas 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.

4. MODEL ESTIMATES

This section presents the estimated structural 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 percentage point rise in the growth rate of credit. These increases are relative to the steady-state levels, and all variables in the model equations are intended to be measured that way. Thus, what is being explained is the nature of the adjustment back to equilibrium. There may be forces here that are not present in equilibrium (for example, nominal interest rates may affect disequilibrium expenditure), but in equilibrium we expect that expenditure will be governed by the real interest rate.

As we observed in the introduction, papers based on SVAR models rarely present the structure, but rather only report the impulse responses. One reason for this is that it is quite possible to have reasonable impulse responses resulting from what might appear to be odd-looking structural equations. But there are three compelling

reasons to present the structural equations of our model. First, it reveals how the skeletal model needs to be modified to fit the data. This provides some information for those who wish to work with a theoretical model that is close to our skeletal one, and quite a few papers use variants of our skeletal model for policy analysis in emerging economies.

Second, it is sometimes useful to refer back to the structural equations when seeking an explanation for either the pattern or the magnitude of the 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 our small sample sizes, we cannot precisely determine the values of these parameters, so it is useful to be able to assess how sensitive our conclusions are to the point estimates of the structural equations used in constructing the impulse responses. This is a central theme in Del Negro and Schorfheide (2008).

Third, a principle of full disclosure seems desirable in empirical work. This would seem to demand the provision of information on the structural equations, even though this is rarely done.

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

4.1 Brazil

The structural VAR that was fitted is given below, with *t* ratios in parentheses:

$$pc_t = \underset{(-2.37)}{-0.35}\, pc_{t-1} + \underset{(2.18)}{2.67}\, \tilde{y}_{t-1} - \underset{(-2.19)}{1.26}\, \hat{\imath}_{t-1} + \underset{(3.44)}{0.603}\, \tilde{z}_{t-1} - \underset{(-1.71)}{0.326}\, \tilde{z}_{t-2} + \varepsilon_t^c;$$

$$\tilde{n}_t - \tilde{n}_{t-1} = \underset{(12.86)}{0.61}\big[E_t(\tilde{n}_{t+1}) - \tilde{n}_{t-1}\big] - \underset{(-2.66)}{0.025}\, \hat{r}_{t-1} + \underset{(1.22)}{0.004}\, pc_t + \underset{(2.29)}{0.006}\, pc_{t-1} + \varepsilon_t^n;$$

$$\tilde{y}_t = \underset{(1.92)}{0.61}\, \tilde{n}_t - \underset{(-1.92)}{0.02}\, \tilde{z}_{t-1} + \underset{(2.92)}{0.42}\, \tilde{y}_{t-1} + \varepsilon_t^y;$$

$$\hat{\pi}_t - \hat{\pi}_{t-1} = \underset{(4.86)}{0.58}\big[E_t(\hat{\pi}_{t+1}) - \hat{\pi}_{t-1}\big] + \underset{(0.84)}{0.79}\, \tilde{n}_t - \underset{(-3.01)}{0.17}\, \tilde{z}_{t-1} + \underset{(2.49)}{0.15}\, \tilde{z}_{t-2} + \varepsilon_t^\pi;$$

$$\hat{\imath}_t = \underset{(10.48)}{1.08}\, \hat{\imath}_{t-1} + \underset{(3.20)}{0.26}\, E_t\hat{\pi}_{t+1} - \underset{(-3.93)}{0.32}\, \hat{\imath}_{t-2} - \underset{(-1.54)}{0.02}\, \tilde{z}_{t-1} + \varepsilon_t^i;$$

$$\zeta_t = \underset{(1.10)}{0.27}(\tilde{z}_{t-1} - E_{t-1}\tilde{z}_t) - \underset{(2.80)}{0.43}\hat{\imath}_{t-1} + \underset{(2.90)}{0.56}\hat{r}^*_{t-1} + \underset{(2.10)}{0.28}\hat{d}_{t-1}$$

$$+ \underset{(1.93)}{0.17}\tilde{z}_{t-1} + \underset{(0.47)}{0.08}E_{t-1}\hat{\pi}_t + \varepsilon^z_t;$$

$$\zeta_t = \tilde{z}_t - E_t\tilde{z}_{t+1} - \left[\frac{(\hat{r}_t - \hat{r}^*_t)}{4}\right].$$

