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# An Estimated Small Open Economy Model of the Financial Accelerator

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*This paper develops a small open economy model in which entrepreneurs partially finance investment using foreign currency–denominated debt subject to an external finance premium. We use Bayesian estimation techniques to evaluate the importance of balance sheet–related credit market frictions for emerging market countries by incorporating the financial accelerator mechanism. We obtain a sizable value for the external finance premium, which is tightly estimated away from zero. Our results support the inclusion of the financial accelerator in an otherwise standard model that—acting through balance sheets—magnifies the impact of shocks, thereby increasing real and financial volatility.* [JEL C11, F41]

Episodes of severe financial crises in recent years have renewed interest in the debate on the most appropriate exchange rate regime for emerging market countries (EMCs). For a small open economy, the common policy prescription—dating back to Friedman (1953)—has advocated exchange rate flexibility. As emphasized by the textbook Mundell-Fleming model, a freely floating nominal exchange rate can act as a shock absorber, insulating the economy against potentially destabilizing external disturbances.

However, EMCs face two fundamental issues that complicate the conduct of monetary policy. First, these countries can typically borrow only in foreign

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currency denominations, a phenomenon labeled "original sin" by Eichengreen and Hausmann (1999). This feature increases vulnerability to external shocks because a potential depreciation can substantially inflate debt service costs—owing to currency mismatches—and thus increase rollover risk. Second, EMCs usually have imperfect access to capital markets. Foreign credit is typically associated with an external finance premium that is linked to the condition of borrower balance sheets. Through the impact on the balance sheets of an EMC, these credit market frictions may substantially magnify the effects of shocks to the economy.

There has been a recent surge in theoretical models attempting to capture these salient features of EMCs. The work of Céspedes, Chang, and Velasco (2004); Devereux, Lane, and Xu (2004); Elekdag and Tchakarov (2004); as well as Gertler, Gilchrist, and Natalucci (2003), are some of the contributions to this literature. Building upon the framework developed by Bernanke, Gertler, and Gilchrist (1999), these authors explore the role of balance sheet–related credit market frictions for EMCs in an open economy context.

The key element in all of these papers is that lenders must incur a cost to monitor the business activity of borrowers caused by information asymmetries. Thus, lenders must be compensated in the form of an external finance premium that the borrower has to shoulder on top of the international interest rate. This premium in turn depends on the debt-to-net worth (leverage) ratio, which underpins what is known as the financial accelerator. In this context, an external shock that triggers an exchange rate depreciation could generate a vicious cycle. Owing to balance sheets with significant currency mismatches, a depreciation would inflate the value of foreign currency debt, thus eroding the value of domestic currency–denominated net worth. The deterioration in net worth would increase the premium, raising the cost of financing capital outlays, thereby amplifying the swings in borrowing and thus in investment, spending, and production.[1]

Against this background, the main question we ask in this paper is whether there is evidence in the data that favors the incorporation of the financial accelerator in an open economy setting. If the data support the inclusion of balance sheet–related credit frictions, then models that incorporate an endogenous external finance premium along with foreign currency–denominated debt could yield important insights into the debate regarding the most appropriate exchange rate regime for EMCs. Despite the potential relevance of the financial accelerator, the literature cited above has only used calibrated models to highlight the implica-

[1]Krugman (1999) and Aghion, Bacchetta, and Banerjee (2001), among others, argue that exchange rate and interest rate fluctuations—through balance sheet constraints impacting investment spending—affect borrowers in EMCs disproportionately more than entrepreneurs in industrialized economies. Calvo and Reinhart (2000) argue that the reluctance to implement a pure float ("fear of floating") could be justified by the fact that large exchange rate movements may devastate corporate and financial balance sheets, because of large outstanding foreign currency–denominated debt obligations. Therefore, one way EMCs apparently deal with such vulnerabilities is by attempting to minimize exchange rate fluctuations. In this context, Elekdag and Tchakarov (2004) show that when the foreign currency–denominated debt-to-GDP ratio exceeds a certain threshold, the welfare costs associated with a pure float could exceed those of managed exchange rate regimes.

tions of credit market frictions. The main contribution of this paper is therefore the estimation of a model incorporating the financial accelerator, which allows us to infer the average premium and to quantify balance sheet vulnerabilities.[2]

To this end, we estimate a stylized model that includes the financial accelerator channel relying on recent developments in Bayesian estimation techniques following Schorfheide (2000), Lubik and Schorfheide (2003), and Smets and Wouters (2003), as well as Justiniano and Preston (2004). One of the main advantages of a Bayesian approach to estimating dynamic stochastic general equilibrium (DSGE) models is that it allows a complete characterization of uncertainty in estimating structural parameters by simulating posterior distributions. It also provides an elegant and coherent way to incorporate prior information about parameters, coming either from microeconomic studies or from previous macroeconomic analyses.

Our main findings can be summarized as follows. First, based on Korean data, the median estimate of the external finance premium is 2.75 percent per quarter, with 90 percent probability bands ranging from about 1 percent to 6 percent. The existence of a sizable premium indicates that there is evidence in favor of the financial accelerator mechanism, and highlights the importance of balance sheet–related credit market frictions. Although the annualized premium corresponds to 11 percent—most likely reflecting the impact of the East Asian crisis when Korean spreads exceeded 900 basis points—the lower end of our 90 percent probability bands covers the actual realization of Korean spreads over our sample period.

Second, the elasticity of the external finance premium to the capital-to–net worth ratio is tightly estimated away from zero with a median value of 0.048. This elasticity could be interpreted as quantifying the importance of balance sheet vulnerabilities. Our estimate is consistent with calibration-based models incorporating the financial accelerator and confirms, once again, that the data support the inclusion of a mechanism capturing balance sheet–related credit market frictions.

Third, the median estimates for the main parameters of the model are broadly consistent with calibrated values found in the previous literature, despite our choice of fairly uninformative priors. Furthermore, it is important to emphasize that our results are robust across different priors and model specifications. To summarize, our results support the inclusion of the financial accelerator in an otherwise standard model, which—acting through balance sheets—magnifies the impact of shocks, thereby increasing real and financial volatility.

## I. The Model

Our modeling framework is an extension of Céspedes, Chang, and Velasco (2004) as well as Elekdag and Tchakarov (2004), in which we focus on a small open economy with a representative household, producers, entrepreneurs, and a central bank. The consumption good is a composite of a domestically produced good and an imported foreign good. The domestic good is a bundle composed of a continuum

[2]Related research includes Christiano, Motto, and Rostagno (2003); Meier and Müller (2006); and Linaa and Rand (2004), as well as Tovar (2005).

of goods produced by domestic firms in a monopolistically competitive environment. Because these firms supply a unique differentiated good, they enjoy market power, which they exploit to maximize their profits. This setup motivates price stickiness, which warrants active cyclical monetary policy. To this end, we incorporate a central bank that uses an interest rate rule to achieve specific policy objectives. The representative household is allowed to accumulate financial assets and is thus responsive to interest rate fluctuations.

## The Representative Household

The household is infinitely lived and its preferences are defined over processes of aggregate consumption, $C_t$, and labor effort, $L_t$, which are described by the following utility function:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[ \frac{C_t^{1-\sigma}}{1-\sigma} - \chi_t \frac{L_t^{1+\psi}}{1+\psi} \right], \tag{1}$$

where $E_t$ denotes the mathematical expectation conditional on information available in period $t$, $\beta \in [0,1]$ is the subjective discount factor, and $\chi_t$ is a labor disutility shock.[3]

Aggregate consumption is a bundle consisting of a domestically produced good and an imported foreign good:

$$C_t = \aleph C_{Ht}^{\gamma} C_{Ft}^{1-\gamma}, \tag{2}$$

where $C_{Ht}$ denotes the consumption of the home good, $C_{Ft}$ denotes consumption of the imported good, and $\aleph = [\gamma^{\gamma} (1-\gamma)^{1-\gamma}]^{-1}$ is a normalizing constant.

