WP/05/44

IMF Working Paper

An Estimated Small Open Economy Model of the Financial Accelerator

Selim Elekdag, Alejandro Justiniano, and Ivan Tchakarov

INTERNATIONAL MONETARY FUND

 WP/05/44

IMF Working Paper

Asia and Pacific Department

An Estimated Small Open Economy Model of the Financial Accelerator

Prepared by Selim Elekdag, Alejandro Justiniano, and Ivan Tchakarov [1]

Authorized for distribution by Jahangir Aziz

March 2005

Abstract

This paper develops a small open economy model where entrepreneurs partially finance investment using foreign currency denominated debt subject to a risk premium above and beyond international interest rates. We use Bayesian estimation techniques to evaluate the importance of balance sheet vulnerabilities combined with the presence of the financial accelerator for emerging market countries. Using Korean data, we obtain an estimate for the external risk premium, indicating the importance of the financial accelerator and potential balance sheet vulnerabilities for macroeconomic fluctuations. Furthermore, our estimates of the Taylor rule imply a strong preference to smooth both exchange rate and interest rate fluctuations.

JEL Classification Numbers: C11; F41

Keywords: Financial accelerator, balance sheets, Bayesian estimation of DSGE models, Taylor rules

Author(s) E-Mail Address: selekdag@imf.org; ajustiniano@imf.org; itchakarov@imf.org

---

[1] Selim Elekdag is an Economist in the Policy Development and Review Department, Alejandro Justiniano is an Economist in the Western Hemisphere Department, and Ivan Tchakarov is an Economist in the Asia and Pacific Department. We are grateful to Gian Maria Milesi-Ferretti, Douglas Laxton, Alessandro Rebucci, and other colleagues in the Research Department of the International Monetary Fund for helpful comments and discussions.

Contents Page

I. Introduction

After episodes of severe crises that have plagued many emerging market economies in recent years, there has been renewed interest in the classic question of the most appropriate monetary policy regime. 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 shocks.

However, emerging market economies face two fundamental issues that complicate the conduct of monetary policy. First, these countries can typically borrow only in foreign currency denominations, labeled "original sin" by Eichengreen and Hausman (1999). This feature increases the susceptibility to external shocks since a potential depreciation can substantially inflate debt service costs—due to currency mismatches—and thus increase rollover risk. Second, emerging market countries usually have imperfect access to capital markets. Foreign credit is typically associated with a risk premium above and beyond the international lending rate. Moreover, this premium typically moves in tandem with the business cycles of the borrowing country, which implies a higher risk premium when an emerging market country (EMC) is experiencing a recession.[2] Hence, when an EMC is in dire need of macroeconomic stimulus, the higher risk premium exacerbates the severity of the recession by increasing the opportunity cost of investment and thus choking off the process of capital accumulation.

As a result of outstanding foreign currency denominated debt, entrepreneurs in EMCs may find that exchange rate fluctuations have large effects on their net worth positions. As emphasized by Krugman (1999) as well as Aghion, Bacchetta, and Banerjee (2001), among others, exchange rate and interest rate fluctuations—through balance sheet constraints affecting investment spending—have much more serious macroeconomic consequences than for entrepreneurs in industrialized economies. One way EMCs apparently deal with such excessive vulnerability is to minimize exchange rate fluctuations by engaging in what Calvo and Reinhart (2001) call "fear of floating." These authors argue that the reluctance to implement a pure float could be justified by the fact that large exchange rate movements may devastate corporate and financial balance sheets due to large outstanding foreign currency denominated debt obligations.[3]

There has been a recent surge in the number of papers attempting to model the behavior of EMCs. The work of Cespedes, Chang, and Velasco (2004), Devereux, Lane, and Xu (2004), Elekdag and Tchakarov (2004), as well as Gertler, Gilchrist, and Natalucci (2003) are particularly notable contributions. Building upon the model developed by Bernanke, Gertler, and Gilchrist (1999), these authors explore the role of balance sheets and imperfect capital market access in investment financing for EMCs in an open economy context.

---

[2] In fact, the correlation of real GDP growth and the EMBI spread for Korea was -0.33 during the sample spanning the fourth quarter of 1997 to the fourth quarter of 2000. Recall that the EMBI spread for Korea was not calculated before the fourth quarter of 1997.

[3] Calvo (1999) also highlights the reluctance of emerging market economies to pursue a clean float.

The key element in all of these papers is that lenders must incur a cost to monitor the business activity of borrowers due to information asymmetries. Thus lenders must be compensated in the form of a risk premium that the borrower has to shoulder on top of the foreign interest rate. This risk premium in turn depends on the ratio of debt to net worth which underpins what is known as the financial accelerator.[4] In this context, an external shock which triggers an exchange rate depreciation could generate a vicious cycle. Balance sheets with significant currency mismatches would imply that a depreciation would increase the value of foreign currency debt which would erode the value of domestic currency denominated net worth. The deterioration in net worth would increase the risk premium, increasing the cost of financing capital projects and therefore causes substantial macroeconomic instability.