The equations generating expectations are as follows:

$$E_t\hat{\pi}_{t+1} = \underset{(2.89)}{0.38}\hat{\pi}_t - \underset{(-4.58)}{0.28}\tilde{z}_t + \underset{(3.72)}{0.22}\tilde{z}_{t-1} + \underset{(1.74)}{1.29}\tilde{n}_t;$$

$$E_t\tilde{n}_{t+1} = -\underset{(-2.46)}{0.07}\hat{\imath}_t + \underset{(2.25)}{0.016}\tilde{z}_t - \underset{(-2.67)}{0.02}\tilde{z}_{t-1} + \underset{(6.90)}{1.02}\tilde{n}_t - \underset{(-5.83)}{0.70}\tilde{n}_{t-1} + \underset{(1.65)}{0.01}pc_t;$$

$$E_t\tilde{z}_{t+1} = \underset{(1.79)}{0.26}\tilde{z}_t - \underset{(-0.95)}{0.91}\tilde{y}_t - \underset{(-2.07)}{4.12}\tilde{n}_t + \underset{(3.76)}{6.35}\tilde{y}^*_t - \underset{(-4.21)}{0.80}\hat{d}_t.$$

In constructing these equations, we tried to err on the side of caution given the small number of available observations. Consequently, we generally used 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 the Akaike information criterion (AIC) to decide whether a regressor should be retained. In some instances, we introduced a variable into the model even though it was not significant. If the variable was supposed to be present in the skeletal model, then the lack of significance represents evidence in the data against that part of the model. In other cases, we included the variables in anticipation of queries from readers.

We start our analysis with the credit growth equation, which is negatively affected by interest rates (in particular, by nominal rather than real rates, although these rates are measured relative to their equilibrium, 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 highlights the importance of improving the 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.[13] Moreover, forward-looking expectations seem to dominate the backward component (0.61 versus 0.39). This might be regarded as unusual for many advanced economies, but it is consistent with the conclusion of Caputo, Liendo, and Medina (2007) for Chile. However, setting the coefficient to 0.50 would not be inconsistent with the data at standard levels of statistical significance.

Since the skeletal model has an identity for GDP connecting absorption, imports, and exports, we need import and export price and income elasticities to complete it. Given that those are unavailable, we fitted a regression to capture these missing functions. The import and export shares are much the same in Brazil, so 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 that affects inflation is easily accepted, but we have chosen to leave the exchange rate in levels rather than differences.

The interest rate rule has neither an absorption gap nor an output gap, which is in line with Furlani, Portugal, and Laurini (2008). Expected inflation produced a better fit than actual inflation. The exchange rate has a small impact on interest rate decisions, and we therefore left it in the relation. The most striking difference with the skeletal model is the presence of a second lag in interest rates. As noted earlier, this equation can be rewritten 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, that is, $0.76\hat{\imath}_{t-1} + 0.32\Delta\hat{\imath}_{t-1}$.

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 earlier. To clarify, we generalized the skeletal model by replacing $E_t\tilde{z}_{t+1}$ with $[\phi\tilde{z}_{t-1}+(1-\phi)E_t\tilde{z}_{t+1}]$. A number of empirical implementations of theoretical models perform this step to reflect the well-known failure of UIP in exchange rates in the short-run (see Berg, Karam, and Laxton, 2006). It produces the