For simplicity, we assume that the price of the imported good is normalized to unity in terms of foreign currency. Also, imports are assumed to be freely traded and the law of one price holds, thus the domestic currency price of imports is just equal to the nominal exchange rate, $s_t$. The aggregate price level, $\rho_t$, is then derived by solving for the minimum expenditure required to obtain one unit of the aggregate consumption good.[4] Denoting the price of the domestic good as $p_t$, the aggregate price level is then

$$\rho_t = p_t^{\gamma} s_t^{1-\gamma}. \tag{3}$$

The domestic good, however, is itself a composite good, consisting of an Armington aggregate of a continuum of differentiated domestic goods; more specifically,

$$C_{Ht} = \left[ \int C_{Hjt}^{\frac{\lambda_t - 1}{\lambda_t}} dj \right]^{\frac{\lambda_t}{\lambda_t - 1}}, \tag{4}$$

[3]In fact, we consider a continuum of households indexed with $h \in [0,1]$. However, in a symmetrical equilibrium, these households will behave identically; therefore, we suppress this index.

[4]Given the aggregate price index defined in equation (3), the individual consumption demands for each good are $C_{Ft} = (1-\gamma)\rho_t C_t / s_t$ and $C_{Ht} = \gamma \rho_t C_t / p_t$.

where $j \in [0,1]$ and $\lambda_t > 1$.[5]

The household's budget constraint in period $t$ is as follows:

$$\rho_t C_t + \rho_t \Gamma_t + b_t + s_t F_t^* = w_t L_t + \Pi_t + (1+i_{t-1})\, b_{t-1} + \left(1+i_{t-1}^*\right) s_t F_{t-1}^*, \tag{5}$$

where $b_t$ and $F_t^*$ are nominal stocks of domestic and foreign currency–denominated bonds maturing in period $t$, which earn interest $i_{t-1}$ and $i_{t-1}^*$, respectively. Households earn wage $w_t$ for their labor services, $L_t$. Because they own the domestic production firms, they retain any profits, $\Pi_t$. Finally, households must incur an intermediation cost, $\Gamma_t$, which has the following specification:

$$\Gamma_t = \frac{\theta_B}{2}\left(\frac{b_t}{\rho_t}\right)^2 + \frac{\theta_{F^*}}{2}\left(\frac{s_t F_t^*}{\rho_t}\right)^2, \tag{6}$$

where $\theta_B$, $\theta_{F^*} \geq 0$ and $\theta_B + \theta_{F^*} > 0$. Without this cost, the stocks of bonds and consumption would not be stationary.

The household chooses the paths of $\{C_t, L_t, b_t, F_t^*\}_{t=0}^{\infty}$ to maximize expected lifetime utility (equation (1)) subject to the constraint (equation (5)) and initial values for $b_0$ and $F_0^*$. Ruling out Ponzi-type schemes, we get the following first-order conditions:

$$\chi_t \frac{L_t^{\psi}}{C_t^{-\sigma}} = \frac{w_t}{\rho_t}, \tag{7}$$

$$C_t^{-\sigma}\left[1+\theta_B\left(\frac{b_t}{\rho_t}\right)\right] = \beta(1+i_t)E_t\left[C_{t+1}^{-\sigma}\frac{\rho_t}{\rho_{t+1}}\right], \tag{8}$$

$$C_t^{-\sigma}\left[1+\theta_{F^*}\left(\frac{s_t F_t^*}{\rho_t}\right)\right] = \beta\left(1+i_t^*\right)E_t\left[C_{t+1}^{-\sigma}\frac{\rho_t}{\rho_{t+1}}\frac{s_{t+1}}{s_t}\right], \tag{9}$$

where the first condition implies that the household equates its marginal rate of substitution between consumption and leisure to the real wage, $w_t / \rho_t$. The last two first-order conditions are familiar Euler equations that state the household's preference to smooth consumption. Abstracted from the intermediation costs, they can be combined to yield the familiar uncovered interest parity condition.

## Production Firms

The economy is populated by a multitude of monopolistically competitive firms, each producing a unique good. The production technology for an arbitrary firm $j \in [0,1]$ is

$$Y_{jt} = A_t K_{jt}^{\alpha} L_{jt}^{1-\alpha} \tag{10}$$

[5]As in Smets and Wouters (2003), in the log-linearized version of the model, the mean value of $\lambda_t$ is perturbed by an $AR(1)$ disturbance, which can be interpreted as a cost-push shock to inflation.

where $A_t$ is the technology shock common to all production firms. The household provides the labor services whereas capital, $K_{jt}$, is provided by the entrepreneurs as described below. Production firms exploit their market power and set prices to maximize profits along a downward sloping demand curve given by

$$Y_{jt} = \left(\frac{p_{jt}}{p_t}\right)^{-\lambda_t} Y_t. \tag{11}$$

When $mc_t$ is defined as the common marginal cost for all firms in the economy, cost-minimizing behavior implies the following optimal conditions:

$$w_t L_t = (1-\alpha)\frac{\lambda_t}{\lambda_t - 1} mc_t Y_t, \tag{12}$$

$$r_t K_t = \alpha \frac{\lambda_t}{\lambda_t - 1} mc_t Y_t, \tag{13}$$

where $r_t$ is the rental rate on capital. Equations (12) and (13) are implicit demand curves for labor and capital, respectively.

With aggregate investment denoted as $I_t$, the stock of capital used by all of the firms in the economy evolves according to

$$K_{t+1} = \Phi\left(\frac{I_t}{K_t}\right) K_t + (1-\delta) K_t, \tag{14}$$

where $\delta$ is the depreciation rate and the adjustment cost function is increasing and concave, satisfying $\Phi'(\delta) > 1$ and $\Phi(\delta) = \delta$. Adjustment costs allow a variable price of capital and therefore contribute to the volatility of entrepreneurial net worth. In equilibrium, the price of capital, $q_t$, is given by

$$q_t = \left[\Phi'\left(\frac{I_t}{K_t}\right)\right]^{-1} \rho_t. \tag{15}$$

## Price Setting

According to the staggered contract set up in Calvo (1983) and Yun (1996), firms are assumed to reset new prices with probability $(1 - \kappa)$ in every period, independent of how long their price has been fixed. If prices are not reset, the old price is adjusted by the rate of steady state gross inflation, $\bar{\pi}$.

When firm $j \in [0,1]$ is allowed to reset its price in period $t$, the firm chooses $p_t(j)$ to maximize the following profit function:

$$E_0 \sum_{\tau=0}^{\infty} \kappa^{\tau} \Lambda_{t,t+\tau} [\pi^{\tau} p_t(j) - mc_{t+\tau}] Y_{t+\tau}(j). \tag{16}$$

In the case where $\kappa \to \infty$, the optimal choice for $p_t(j)$ becomes a simple markup rule:

$$p_t(j) = \frac{\lambda_t}{\lambda_t - 1} mc_t, \tag{17}$$

which implies that there is a markup of price over marginal cost.[6]

## Entrepreneurs

One of our main objectives in this paper is to uncover evidence of balance sheet vulnerabilities. To this end, we adapt the modeling framework provided by Céspedes, Chang, and Velasco (2004) when introducing the financial accelerator in an open economy context. Although the inclusion of entrepreneurs is crucial to our investigation, we attempt only a concise presentation and refer the reader to the work of Céspedes, Chang, and Velasco (2004); Elekdag and Tchakarov (2004); and Bernanke, Gertler, and Gilchrist (1999) for further details.

Entrepreneurs finance investment partly with foreign loans, which are subject to frictions. At any given period, the entrepreneur is assumed to have some net worth denominated in domestic currency. When the entrepreneur engages in capital accumulation, investment outlays are financed partly with net worth and partly with foreign currency–denominated debt. Hence, the entrepreneur is subject to the following budget constraint:

$$p_t N_t = q_t K_t - s_t D_t^*, \tag{18}$$

where $N_t$ and $D_t^*$ denote net worth and foreign currency–denominated debt, respectively. It is important to note that equation (18) is a standard accounting identity that states that net worth is equal to assets minus liabilities. Therefore, an unanticipated depreciation—an increase in $s_t$—will inflate liabilities and reduce net worth, highlighting the vulnerability of balance sheets to exchange rate fluctuations.