Against this background, the main question we ask in this paper is whether there is evidence that favors the incorporation of the financial accelerator in an open economy setting. If the data support the importance of collateral constraints, then models that incorporate an endogenous risk premium and foreign currency denominated debt could yield important insights on the debate regarding the most appropriate exchange rate regime for EMCs. To this end, we estimate a stylized model which includes the financial accelerator channel and assess how well our model is supported by the data.

Despite the potential relevance of the financial accelerator, the literature cited above has used calibrated models to highlight the implications of collateral constraints. The main contribution of this paper is therefore the estimation of a model incorporating the financial accelerator which allows us to *infer* the average risk premium. The sensitivity of the risk premium to the capital to net worth ratio in turn sheds light on the potential vulnerabilities of the economy to shocks affecting aggregate balance sheets.

Relying on recent developments in Bayesian estimation techniques following Geweke (1999), Shorfheide (2000), Smets and Wouters (2003), as well as Justiniano and Preston (2004), we estimate our model using Korean data. To our knowledge, this is the first paper that bridges the gap between theory and empirical research in the context of the financial accelerator and the presence of a balance sheet channel.[5] The main advantages of the Bayesian methodology is that it allows a complete characterization of uncertainty in estimating structural parameters by simulating posterior distributions. It also provides an elegant way to incorporate prior information about parameters coming either from microeconomic studies or from previous macroeconomic exercises, and therefore creates a direct connection with the calibration-based literature and rigorous policy analysis.[6]

---

[4] To ease exposition, we sometimes refer to the combination of the financial accelerator channel and existence of foreign currency denominated debt as collateral constraints. We also abuse this convention and sometimes refer to the financial accelerator with the understanding that (entrepreneurial) debt is always denominated in foreign currency.

[5] Related research includes Linaa and Rand (2004) as well as Tovar (2004).

[6] Furthermore, as emphasized by Smets and Wouters (2003), the use of Bayesian methods provides greater stability to maximum likelihood algorithms for the estimation of the parameters.

Our main findings can be summarized as follows. First, based on Korean data, the mean estimate of the endogenous risk premium is 3.27 percent per quarter with the 10th and 90th percentiles corresponding to 1.93 and 4.87 percent, respectively. The existence of a significant average risk premium indicates that there is evidence in favor of the financial accelerator mechanism and that the balance sheet channel matters. Although the annualized risk premium corresponds to 13.1 percent–most likely reflecting the impact of the East Asian crisis–using the 10th percentile implies an annual risk premium of 7.95, which is much more in line with the Korean data.

Second, our estimate of the elasticity of the risk premium to the capital to net worth ratio is 0.066, which is in line with calibration-based models incorporating the financial accelerator.[7] This elasticity could be interpreted as a summary statistic indicating how vulnerable the economy is to shocks affecting aggregate balance sheets. Moreover, the significance of this parameter highlights the sensitivity of partially foreign credit financed investment spending to exogenous shocks, which could be amplified further through the presence of currency mismatches and limited market access.

Third, in the context of Taylor rules, the parameter that determines the interest rate response to exchange rate fluctuations is estimated away from zero. This could be interpreted as evidence in favor of the "fear of floating" hypothesis, perhaps reflecting that importance of potential balance sheet vulnerabilities. The other parameters in the Taylor rule–including those that determine the authorities' desire to smooth interest rate fluctuations–are also in line with the managed float implemented by the Korean authorities' throughout the sample period we analyze.

Finally, the median estimates for the main behavioral parameters are broadly consistent with values found in previous literature. In particular, the intertemporal elasticity of substitution, the elasticity of labor supply, as well as the parameters governing price stickiness and the level of monopolistic competition are estimated reasonably well despite loose priors.

Section II lays out the model, Section III describes the estimation methodology and the results, and Section IV concludes.

II. The Model

Our modeling framework is an extension of Cespedes, Chang, and Velasco (2004), where we focus on a small open economy with a representative household, producers, entrepreneurs, and a central bank.[8] The consumption good is a composite of a domestically produced good and an imported foreign good. The domestic good is a bundle which is composed of a continuum of goods produced by domestic firms in a monopolistically competitive environment. Since 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 which uses an interest rate rule

---

[7] As is shown in Table 2, we estimate the 10th and 90th percentiles to be 0.039 and 0.097, respectively.

[8] The augmented modeling framework is based on Elekdag and Tchakarov (2004).

to achieve specific policy objectives. The representative household is allowed to accumulate financial assets and is thus responsive to interest rate fluctuations.