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13. The estimated equation involved $\Delta\tilde{n}_t$ as the dependent variable and $E_t(\tilde{n}_{t+1}-\tilde{n}_{t-1})$ as a regressor. We used $E_{t-1}\tilde{n}_t$ and $\tilde{n}_{t-1}$ as separate instruments for the latter variable. This was also true of the inflation equation.
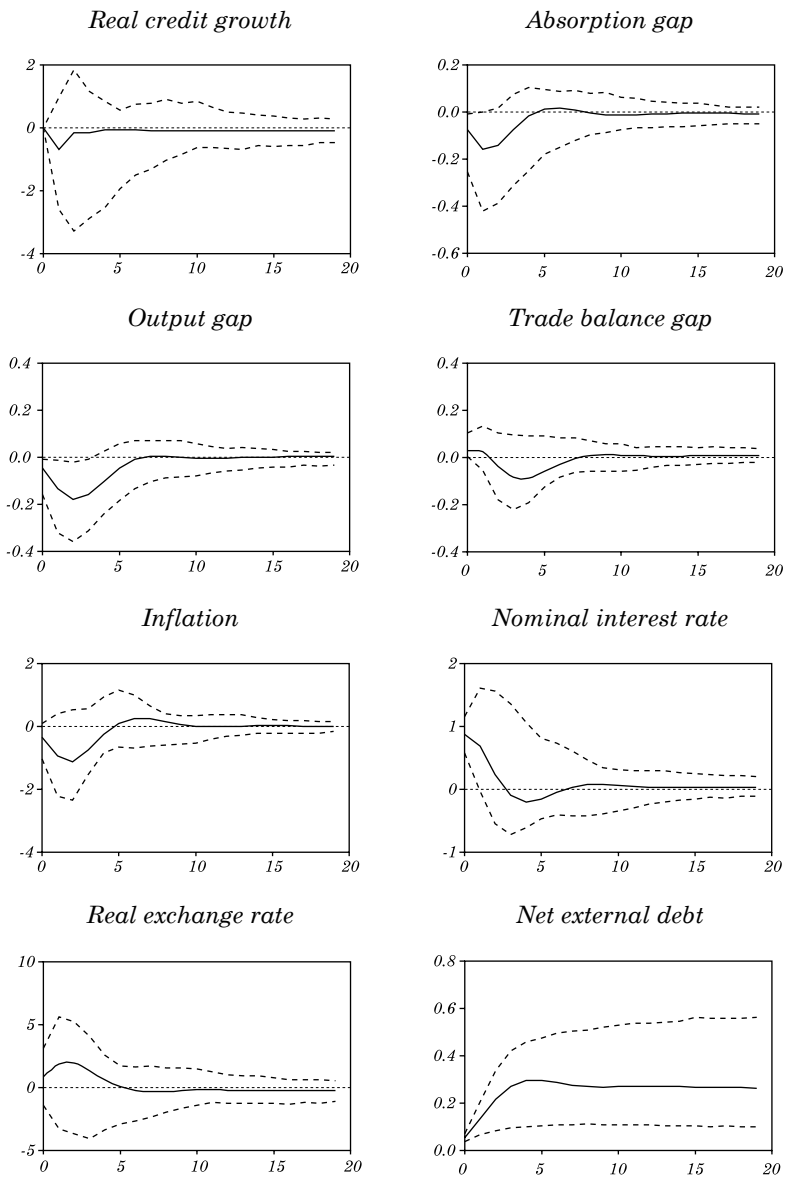
first term on the right-hand side of our estimated equation, except for a component $\phi(E_t\tilde{z}_{t+1} - E_{t-1}\tilde{z}_t)$. However, this latter term had a zero coefficient when we added it to the regression. The lagged nominal interest rate might seem to be an odd regressor. To show that it is not proxying for a real interest rate, we added in the expected inflation rate, which clearly is not accepted by the data. Once again, however, we note that $\hat{i}_t$ is measured as a deviation from an equilibrium rate. Since this includes the expected inflation rate, $\hat{i}_t$ cannot be thought of as a purely nominal rate.

Figure 1 gives the impulse responses for the 100 basis point rise in interest rates; figure 2 contains those for a 1 percent increase in credit growth. The figures also present the confidence intervals, which were chosen by simulating the model with the point estimates of the parameters and then choosing the range from 2.5 percent to 97.5 percent for the simulated parameter estimates. In these simulations, the estimated parameters might imply an unstable VAR, as the debt equation is always close to a unit root. When it was unstable, we discarded the simulated values.[14]