Entrepreneurs are risk neutral and choose the stock of firm capital, $K_t$, as well as the associated level of borrowing, $D_t^*$, to maximize profits. Owing to informational asymmetries, the expected return on capital is equal to the foreign interest rate adjusted for expected exchange rate fluctuations, augmented by an external finance premium, that is,

$$E_t\left[\frac{r_{t+1}^K}{\rho_t}\right] = \left(1 + i_t^*\right) E_t\left[\frac{s_{t+1}}{s_t}\right] \Psi\left(\frac{q_t K_t}{p_t N_t}\right), \tag{19}$$

[6]However, in the model, we use the general specification of the optimal choice for $p_t(j)$, which is

$$p_t(j) = \frac{\lambda_t}{\lambda_t - 1} \frac{E_t \sum_{\tau=0}^{\infty} \kappa^\tau \Lambda_{t,t+\tau} mc_{t+\tau} Y_{t+\tau}(j)}{E_t \sum_{\tau=0}^{\infty} \kappa^\tau \Lambda_{t,t+\tau} \bar{\pi}^\tau Y_{t+\tau}(j)},$$

where the discount factor is formally defined as $\Lambda_{t,t+\tau} = \beta C_t / C_{t+\tau}$.

where $\Psi(\cdot)$ denotes the endogenous external finance premium, which satisfies $\Psi(1)=1$ and $\Psi'(\cdot) > 0$. Denoting the capital-to-net worth ratio as $k = qK/pN$, we define the elasticity of the external finance premium with respect to $k$ as $\nu = [\Psi'(k)/\Psi(k)]k$.[7] The elasticity determines the percentage increase in the premium when the capital-to–net worth ratio increases by 1 percent.[8]

At the beginning of each period, entrepreneurs collect their returns on capital and honor their foreign debt obligations. Because it will be assumed that they consume a fraction $(1 - \eta_t)$ of the remainder on imports, net worth evolves according to the following formulation:

$$p_t N_t = \eta_t \left[ r_t^K q_{t-1} K_{t-1} (1-\zeta_t) - \left(1+i_{t-1}^*\right) \Psi\left( \frac{q_{t-1}K_{t-1}}{p_{t-1}N_{t-1}} \right) s_t D_{t-1}^* \right], \tag{20}$$

where, as in Céspedes, Chang, and Velasco (2004), $\zeta_t$ reflects the deadweight cost associated with imperfect capital markets.[9] As can be seen from equation (20), liabilities are susceptible to foreign interest rate and exchange rate fluctuations.[10] An unanticipated increase in the foreign interest rate or a sudden depreciation can inflate the debt service obligations of the entrepreneur, potentially deteriorating net worth, thus increasing the premium. This further increases the opportunity cost of investment, which hampers capital accumulation, thus exacerbating the severity of the potential recession.

In a model with investment adjustment costs and capital depreciation, we need to differentiate between the entrepreneur's return on capital, $r_t^K$, and the rental rate of capital, $r_t$. The former depends on the latter as well as on the value of the capital stock net of depreciation, adjusted for asset price valuation effects (fluctuations in $q_t$), that is,

$$E_t r_t^K = E_t \left[ \frac{r_{t+1} + (1-\delta) q_{t+1}}{q_t} \right]. \tag{21}$$

## The Central Bank

In our model, we include a central bank that implements a general interest rate rule to achieve specific policy objectives. The interest rate rule takes the following form:

$$i_t = i + \zeta_i i_{t-1} + \zeta_\pi \hat{\pi}_t + \zeta_s \left(\hat{s}_t - \hat{s}_{t-1}\right) + \zeta_Y \left(Y_t - Y\right) + \varepsilon_{it}, \tag{22}$$

[7]Bernanke, Gertler, Gilchrist (1999) provide further details on the increasing relationship between the entrepreneur's capital-to-net worth ratio and the external finance premium.

[8]Equivalently, we could have used the leverage ratio—also referred to as the (foreign) debt-to-equity ratio defined as $sD/pN$, based on $qK/PN = 1 + sD/pN$, as implied by equation (18).

[9]More specifically, it is the cost associated with monitoring, and it is an increasing function of the risk premium; see Bernanke, Gertler, and Gilchrist (1999) for the full exposition.

[10]The term $(1 - \eta_t)$ may also be interpreted as the entrepreneur's bankruptcy rate. In the log-linearized version of the model, $\eta_t$ is perturbed from its mean by an $AR(1)$ disturbance, which—as in Christiano, Motto, and Rostagno (2003)—could be interpreted as a shock to the rate of destruction of entrepreneurial financial wealth mimicking the bursting of a stock market bubble.

where the monetary authority sets the nominal interest rate, taking into consideration the inflation rate, the output gap, the rate of exchange rate depreciation, and the previous period's interest rate—allowing for a wide array of monetary policy regimes. Note that as $\zeta_\pi \to \infty$, the central bank is implementing strict inflation targeting, and when $\zeta_s \to \infty$, the bank is implementing a fixed exchange rate regime. If $\zeta_\pi$ is finite and $\zeta_s > 0$, then a managed float is being implemented. If $\zeta_i > 0$, then the central bank is engaging in interest rate smoothing. Finally, with $\varepsilon_{it}$, we denote an independent and identically distributed domestic monetary policy shock.

### Market Clearing

Domestic expenditure on home goods is a fraction, $\gamma$, of final expenditures. The home good is also sold to foreigners, whose demand is assumed to be an exogenous stochastic process, $X_t$. The market clearing condition for home goods is thus

$$p_t Y_t = \gamma \rho_t [C_t + I_t] + s_t X_t. \tag{23}$$

We close the model by imposing a market clearing condition for domestic bonds and by specifying stochastic processes for the exogenous variables.[11] Assuming all firms behave symmetrically, the stationary rational expectations equilibrium is a set of stationary stochastic processes, $\{\rho_t, p_t, q_t, s_t, b_t, F^*_t, D^*_t, w_t, L_t, r_t, r^K_t, i_t\}_{t=0}^{\infty}$ and $\{i^*_t, Y_t, K_t, I_t, N_t, mc_t, \pi_t, A_t, X_t, cp_t, \chi_t, \eta_t\}_{t=0}^{\infty}$, satisfying equations (3), (5), (7)–(10), (12)–(15), (18)–(23), and the market clearing conditions, along with initial values $b_0$, $F^*_0$, and $D^*_0$.

## II. Estimation Methodology

We estimate the model using Bayesian methods based on the influential work of Schorfheide (2000). Papers using a Bayesian approach in the estimation of open economy DSGE models include Lubik and Schorfheide (2003), Justiniano and Preston (2004), and Adolfson and others (2005). There are several advantages of using Bayesian methods for inference in estimating macroeconomic models. For our purposes, we highlight the fact that because Bayesian methods seek to characterize the posterior distribution of the parameters, they facilitate an accurate assessment of all of the uncertainty surrounding the model's coefficients. Indeed, posterior inference provides us with posterior probability bands without having to assume, for instance, symmetry in these distributions.[12]

[11]More specifically, the domestic bond clearing condition implies that $b_t = 0$. The six exogenous variables—$i^*_t$, $A_t$, $X_t$, $\chi_t$, $\eta_t$, and $cp_t$—are all assumed to be $AR(1)$ processes. Finally, to close the model, we need to explicitly define the inflation rate: $\pi_t = \rho_t / \rho_{t-1}$.

[12]There are also clear advantages when it comes to model comparisons because the models are not required to be nested and numerical methods for the computation of the marginal likelihood permit constructing posterior model probabilities. These probabilities can in turn be used for model averaging, thereby producing parameter estimates that also explicitly incorporate model uncertainty. Furthermore, as emphasized by Smets and Wouters (2003), the use of Bayesian methods provides greater stability to optimization algorithms relative to maximum likelihood.