First, the interest rate rise. The strong exchange rate appreciation, rather than the output gap effect, is probably what produces the inflation response. 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 United States and the euro area (see Angeloni and others, 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 rate differential appreciates the real exchange rate (UIP-type effects) and boosts external debt (for example, 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 a higher relative price of nontradeables and a positive balance sheet effect), which somewhat offsets the negative effect of monetary tightening on absorption through the intertemporal channel. Thus, the positive effects of the appreciation on credit growth
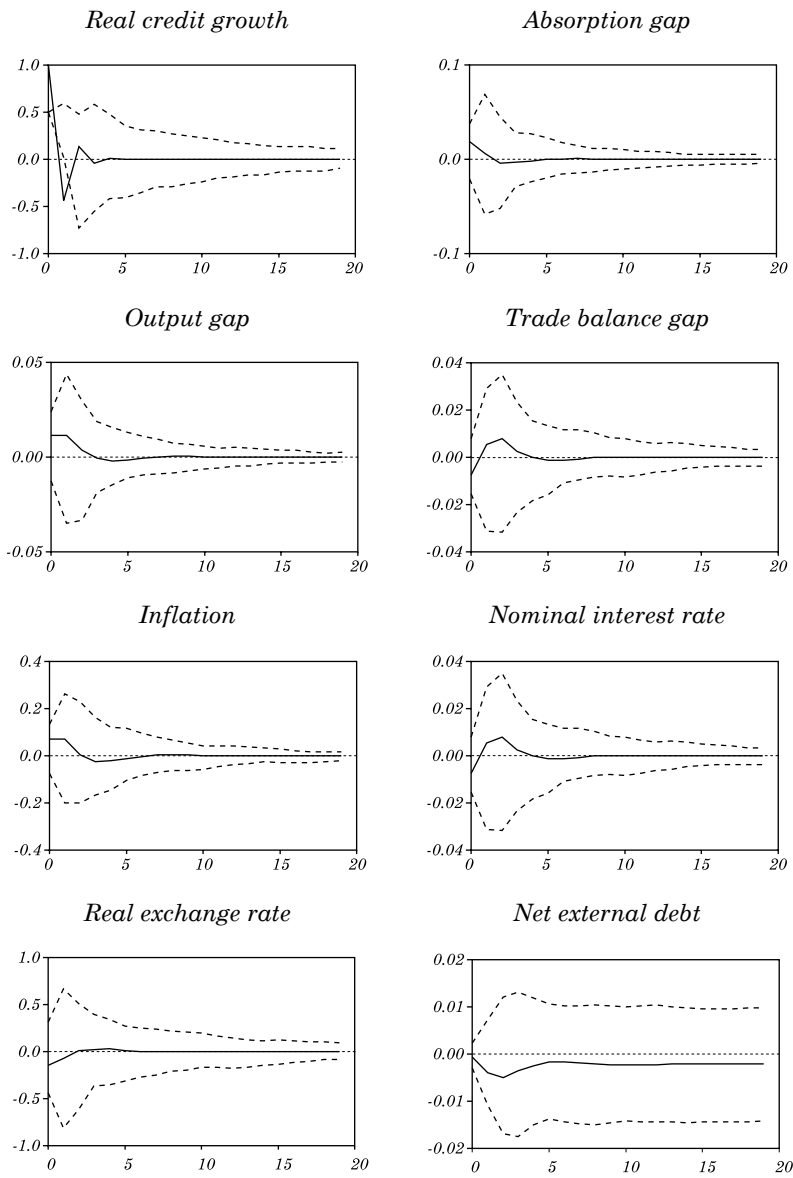
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14. The oscillations in the credit shock confidence intervals come from the fact that the autoregressive parameter in the credit growth equation is negative. Because the point estimate is small, the effect dies out quickly in the estimated impulse responses. Some simulations, however, yield a large negative value, in which case the oscillations persist for quite some time.

Figure 1. Brazil: Impulse Responses to a One Percent Rise in the Annualized Interest Rate

Real credit growth

Absorption gap

Output gap

Trade balance gap

Inflation

Nominal interest rate

Real exchange rate

Net external debt

Figure 2. Brazil: Impulse Responses to a One Percent Rise in Credit Growth

Real credit growth

Absorption gap

Output gap

Trade balance gap

Inflation

Nominal interest rate

Real exchange rate

Net external debt

Source: Authors' calculations.

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

In the experiment involving credit growth, we see 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 percent. It rapidly disappears, however, as interest rates rise and an exchange rate depreciation chokes off credit growth and absorption. The effects are even more short-lived than those associated with interest rate shocks, particularly regarding credit growth, in that they virtually vanish after four quarters. The same rise and fall is true for inflation, although it lasts a quarter or two longer. Given that there have been very large movements in real credit growth—with one standard deviation being the equivalent of 9 percent (annualized) growth—the impulse responses understate the impact of credit over the period, since the norm is not 1 percent growth but variations that are about nine times as high. This indicates that the macroeconomic effects of a standard deviation in credit growth appear to be of a higher magnitude and shorter duration than in developed countries.[15] The 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.[16] This indicates that the credit channel of monetary transmission is important in itself, quite separately from the open economy features of the skeletal model.

4.2 Chile

The structural VAR for Chile is fitted over 1999:1–2008:4, and its equations are given below. As for Brazil, the absolute *t* ratios are in parentheses. The import, export, and debt ratios were set to their

15. See the discussion by Eichenbaum (1994) on the difficulties of identifying credit channel effects in empirical work on the United States, for which longer and better data series and more disaggregated empirical evidence are available.

16. 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 space, but they are readily available on request.

averages over the period, and the nominal growth in potential GDP was set to 1.95 percent per quarter, based on a potential annual growth rate of 4.0 percent and target inflation of 3 percent per year. The results are not sensitive to this choice.

$$pc_t = \underset{(2.93)}{0.5}\, pc_{t-1} - \underset{(-0.83)}{0.57}\, \hat{r}_{t-1} + \underset{(0.66)}{0.15}\, \tilde{z}_{t-1} - \underset{(-0.77)}{0.17}\, \tilde{z}_{t-2} + \varepsilon_t^c;$$

$$\tilde{n}_t - \tilde{n}_{t-1} = \underset{(5.53)}{0.65}\big[E_t(\tilde{n}_{t+1}) - \tilde{n}_{t-1}\big] - \underset{(-1.40)}{0.15}\, \hat{r}_{t-2} + \underset{(1.33)}{0.04}\, pc_{t-1} - \underset{(-2.25)}{0.06}\, pc_{t-2} + \varepsilon_t^n;$$

$$\tilde{y}_t = \underset{(2.43)}{0.30}\, \tilde{n}_{t-1} - \underset{(-1.58)}{0.05}\, \tilde{z}_{t-1} + \underset{(1.65)}{0.05}\, \tilde{z}_{t-2} + \underset{(3.02)}{0.43}\, \tilde{y}_{t-1} + \varepsilon_t^y;$$

$$\hat{\pi}_t - \hat{\pi}_{t-1} = \underset{(10.35)}{0.49}\big[E_t(\hat{\pi}_{t+1}) - \hat{\pi}_{t-1}\big] + \underset{(2.86)}{0.21}\, \tilde{y}_t - \underset{(-3.17)}{0.05}\, \tilde{z}_t$$
$$+ \underset{(3.06)}{0.04}\, \tilde{z}_{t-1} + \underset{(3.57)}{0.03}\, pc_t + \varepsilon_t^\pi;$$

$$\hat{i}_t = \underset{(8.73)}{1.23}\hat{i}_{t-1} - \underset{(-3.46)}{0.45}\, \hat{i}_{t-2} + \underset{(0.98)}{0.16}\, \tilde{y}_t + \underset{(2.87)}{0.14}\, E_t(\hat{\pi}_{t+1}) + \varepsilon_t^i;$$

$$\zeta_t = \underset{(7.38)}{0.60}\, \tilde{z}_{t-1} - \underset{(-3.17)}{0.22}\, \tilde{z}_{t-2} + \underset{(4.38)}{0.29}\, \hat{d}_t + \varepsilon_t^z;$$

$$\zeta_t = \tilde{z}_t - E_t \tilde{z}_{t+1} - \left[\frac{(\hat{r}_t - \hat{r}_t^*)}{4}\right].$$

The equations generating expectations are as follows:

$$E_t \hat{\pi}_{t+1} = \underset{(10.19)}{1.41}\, \hat{\pi}_t - \underset{(-4.34)}{0.63}\, \hat{\pi}_{t-1} + \underset{(1.16)}{0.117}\, \tilde{n}_t + \underset{(1.59)}{0.025}\, pc_{t-1} + \underset{(2.65)}{0.325}\, \tilde{y}_t^*;$$
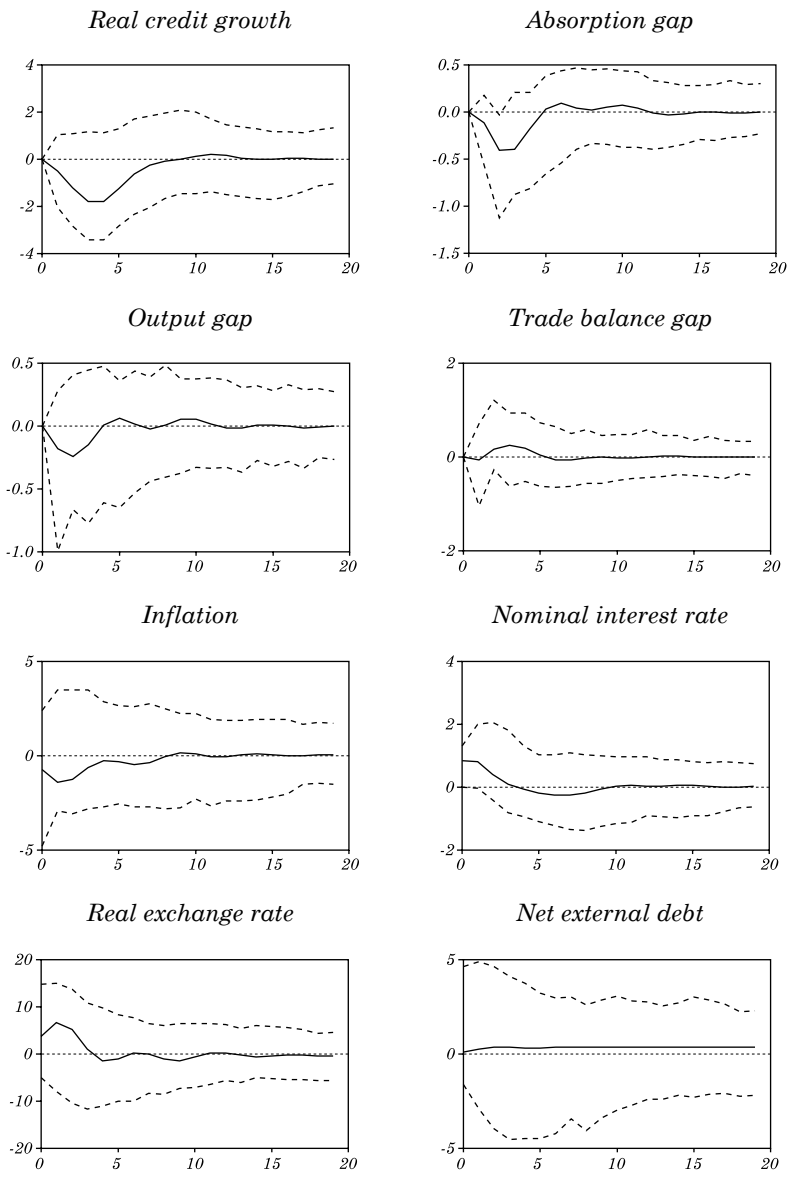
$$E_t \tilde{n}_{t+1} = -\underset{(-1.16)}{0.13}\, \hat{i}_{t-1} + \underset{(1.72)}{0.4}\, \tilde{y}_t - \underset{(-1.39)}{0.26}\, \tilde{y}_{t-1} - \underset{(-3.81)}{0.513}\, \tilde{n}_{t-1} + \underset{(1.89)}{0.314}\, \tilde{y}_t^*;$$

$$E_t \tilde{z}_{t+1} = \underset{(2.38)}{1.15}\, \hat{i}_t + \underset{(3.49)}{0.64}\, \tilde{z}_t - \underset{(-2.21)}{0.32}\, \tilde{z}_{t-1} - \underset{(-2.55)}{0.28}\, pc_{t-1} - \underset{(-3.43)}{0.63}\, \hat{d}_t + \underset{(1.13)}{0.96}\, \tilde{y}_t^*.$$

There are some notable similarities and differences with the Brazilian case. First, the exchange rate effect on credit growth is much weaker than in Brazil and more imprecisely estimated. The weaker effect may reflect the greater hedging of private sector balance sheets in Chile, which makes them less sensitive to currency valuation effects. Second, as in Brazil, credit affects absorption, probably through the growth of credit rather 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, the exchange rate effects are more than twice as strong for Chile as for Brazil, and they 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 not as strong, consistent with Chile being a much more open economy. Credit has a direct effect on inflation, and the coefficient on the output gap is very statistically significant. As in Brazil, the exchange rate clearly plays a nontrivial role in inflation, despite the general wisdom that exchange rate pass-through is lower in Chile. Indeed, the point estimate of −0.05 on the real exchange rate change ($\Delta\tilde{z}_t$) in the inflation equation suggests that a 10 percent nominal appreciation lowers CPI inflation by 50 basis points, all else constant. This is a nonnegligible effect. When we combine this estimate with the evidence that the exchange rate is highly responsive to shocks to the domestic interest rate (as illustrated in the impulse responses below), 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.