We briefly sketch our approach to inference, and the reader is referred to the above references for further details. Defining $\Theta$ as the parameter space, we wish to estimate the model parameters denoted by $\theta \in \Theta$. Given a prior $p(\theta)$, the posterior density of the model parameters, $\theta$, is given by

$$p(\theta|Y^T) = \frac{L(Y^T|\theta)p(\theta)}{\int L(Y^T|\theta)p(\theta)d\theta},$$

where $L(Y^T|\theta)$ is the likelihood conditional on observed data, $Y^T$. The likelihood function is computed under the assumption of normally distributed disturbances by combining the state-space representation implied by the solution of the linear rational expectations model and the Kalman filter.[13]

Our goal is to therefore characterize the posterior density of the parameters. To do so we follow a two-step approach. In the first step, a numerical algorithm is used to find an initial guess of the posterior mode by combining the likelihood $L(Y^T|\theta)$ with the prior. To deal with the possibility of multiple modes, following Justiniano and Preston (2004), several (at least 20) optimization algorithms are run with random starting values chosen from the prior to determine whether they all converge to a unique mode. The unique posterior mode obtained from this first step is used as the starting value ($\theta^0$) of a multiple chain Random Walk Metropolis algorithm. This Markov Chain Monte Carlo (MCMC) method allows us to generate draws from the posterior density $p(\theta|Y^T)$. At each step $i$ of the Markov Chain, the proposal density is used to draw a new candidate parameter $\theta^* \sim N(\theta^i, c\Sigma)$. The new draw is then accepted with the following probability:

$$\omega = \min\left\{1, \frac{L(Y^T|\theta^*)p(\theta^*)}{L(Y^T|\theta^i)p(\theta^i)}\right\}.$$

If accepted, $\theta_k^{i+1} = \theta_k^*$, otherwise, $\theta_k^{i+1} = \theta_k^i$. We generate multiple chains of 70,000 replications in this manner, discarding the first 10,000 iterations while monitoring the convergence of the generated draws using potential scale reduction factors and trace plots.

The scaling constant for the variance covariance matrix, $c$, is chosen to attain a 25 percent acceptance rate. With the generated draws, point estimates of $\theta$ can be obtained from the simulated values by using various location measures, such as means or medians. Similarly, measures of uncertainty follow from computing the percentiles of the draws.

## Data

To estimate the model, we use five key macroeconomic time series for Korea: real GDP, inflation, hours worked, the nominal interest rate, and the real CPI-based exchange rate.[14] The average annualized Korean money market rate and the annu-

[13]The model is solved using Sims' (2003b) method.

[14]All series are extracted from Datastream International. Korea was chosen primarily because of data availability and the fact that it is an EMC that is not a net hydrocarbon or primary commodity exporter.

alized log percentage change in the GDP deflator correspond to the interest and inflation rates, respectively. Meanwhile, real GDP, the real exchange rate, and hours worked are expressed in log-deviations from a linear time trend.

The sample period runs from the first quarter of 1990 to the third quarter of 2003. Even though longer time series for the five observable variables were available, the sample is restricted because Korea was pegging its nominal exchange rate to a basket of major currencies throughout the 1980s.[15] During March 1990, Korea switched to a managed float, with bands of 2.25 percent around the won-to-dollar exchange rate.[16] On November 19, 1997, the band was widened to 10 percent, and on December 16, 1997, Korea abolished its band and allowed the won to float freely. To avoid any nonlinear regime shifts, we therefore focus on the sample period during which Korea was pursuing a managed or a free float.[17]

The most striking feature of our data sample is the turbulence generated by the Korean crisis. As in other financial crises, the Korean data feature a dramatic real exchange rate depreciation, a spike in interest rates and spreads—the J.P. Morgan Emerging Markets Bond Index Global (EMBIG) stripped spread hit 940 basis points during August 1998—as well as a severe recession, with the associated sharp declines in investment and consumption.[18] In this context, one of the objectives of this paper is to assess how well a stylized model can account for the macroeconomic instability that occurred during our sample period.

## Prior Distribution of the Parameters

In this section we review the assumptions that underpin our prior distributions for the parameters. So that the structure of our model is immediately comparable with those used in Céspedes, Chang, and Velasco (2004) and Elekdag and Tchakarov (2004), we initially abstract from investment adjustment costs and impose full depreciation, but later analyze a specification that relaxes these assumptions. To avoid identification problems, we calibrate a few parameters: we set $\alpha$, $\beta$, $\theta_F{}^*$, and $\phi_\eta$ to 0.37, 0.99, 0.01, and zero, respectively, as is standard in the literature.[19] Table 1 through Table 4 report the type of density, as well as the mean and standard deviation, for each estimated parameter. We denote the capital-to-net worth ratio

[15]Nonetheless, we use observations from 1988 to 1989 for the initialization of the Kalman filter although these observations are not used in the computation of the likelihood and the estimation of the parameters.

[16]Under the market average exchange rate system introduced in March 1990, the won was allowed to float against the U.S. dollar within a daily trading range of the weighted average of the previous day's rates in the interbank market.

[17]It might be argued that we should have restricted ourselves to a period of either pure float or managed float. Although Markov switching methods might allow for incorporating the transition to an alternative exchange rate policy, our limited sample prevents us from considering the estimation under the two regimes. Furthermore, this would entail estimating a nonlinear model.

[18]The J. P. Morgan EMBIG series code is JPSSGKOR Index.

[19]As argued in Smets and Wouters (2003), these parameters present difficulties in the estimation unless the absolute values of the time series are taken into account through the definition of the steady state. Furthermore, the adjustment cost parameter, $\theta_F{}^*$, is calibrated because it primarily serves to overcome the unit root problem in open economy models. In the working paper version of this paper (Elekdag, Justiniano, and Tchakarov, 2005), we have considered an alternative calibration for $\beta$ that does not affect any of our conclusions.

with $k$ and the elasticity of the external finance premium to $k$ with $\nu$.[20] These two parameters are of most interest to us because they underpin the external finance premium. Kawai, Hahm, and Iarossi (2000) conduct a survey of 850 industrial Korean firms, showing that the total debt-to-capital ratio is around 70 percent. If we assume that all of this debt is denominated in foreign currency, we obtain an upper bound for $k$ of 3.3. Moreover, the authors report that the share of debt in foreign currency denominations was at least 20 percent, which then determines a lower bound for $k$ of 1.25. Therefore, we choose a gamma prior distribution for $k$ with a mean of 2 and standard deviation of 0.3, which incorporates these bounds. Lacking a reliable empirical benchmark in specifying our prior for $\nu$, we choose a beta distribution for this parameter with a mean of 0.07 and standard deviation of 0.03, which encompasses the calibrated values used in the literature cited above. Taken together, these prior densities allow for a wide range of values for the external finance premium. Nonetheless, in the robustness section below, we also consider an alternative prior for $\nu$ that is less informative.

The degree of openness is determined by the value of $(1 - \gamma)$ and, along with the Calvo (1983) price adjustment probability, $\kappa$, both parameters are by definition constrained to the unit interval. Therefore, we specify the prior densities for $\gamma$ and $\kappa$ as belonging to the beta family with means of 0.6 and standard deviations of 0.1 for both. For the remaining three structural parameters, the priors are specified as gamma distributions. For the inverse of the intertemporal elasticity of substitution, $\sigma$, and the elasticity of labor supply, $\psi$, we center the gamma densities at 3, with a standard deviation of 1. Meanwhile, the intratemporal elasticity of substitution between varieties of domestically produced goods, $\lambda$, has a gamma prior with a mean of 8 and a standard deviation of 3, implying an average markup of about 14 percent, although allowing for a range of values from 6 percent to 59 percent.

The parameters describing monetary policy are based on a Taylor rule that allows for interest rate smoothing and is augmented to include responses to the exchange rate in addition to output and inflation. We conjecture that the central bank smoothes interest rate adjustments; thus, we set the prior for $\zeta_i$ as a beta density centered around 0.8 with a standard deviation of 0.2. Priors for the coefficients on inflation, the output gap, and the exchange rate—$\zeta_\pi$, $\zeta_Y$, and $\zeta_S$—belong to the gamma density and have means of 3, 1.2, and 1, respectively, as well as standard deviations of 0.5, 0.8, and 0.8, respectively. As noted above when describing equation (22), these priors allow for a wide range of monetary policy regimes.

The priors for the variances of the exogenous stochastic processes correspond to an inverse Wishart distribution with a mean of unity and standard deviation of 0.75. The exogenous shocks are all $AR(1)$ processes; we therefore specify a fairly flat beta distribution with a mean of 0.5 and a standard deviation of 0.25 for each autoregressive coefficient. It is worth emphasizing that, as opposed to many other papers that use Bayesian techniques for estimating DSGE models, we choose prior distributions that are very flexible and permit a broad array of possible values.

[20]Recall that $\nu = [\Psi'(k)/\Psi(k)]k$ and $k = qK/pN$.

Table 1. Parameter Estimates for the Baseline Model

| | Prior Distribution | | | Posterior Distribution | | |
|---|---|---|---|---|---|---|
| Parameter | Type | Mean | Standard error | 10% | Median | 90% |
| $\gamma$ | Beta | 0.600 | 0.100 | 0.661 | 0.689 | 0.718 |
| $\kappa$ | Beta | 0.600 | 0.100 | 0.317 | 0.387 | 0.458 |
| $\lambda$ | Gamma | 8.000 | 3.000 | 5.521 | 8.545 | 12.644 |
| $\psi$ | Gamma | 3.000 | 1.000 | 1.516 | 1.889 | 2.373 |
| $\sigma$ | Gamma | 3.000 | 1.000 | 0.590 | 0.741 | 0.999 |
| $\nu$ | Beta | 0.070 | 0.030 | 0.025 | 0.048 | 0.081 |
| $k$ | Gamma | 2.000 | 0.300 | 1.435 | 1.759 | 2.141 |
| $\zeta_\pi$ | Gamma | 3.000 | 0.500 | 1.913 | 2.307 | 2.798 |
| $\zeta_S$ | Gamma | 1.000 | 0.800 | 0.014 | 0.040 | 0.085 |
| $\zeta_i$ | Beta | 0.800 | 0.200 | 0.567 | 0.643 | 0.708 |
| $\zeta_Y$ | Gamma | 1.200 | 0.800 | 0.027 | 0.054 | 0.086 |
| $\phi_A$ | Beta | 0.500 | 0.250 | 0.975 | 0.987 | 0.995 |
| $\phi_X$ | Beta | 0.500 | 0.250 | 0.970 | 0.980 | 0.990 |
| $\phi_{i*}$ | Beta | 0.500 | 0.250 | 0.780 | 0.825 | 0.864 |
| $\phi_\chi$ | Beta | 0.500 | 0.250 | 0.503 | 0.731 | 0.882 |
| $\phi_{CP}$ | Beta | 0.500 | 0.250 | 0.214 | 0.476 | 0.748 |
| $\sigma_A$ | Inverse Wishart | 1.000 | 0.750 | 0.680 | 0.765 | 0.877 |
| $\sigma_X$ | Inverse Wishart | 1.000 | 0.750 | 5.148 | 5.896 | 6.812 |
| $\sigma_i*$ | Inverse Wishart | 1.000 | 0.750 | 0.900 | 1.176 | 1.506 |
| $\sigma_i$ | Inverse Wishart | 1.000 | 0.750 | 0.497 | 0.579 | 0.691 |
| $\sigma_\chi$ | Inverse Wishart | 1.000 | 0.750 | 0.889 | 1.306 | 1.751 |
| $\sigma_{CP}$ | Inverse Wishart | 1.000 | 0.750 | 0.697 | 0.933 | 1.252 |
| $\sigma_{SM}$ | Inverse Wishart | 1.000 | 0.750 | 0.635 | 0.848 | 1.231 |
| Log Marginal Likelihood = −621.4160 | | | | | | |

Source: Authors' estimates.
Note: The parameters $\alpha$, $\beta$, $\theta$, and $\phi_{SM}$ have been calibrated to 0.37, 0.99, 0.01, and 0, respectively.

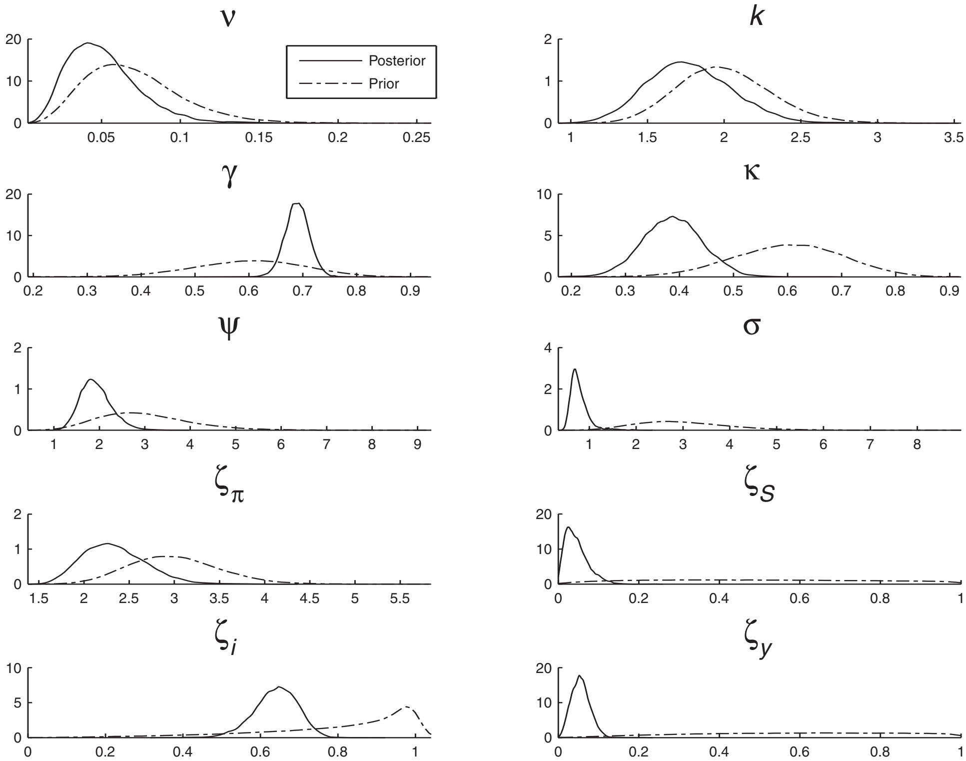
## III. Results

In our baseline model, we abstract from investment adjustment costs and impose full capital depreciation.[21] Although we relax these restrictions below, these assumptions facilitate comparisons with previous calibration-based studies that have incorporated the financial accelerator, namely, Céspedes, Chang, and Velasco (2004), as well as Elekdag and Tchakarov (2004).

Table 1 reports the posterior estimates of each parameter from our baseline model. Along with the medians, we present the 10th and 90th percentiles for the posterior distributions, which serve to quantify the uncertainty surrounding these estimates. Additional information on our results is presented in Figure 1, which plots kernel density estimates for the posteriors, together with the priors, for a subset of the parameters.

[21]With these assumptions, equations (14), (15), and (21) imply that $I_t = K_t$, $q_t = \rho_t$, and $r_t^K = r_t$.

Figure 1. Prior and Posterior Distributions of Selected Parameters




Source: Authors' estimates.

One of the main objectives of this paper is to assess whether the data support the inclusion of the financial accelerator in a small open economy model. This can be gauged by determining whether the elasticity of the external finance premium, ν, is estimated away from zero, because this would support the notion that the financial accelerator captures important credit market frictions. Indeed, Table 1 presents our median estimate for ν as 0.048, with 10th and 90th percentiles of 0.025 and 0.081, respectively—which is different than zero despite our prior. Note in Figure 1 that the posterior is sharply peaked relative to our prior distribution, suggesting that the data are quite informative about this parameter. We interpret the estimate of this critical parameter as strongly supporting the inclusion of a financial accelerator channel in this class of models.[22]

[22]Relatedly, it is encouraging that our estimates are in the range of values used previously by calibration-based studies of the financial accelerator mentioned above. Furthermore, our results are also consistent with those of Meier and Müller (2006).