The interest rate equation resembles the results for Brazil in that it has two lags of the interest rate and a significant positive response to expected inflation. The output gap has a stronger effect on interest rate setting than was the case for Brazil, though neither is precisely estimated. The coefficient on expected inflation is less than $1-\beta_{ii}-\gamma_{ii}$, so the Taylor principle fails. This does not lead to explosive inflation, however, because there is a separate exchange-rate-induced effect on inflation in an open economy and because 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.

Figure 3. Chile: Impulse Responses to a One Percent Rise in the Annualized Interest Rate


Real credit growth

Absorption gap

Output gap

Trade balance gap

Inflation

Nominal interest rate

Real exchange rate

Net external debt

Source: Authors' calculations.

If a recursive SVAR(2) was fitted to a standard ordering of variables $\{\tilde{y}_t, \hat{\pi}_t, \hat{\imath}_t, \tilde{z}_t\}$, one would find that a rise in interest rates causes a rise in inflation and an initial depreciation in the exchange rate. This finding holds when the system is expanded to the full set of variables $\{pc_t, \tilde{n}_t, \tilde{y}_t, \hat{\pi}_t, \hat{\imath}_t, \tilde{z}_t, \tilde{d}_t\}$. The structural VAR impulses given in figure 3, however, are 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 in Chile relative to other emerging market countries, despite the relatively low pass-through. 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 in Brazil, with the real exchange rate appreciating by over 6.7 percent at its peak in response to a 100 basis point rise in the domestic policy rate, all else constant. This compares with a 2.0 percent response in Brazil, whereas the effect through the external debt term is similar in the two countries (0.28 in Chile versus 0.29 in Brazil). So, while our single equation estimates shown above indicate that the exchange rate pass-through to CPI is nearly three times lower in Chile than in Brazil, the effect of monetary policy on inflation through the exchange rate channel is quite strong in Chile because the exchange rate is so responsive to onshore-offshore interest rate differentials.

Figure 4 shows the results of a shock of a 1 percent increase in credit growth in Chile. Again, the results are similar to Brazil, although the 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.[17] This seems consistent with the evidence of a much greater banking sector penetration in Chile than in Brazil, as gauged by standard indicators of financial depth such as the ratio of bank credit to GDP (72 percent in Chile versus 40 percent in Brazil in 2008).

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

Figure 4. Chile: Impulse Responses to a One Percent Rise in Credit Growth
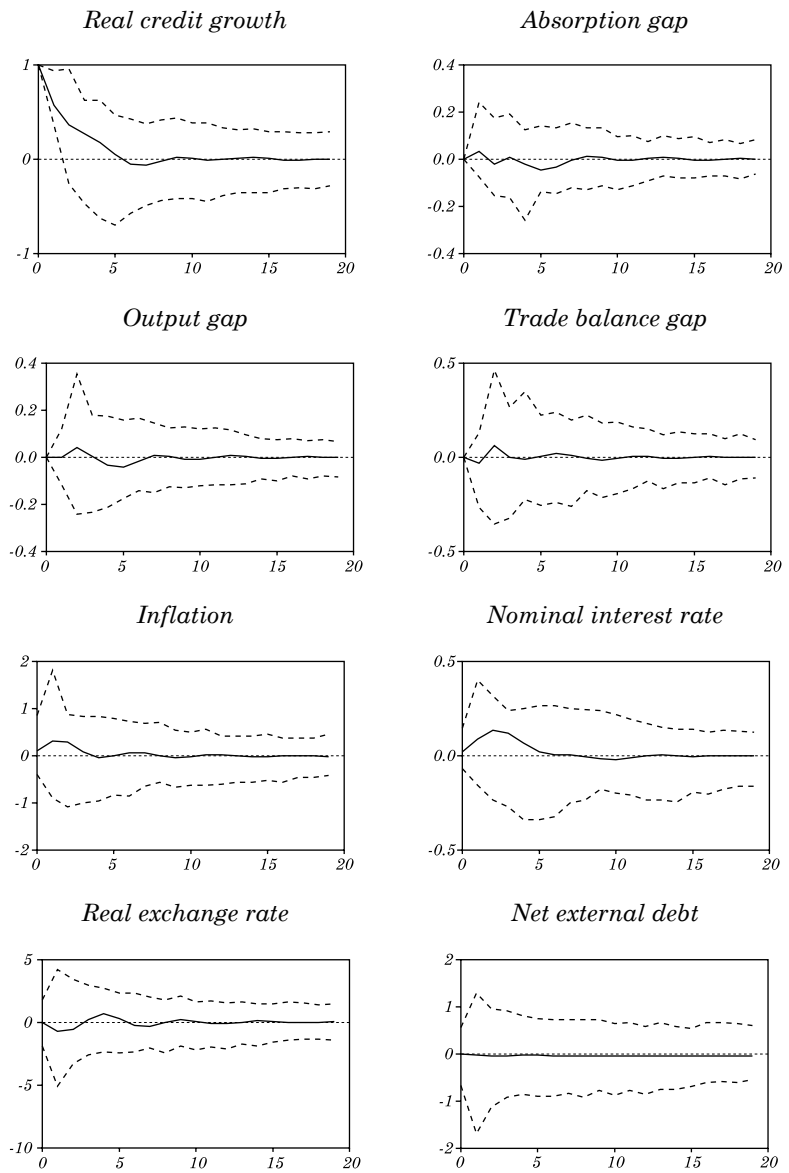
5. Conclusion