The parameter $\nu$ could be interpreted as a summary statistic indicating how vulnerable the economy is to shocks affecting aggregate balance sheets. When the economy is hit with an unfavorable shock, the likely exchange rate depreciation and decline in asset prices will raise the leverage ratio and thus the external finance premium. The increase in the premium—determined by the size of $\nu$—further dampens investment and output. The financial accelerator mechanism—acting through balance sheets—adds another channel to an otherwise standard model, which magnifies the impact of shocks on real and financial macroeconomic indicators.

The estimated median external finance premium we obtain is 2.75 percent per quarter, with the 10th and 90th percentiles corresponding to 0.91 percent and 6.36 percent, respectively.[23] The median estimate implies an annualized premium of 11 percent, which may seem somewhat high, although it most likely reflects the impact of the financial crisis when spreads exceeded 900 basis points.[24] However, using the 10th percentile implies an annual premium of 3.6 percent, more in line with the historical average data.[25] We take the model's ability to mimic the Korean external finance premium as quite remarkable considering that this series was not used in the estimation procedure; this further strengthens our confidence in our parameter estimates.

We obtain a reasonable characterization of monetary policy as summarized by a Taylor rule. The median value of the coefficient on inflation, $\zeta_\pi$, is 2.31, which suggests an aggressive response to inflationary pressures and likely reflects the authorities' commitment to price stability. As shown in Figure 1, our estimates also indicate a considerable degree of interest rate smoothing because the parameter $\zeta_i$ is inferred to be 0.64. The estimates of the interest rate response to the output gap, $\zeta_Y$, and exchange rate fluctuations, $\zeta_S$, both concentrate close to the lower bound of zero, despite our flat prior.

Turning to the other parameters, we note that the degree of openness is estimated remarkably well and is summarized by the value of $(1 - \gamma)$. The imports-to-GDP ratio throughout our sample period is 32 percent, whereas the model-based measure of openness, $\gamma$, is estimated to be 31 percent.[26] Our median inferred value of the inverse of the Frisch elasticity of labor supply, $\psi$, is 1.9. Contrary to other studies—including Smets and Wouters (2003)—that have reported difficulties in pinning down this parameter, we find fairly tight posterior probability bands around this estimate. Meanwhile, the inverse of the intertemporal elasticity of substitution, $\sigma$, has a median value of 0.74, which mimics the estimates found in other papers (for instance, see Rabanal, 2006). A 90 percent posterior density interval for the

[23]The median estimate of the capital-to–net worth ratio, $k$, was 1.76, implying a (foreign currency–denominated) debt-to-capital ratio of about 43 percent, consistent with the industry-level evidence discussed above. Also as discussed above, this is one of the key parameters that underpins the external finance premium; see Bernanke, Gertler, and Gilchrist (1999) for further details.

[24]Recall that the EMBIG stripped spread hit 940 basis points at end-August 1998.

[25]The average for the EMBIG from inception until the expiration of the IMF-supported program averaged about 321 basis points.

[26]The import-to-GDP ratio was calculated for Korea using annual data from 1990 to 2003, using IMF *International Financial Statistics* series 98C and 99B.

Calvo price resetting probability, $\kappa$, is fairly tight, covering the 0.32 to 0.46 interval, with a median value of 0.39. This implies that the average duration of a contract is slightly below two quarters, most likely reflecting a relatively higher average inflation rate during our sample period. Turning to the elasticity of substitution between varieties of domestically produced goods, we note that $\lambda$ takes a median value of 8.5, implying a price markup of approximately 13 percent. The 10th and 90th percentiles also cover a wide spectrum of markups ranging from about 8 percent to 22 percent, respectively, which is remarkably close to the values commonly used in calibration-based studies.

Overall, we are able to obtain very reasonable and tight estimates of most parameters. More important—and pertinent to the heart of this paper—our range of estimates for the external finance premium coincides with actual historic outcomes, highlighting the importance of balance sheet–related credit market frictions.

## Robustness Analysis

We check the robustness of our main result—the structural estimate of the external finance premium—by reestimating the model with a different prior for the key parameter of the paper, the elasticity of the premium, $\nu$. As described in Section III, our prior for $k$ is based on firm-level data and we therefore keep it unchanged. In contrast, there is substantive ambiguity on the value of $\nu$, which prompts consideration of an alternative prior for this parameter. In our baseline, the prior for $\nu$ was centered at 0.07 and delivered a median estimate of 0.048. Although the posterior percentiles are bounded away from zero, it is important to recognize that the financial accelerator mechanism gets suppressed as $\nu$ approaches zero. In this case, entrepreneurs would still borrow from abroad in foreign currency, but the cost associated with this source of financing would be given by the foreign interest rate and would not be augmented by a premium.

Therefore, to gauge the robustness of our baseline estimates, we allow for an even looser prior for $\nu$. We consider an alternative specification that is centered at 0.2 and has a standard deviation of 0.1, covering an even wider range of possible values for $\nu$, allowing values of the premium ranging from 1 percent to more than 60 percent. Then, keeping all other priors unchanged, we repeat the estimation procedure.

The results using the alternative prior are presented in Table 2. Our parameter estimates are broadly similar to those in the baseline case. More important, the median estimate for $\nu$ is now 0.077, with a 90 percent posterior probability band in the range of 0.033 and 0.187. Although this may seem like a substantial jump from our earlier estimate, it is important to note that the posterior is concentrated on values well in line with our baseline. This reinforces our conclusion that the data support the inclusion of the financial accelerator, which captures credit market frictions and balance sheet vulnerabilities. Combined with the now lower median estimate of $k$ centered at 1.7, the median estimate of the external finance premium is now about 4 percent, but with wider 90 percent probability bands. Despite a looser prior allowing for very small values of $\nu$, the posterior density is, once again, bounded away from zero, and confirms that our estimate of the external finance premium is not an artifact of the choice of prior.

**Table 2. Parameter Estimates for the Baseline Model Under Alternative Prior**

| | Prior Distribution | | | Posterior Distribution | | |
|---|---|---|---|---|---|---|
| Parameter | Type | Mean | Standard error | 10% | Median | 90% |
| $\gamma$ | Beta | 0.600 | 0.100 | 0.686 | 0.723 | 0.760 |
| $\kappa$ | Beta | 0.600 | 0.100 | 0.349 | 0.426 | 0.493 |
| $\lambda$ | Gamma | 8.000 | 3.000 | 5.536 | 8.387 | 12.623 |
| $\psi$ | Gamma | 3.000 | 1.000 | 2.339 | 2.923 | 3.465 |
| $\sigma$ | Gamma | 3.000 | 1.000 | 0.714 | 1.027 | 1.579 |
| $\nu$ | Beta | 0.200 | 0.100 | 0.033 | 0.077 | 0.187 |
| $k$ | Gamma | 2.000 | 0.300 | 1.355 | 1.697 | 2.094 |
| $\zeta_\pi$ | Gamma | 3.000 | 0.500 | 2.000 | 2.385 | 2.870 |
| $\zeta_S$ | Gamma | 1.000 | 0.800 | 0.014 | 0.041 | 0.082 |
| $\zeta_i$ | Beta | 0.800 | 0.200 | 0.573 | 0.652 | 0.720 |
| $\zeta_Y$ | Gamma | 1.200 | 0.800 | 0.027 | 0.053 | 0.087 |
| $\phi_A$ | Beta | 0.500 | 0.250 | 0.955 | 0.970 | 0.982 |
| $\phi_X$ | Beta | 0.500 | 0.250 | 0.938 | 0.960 | 0.978 |
| $\phi_i{}^*$ | Beta | 0.500 | 0.250 | 0.778 | 0.823 | 0.862 |
| $\phi_\chi$ | Beta | 0.500 | 0.250 | 0.430 | 0.694 | 0.849 |
| $\phi_{CP}$ | Beta | 0.500 | 0.250 | 0.203 | 0.449 | 0.666 |
| $\sigma_A$ | Inverse Wishart | 1.000 | 0.750 | 0.672 | 0.759 | 0.869 |
| $\sigma_X$ | Inverse Wishart | 1.000 | 0.750 | 6.850 | 8.148 | 9.344 |
| $\sigma_i{}^*$ | Inverse Wishart | 1.000 | 0.750 | 0.940 | 1.185 | 1.491 |
| $\sigma_i$ | Inverse Wishart | 1.000 | 0.750 | 0.495 | 0.577 | 0.687 |
| $\sigma_\chi$ | Inverse Wishart | 1.000 | 0.750 | 0.901 | 1.548 | 2.237 |
| $\sigma_{CP}$ | Inverse Wishart | 1.000 | 0.750 | 0.718 | 0.996 | 1.385 |
| $\sigma_{SM}$ | Inverse Wishart | 1.000 | 0.750 | 0.628 | 0.858 | 1.241 |
| Log Marginal Likelihood = −624.5200 | | | | | | |

Source: Authors' estimates.
Note: The parameters $\alpha$, $\beta$, $\theta$, and $\phi_{SM}$ have been calibrated to 0.37, 0.99, 0.01, and 0, respectively.

## Augmenting the Model

In the baseline model, we assumed that there were no investment adjustment costs, and we imposed full depreciation. As mentioned above, this was done so that our model could be readily compared with previous models that incorporated the financial accelerator. However, this raises the possibility that the exclusion of realistic investment dynamics may artificially enhance or subdue the role of the financial accelerator. In this context, are the relatively high values of the estimated external finance premium an outcome of these stringent restrictions?

In this section, we address these issues by augmenting the model to allow for the dynamics in equations (14), (15), and (21) in contrast to the baseline case. The addition of these three equations now entails the estimation of an additional parameter, the elasticity of Tobin's $q$ with respect to the investment-to-capital ratio, $\Omega$.[27]

[27]More specifically, $\Omega = \Phi''(\delta)\delta/\Phi'(\delta)$, and as do Smets and Wouters (2003), we calibrate the depreciation rate, $\delta$, to 0.025.

Table 3. Parameter Estimates for the Augmented Model

| Parameter | Prior Distribution | | | Posterior Distribution | | |
|---|---|---|---|---|---|---|
| | Type | Mean | Standard error | 10% | Median | 90% |
| $\gamma$ | Beta | 0.600 | 0.100 | 0.604 | 0.658 | 0.708 |
| $\kappa$ | Beta | 0.600 | 0.100 | 0.332 | 0.415 | 0.482 |
| $\lambda$ | Gamma | 8.000 | 3.000 | 4.019 | 7.082 | 11.284 |
| $\psi$ | Gamma | 3.000 | 1.000 | 1.846 | 2.439 | 3.341 |
| $\sigma$ | Gamma | 3.000 | 1.000 | 0.266 | 0.332 | 0.414 |
| $\nu$ | Beta | 0.070 | 0.030 | 0.030 | 0.054 | 0.087 |
| $k$ | Gamma | 2.000 | 0.300 | 1.517 | 1.822 | 2.177 |
| $\Omega$ | Normal | 0.300 | 0.100 | 0.320 | 0.418 | 0.526 |
| $\zeta_\pi$ | Gamma | 3.000 | 0.500 | 2.250 | 2.675 | 3.204 |
| $\zeta_S$ | Gamma | 1.000 | 0.800 | 0.008 | 0.028 | 0.066 |
| $\zeta_i$ | Beta | 0.800 | 0.200 | 0.536 | 0.628 | 0.702 |
| $\zeta_Y$ | Gamma | 1.200 | 0.800 | 0.015 | 0.036 | 0.065 |
| $\phi_A$ | Beta | 0.500 | 0.250 | 0.837 | 0.913 | 0.966 |
| $\phi_X$ | Beta | 0.500 | 0.250 | 0.955 | 0.981 | 0.995 |
| $\phi_{i*}$ | Beta | 0.500 | 0.250 | 0.889 | 0.930 | 0.961 |
| $\phi_\chi$ | Beta | 0.500 | 0.250 | 0.283 | 0.698 | 0.907 |
| $\phi_{CP}$ | Beta | 0.500 | 0.250 | 0.437 | 0.857 | 0.953 |
| $\sigma_A$ | Inverse Wishart | 1.000 | 0.750 | 1.161 | 1.304 | 1.464 |
| $\sigma_X$ | Inverse Wishart | 1.000 | 0.750 | 6.618 | 7.576 | 8.764 |
| $\sigma_{i*}$ | Inverse Wishart | 1.000 | 0.750 | 0.936 | 1.141 | 1.401 |
| $\sigma_i$ | Inverse Wishart | 1.000 | 0.750 | 0.549 | 0.649 | 0.791 |
| $\sigma_\chi$ | Inverse Wishart | 1.000 | 0.750 | 0.696 | 1.083 | 1.771 |
| $\sigma_{CP}$ | Inverse Wishart | 1.000 | 0.750 | 0.773 | 1.245 | 1.777 |
| $\sigma_{SM}$ | Inverse Wishart | 1.000 | 0.750 | 0.622 | 0.834 | 1.202 |

Log Marginal Likelihood = −641.7723

Source: Authors' estimates.

Note: The parameters $\alpha$, $\beta$, $\delta$, $\theta$, and $\phi_{SM}$ have been calibrated to 0.37, 0.99, 0.025, 0.01, and 0, respectively.

The estimation results of the extended model are presented in Table 3. When compared with the baseline model estimates, it is reassuring that in general the parameter estimates are quite similar. Once again, $\nu$ is tightly estimated away from zero with a median value of 0.054 and with 90 percent probability bands ranging from 0.030 to 0.087. We interpret this estimate in a more general model as further evidence that the data support the inclusion of the financial accelerator; this reinforces our previous conclusions.

Table 3 reveals some other interesting results. First, the elasticity of Tobin's *q* with respect to the investment-to-capital ratio, $\Omega$, is estimated to be about 0.42, which is larger than those values used in many calibration-based studies. Lubik and Teo (2005) find a similar result and argue that although calibration-based papers set this parameter to match specific features in the data, in the context of *overall* fit

Table 4. Parameter Estimates for the Augmented Model Under Alternative Prior

| | Prior Distribution | | | Posterior Distribution | | |
|---|---|---|---|---|---|---|
| Parameter | Type | Mean | Standard error | 10% | Median | 90% |
| $\gamma$ | Beta | 0.600 | 0.100 | 0.608 | 0.661 | 0.712 |
| $\kappa$ | Beta | 0.600 | 0.100 | 0.345 | 0.421 | 0.484 |
| $\lambda$ | Gamma | 8.000 | 3.000 | 4.088 | 7.022 | 11.197 |
| $\psi$ | Gamma | 3.000 | 1.000 | 1.928 | 2.500 | 3.343 |
| $\sigma$ | Gamma | 3.000 | 1.000 | 0.254 | 0.319 | 0.400 |
| $\nu$ | Beta | 0.200 | 0.100 | 0.041 | 0.086 | 0.178 |
| $k$ | Gamma | 2.000 | 0.300 | 1.383 | 1.686 | 2.061 |
| $\Omega$ | Normal | 0.300 | 0.100 | 0.332 | 0.431 | 0.530 |
| $\zeta_\pi$ | Gamma | 3.000 | 0.500 | 2.240 | 2.652 | 3.162 |
| $\zeta_S$ | Gamma | 1.000 | 0.800 | 0.008 | 0.026 | 0.057 |
| $\zeta_i$ | Beta | 0.800 | 0.200 | 0.530 | 0.624 | 0.698 |
| $\zeta_Y$ | Gamma | 1.200 | 0.800 | 0.017 | 0.038 | 0.068 |
| $\phi_A$ | Beta | 0.500 | 0.250 | 0.829 | 0.905 | 0.961 |
| $\phi_X$ | Beta | 0.500 | 0.250 | 0.940 | 0.972 | 0.992 |
| $\phi_i{}^*$ | Beta | 0.500 | 0.250 | 0.891 | 0.932 | 0.963 |
| $\phi_\chi$ | Beta | 0.500 | 0.250 | 0.282 | 0.689 | 0.911 |
| $\phi_{CP}$ | Beta | 0.500 | 0.250 | 0.470 | 0.853 | 0.949 |
| $\sigma_A$ | Inverse Wishart | 1.000 | 0.750 | 1.169 | 1.306 | 1.475 |
| $\sigma_X$ | Inverse Wishart | 1.000 | 0.750 | 6.705 | 7.623 | 8.789 |
| $\sigma_i{}^*$ | Inverse Wishart | 1.000 | 0.750 | 0.949 | 1.175 | 1.459 |
| $\sigma_i$ | Inverse Wishart | 1.000 | 0.750 | 0.550 | 0.652 | 0.789 |
| $\sigma_\chi$ | Inverse Wishart | 1.000 | 0.750 | 0.694 | 1.038 | 1.777 |
| $\sigma_{CP}$ | Inverse Wishart | 1.000 | 0.750 | 0.813 | 1.282 | 1.756 |
| $\sigma_{SM}$ | Inverse Wishart | 1.000 | 0.750 | 0.631 | 0.859 | 1.257 |