This paper has laid out a structural model of monetary transmission that incorporates key features of emerging markets in a manner parsimonious enough to be estimated with existing data and yet grounded on a DSGE 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 that have not featured in previous studies. An SVAR representation of the model was derived and used to examine the Brazilian and Chilean experiences with full-fledged inflation-targeting regimes since 1999. The two countries display important differences in economic structure and in the track record of economic policymaking. Most notably, Brazil is far more closed to trade, is less reliant on primary commodities, and has a more recent record of monetary and inflation stability than Chile. These differences make a comparative assessment of monetary transmission in the two countries particularly interesting. This diversity also provides a strict test of the general validity of the skeletal model and our estimation approach.

Our SVAR estimates yield very reasonable results for both countries. Indeed, we find no price puzzles, exchange rate puzzles, or any counterintuitive results in the impulse responses, which are common in VAR studies. This suggests that the proposal of a DSGE skeletal model as the basis of a structural VAR representation might provide a useful approach for examining monetary transmission in other emerging markets that are operating inflation-targeting regimes.

A common finding is that the transmission mechanism operates with shorter lags than in advanced countries (notably the United States): the bulk of the effects on output and inflation take place within five quarters. This is arguably consistent with structural factors (such as the shorter maturity of domestic credit) and the still-considerable (albeit reduced) weight of the exchange rate and imported inflation in domestic currency pricing, as is often mentioned in the literature.

The exchange rate effects on disinflation are nontrivial. This is all the more interesting in Brazil, which is still relatively closed to foreign trade, with ratios of exports and imports to GDP below 15 percent. Both countries display a sizable effect on the exchange rate of changes in the domestic interest rate policy, 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. 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 the role of bank credit in monetary transmission, our estimates indicate a nontrivial role for bank credit in monetary propagation. In both countries, there is evidence that changes in the policy rate affect credit growth and that the latter affects absorption. Moreover, at least in the case of Brazil, such a credit channel plays an intratemporal role in moderating the impact of monetary policy shocks on absorption via exchange rate effects: while higher interest rates reduce absorption through the standard intertemporal effect, they also boost bank credit demand via the short-run exchange rate appreciation that monetary tightening typically entails. The attendant balance sheet and wealth effects arising from such currency appreciations (particularly for nontradables 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 to be somewhat mitigated by the intratemporal exchange rate effect. We also find an independent role for credit shocks, which may reflect changes in reserve requirements and other regulations, as well as shocks to intermediation efficiency. Our estimates suggest that there are nontrivial effects on output and inflation in both countries, although these are reasonably short-lived, particularly in the case of Brazil.

An obvious practical implication is that policies that affect bank credit have direct effect 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. The far-reaching countercyclical credit policies implemented in both countries 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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