Log Marginal Likelihood = −643.2308

Source: Authors' estimates.

Note: The parameters $\alpha$, $\beta$, $\delta$, $\theta$, and $\phi_{SM}$ have been calibrated to 0.37, 0.99, 0.025, 0.01, and 0, respectively.

through model estimation, the data seem to support larger values of $\Omega$. Second, note that the persistence, $\phi_{cp}$, and volatility, $\sigma_{cp}$, of the cost-push shock increases compared with the baseline, but to counteract these results, we also estimate a higher interest rate response to inflation, $\zeta_\pi$. Finally, with a higher value of $k$, the associated median external finance premium is 3.29 percent—larger than, but still in line with, the baseline estimates.

Analogous to the sensitivity analysis under the baseline specification, we also reestimate the augmented model using the looser prior for the parameter $\nu$ mentioned above. The results are presented in Table 4 and are, once again, broadly similar to the previous experiments. As before, $\nu$ is inferred away from zero; with the median value $k$, it yields an external finance premium of 4.6 percent. It seems that although adding investment dynamics relaxes some of the stringent features of the baseline model, it also imposes other cross-equation restrictions that result in a higher premium. We elaborate on this further in the next section.

In summary, all estimates, from the baseline and extended version of the model under both sets of priors, suggest a prominent role for the financial accelerator mechanism as manifested by a sizable external finance premium. Indeed, although our estimates of $\nu$ and $k$ may seem rather different in terms of medians, the posterior bands imply overlapping values of the premium.[28]

## Cross-Validation with a BVAR Model

As suggested by Schorfheide (2000), it is desirable to compare the fit of the estimated DSGE model with the one resulting from the estimation of a more densely parameterized and less restrictive reference framework, which is usually taken to be a Bayesian Vector Autoregression (BVAR). Comparing the fit of the model with a BVAR permits, for instance, an evaluation of how useful a DSGE model might be in formulating policy. Indeed, Smets and Wouters (2003) argue that a full-fledged microfounded model can describe macroeconomic aggregates as well as, if not better than, a BVAR, which has led to an increased interest in the role of models for policymaking (see Sims, 2003a). It is important to recognize, however, that at a deeper level this cross-validation is done to control for the possible misspecification of all DSGE models under consideration (for details, see Schorfheide, 2000).

Model comparisons in a Bayesian setting are achieved by computing the posterior model probabilities, which, in the case of equal prior probabilities across models, becomes the ratio of the marginal likelihoods across different models or specifications. Computing the marginal likelihood usually requires relying on approximations such as Laplace asymptotics or simulation-based methods. In this paper, we estimate the marginal density using the Modified Harmonic Mean proposed by Geweke (1998) and then use the draws that are generated for the estimation of the model. Our estimate of the marginal density is presented in the top four rows of Table 5 and is obtained by averaging the estimates of a grid of values between 0.1 and 0.9 for the truncation used to bound the reciprocal importance sampling density. The marginal densities of the BVARs are displayed in the bottom four rows of Table 5. Closed form solutions for the marginal likelihood of BVARs are available; their availability facilitates using various lag lengths, which in our case range between one and four. Our computations follow Sims and Zha (2004) where we use a symmetrized version of the Minnesota prior and include dummy priors to control for the persistence of the data.

Echoing the results of Smets and Wouters (2003), we find that our baseline model does at least as good a job as the BVARs in describing the evolution of the five observable variables over the sample period. This is because the log marginal likelihood of the baseline model (including under the alternative prior) is larger than the best-performing BVAR, which has one lag, as depicted in Table 5.

[28]Finally, we ask how the models presented above compare with one in which the financial accelerator mechanism is shut down, that is, $\nu \rightarrow 0$. We find that the marginal likelihood implies that the posterior odds favor the model *with* the financial accelerator fully operational ($\nu >> 0$) by a ratio of 13 to 1.

Table 5. Cross-Validation

| Model | Log Marginal |
|---|---|
| Baseline model | −621.4160 |
| Alternative prior | −624.5200 |
| Extended model | −641.7723 |
| Alternative prior | −643.2308 |
| BVAR(1) | −624.7053 |
| BVAR(2) | −653.7563 |
| BVAR(3) | −672.0307 |
| BVAR(4) | −690.1469 |

Source: Authors' estimates.

Note: The log marginal likelihood is obtained with a modified-harmonic mean estimator. The results are insensitive to the choice of cutoff point for the reciprocal importance sampling density and correspond to the mean of a grid that takes values between 0.1 and 0.9. Results are also robust to whether the density point corresponds to the mean, median, or the (simulated) maximizing value of the posterior draws.

Although this result is encouraging and further supports the inclusion of the financial accelerator into DSGE models, we also uncover a puzzling outcome. The log marginal likelihood of the extended model (and the version with the alternative prior) is lower than that of the BVARs. It seems that although investment dynamics add flexibility to model in some dimensions, they require the estimation of an additional parameter and—more important—add further cross-equation restrictions. One explanation is that the investment adjustment costs prevent dramatic swings in investment and thereby limit the model's ability to capture the large drop in investment during the crisis. In this respect, we view the financial accelerator as a key mechanism for explaining balance sheet–related vulnerabilities, although we must recognize that linearized models may have difficulties matching sharp and abrupt declines in real activity and, in particular, in investment.

## IV. Concluding Remarks

We use Bayesian methods to estimate a small open economy model with a financial accelerator mechanism that serves to capture balance sheet–related credit market frictions. Using Korean data, we obtain a sizable value for the external finance premium, which is tightly estimated away from zero. This implies that the inclusion of the financial accelerator into an otherwise standard small open economy model is supported by the data. This result is robust to model specification as well as to alternative prior densities. Furthermore, we provide evidence that our baseline model with the financial accelerator provides a fit of the data that outperforms the best BVAR. Our findings emphasize the importance of financial frictions that magnify the impact of shocks, thereby exacerbating real and financial volatility. These results highlight the role of the latent balance sheet vulnerabilities that are believed to have exacerbated the Korean crisis.

Although the estimation of the model has yielded significant insights into the dynamics of credit market frictions and balance sheet vulnerabilities, future research should include augmenting the model to capture even richer dynamics. We thus hope to refine our parameter estimates, especially with regard to the external finance premium, by incorporating endogenous persistence through habit formation as well as price and wage indexation. We would also like to extend our framework to a multicountry setup and consider estimating the model using data from other EMCs. Such extensions could also guide the choice of parameter values when conducting welfare-based comparisons of monetary policies, and sharpen the debate on the appropriate choice of exchange rate regimes for EMCs.

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