A. 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:

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

where E_t denotes the mathematical expectation conditional on information available in period t, $\beta \in [0, 1]$ is the subjective discount factor, and η_t is a labor disutility shock.[9]

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

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

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

Following Cespedes, Chang, and Velasco (2004), 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, so the domestic currency price of imports is just equal to the nominal exchange rate s_t. The aggregate price level, q_t, is then derived by solving for the minimum expenditure required to obtain one unit of the aggregate consumption good.[10] Denoting the price of the of the domestic good as p_t, the aggregate price level is then:

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

The domestic good, however, is itself a composite good, consisting of a CES aggregate of the

---

9 To be discussed below.

10 Given the aggregate price index defined in equation 3, the individual consumption demands for each good are:

$$\begin{aligned} C_{Ht} &= \gamma \left(\frac{p_t}{q_t} \right)^{-1} C_t \\ C_{Ft} &= (1-\gamma) \left(\frac{s_t}{q_t} \right)^{-1} C_t. \end{aligned}$$

continuum of differentiated domestic goods, more specifically:

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

where $j \in [0,1]$ and $\lambda_t > 1$.[11]

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

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

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. Since they own the domestic production firms, they retain any profits, Π_t. Finally, households must incur an intermediation cost, Γ_t, which has the following specification:

$$\Gamma_t = \frac{\phi_B}{2} \left(\frac{b_t}{q_t} \right)^2 + \frac{\phi_{F^*}}{2} \left(\frac{s_t F_t^*}{q_t} \right)^2 \tag{6}$$

where $\phi_B, \phi_{F^*} \geqslant 0$ and $\phi_B + \phi_{F^*} > 0$. 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 (1) subject to the constraint (5) and initial values for b_0 and F_0^*. Ruling out Ponzi type schemes, we get the following first-order conditions:

$$\eta_t \chi \frac{L_t^{\psi}}{C_t^{-\sigma}} = \frac{w_t}{q_t} \tag{7}$$

$$C_t^{-\sigma} [1 + \phi_B \left(\frac{b_t}{q_t} \right)] = \beta (1 + i_t) E_t [C_{t+1}^{-\sigma} \frac{q_t}{q_{t+1}}] \tag{8}$$

[11] More specifically in the log-linearized version of the model we consider:

$$\begin{aligned} \lambda_t &= \lambda + cp_t \\ cp_t &= \phi_{cp} cp_{t-1} + \varepsilon_{cpt} \end{aligned}$$

with ε_{cpt} can be interpreted as a cost-push shock to price inflation.

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

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

B. 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:

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

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 to be described below. Production firms exploit their market power and set prices in order to maximize profits along a downward sloping demand curve given by:

$$Y_{jt} = \left(\frac{P_{jt}}{P_t}\right)^{-\lambda_t} Y_t \tag{11}$$

Denoting mc_t as the common marginal cost for all firms in the economy, cost minimizing behavior implies the following optimally 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, denotes the rental rate on capital. Equations 12 and 13 are implicit demand curves for labor and capital, respectively.

C. Price Setting

Following 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 $\overline{\pi}$.

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 functional:

$$E_t \sum_{\tau=0}^{\infty} \kappa^{\tau} \Lambda_{t,t+\tau} \left[\overline{\pi}^{\tau} p_t(j) - mc_{t+\tau} \right] Y_{t+\tau}(j) \tag{14}$$

In the case where $\kappa \rightarrow 0$, the optimal choice for $p_t(j)$ becomes a simple markup rule:

$$p_t(j) = \frac{\lambda_t}{\lambda_t - 1} mc_t \tag{15}$$

which implies that there is a markup of price over marginal cost.[12]

D. Entrepreneurs

One of our main objectives in this paper is to uncover evidence of balance sheet effects. To this end, we adapt the modeling framework introduced by Cespedes, 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 Cespedes, Chang, and Velasco (2004) along with 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 partially financed with net worth, and partially with foreign currency denominated debt. Hence the entrepreneur is subject to the following budget constraint:

$$p_t N_t = q_t K_t - s_t D_t^* \tag{16}$$

---

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

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

where the discount factor is formally defined as:

$$\Lambda_{t,t+\tau} = \beta \frac{C_t}{C_{t+\tau}}.$$

where, N_t and D^*_{t+1}, denote net worth and foreign currency denominated debt, respectively. It is important to notice that equation 16 is a standard accounting identity which states that net worth is equal to assets minus foreign currency denominated liabilities. Notice that an unanticipated depreciation–an increase in s_t–will inflate liabilities and reduce net worth. This highlights one source of vulnerability as net worth is susceptible to exchange rate fluctuations.

Entrepreneurs are risk neutral and choose D^*_{t+1} and K_{t+1} to maximize profits.[13] Due to informational asymmetries, the expected return to investment is equal to the foreign interest rate adjusted for expected depreciations, augmented by a risk premium, that is:

$$E_t\left[\frac{r_{t+1}}{q_t}\right] = \Psi(\frac{q_t K_t}{p_t N_t})(1+i^*_t)E_t\left[\frac{s_{t+1}}{s_t}\right] \tag{17}$$

where, $\Psi(\cdot)$, denotes the risk premium which satisfies $\Psi(1) = 1$ and $\Psi'(\cdot) > 0$. Notice that the premium $\Psi(\cdot)$ depends on the capital-to-net worth ratio $k = qK/pN$.[14]

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

$$p_t N_t = \delta_t\left[r_t K_{t-1}(1-\varrho_t) - (1+i^*_{t-1})\Psi(\frac{q_{t-1}K_{t-1}}{p_{t-1}N_{t-1}})s_t D^*_{t-1}\right] \tag{18}$$

where, as in Cespedes, Chang, and Velasco (2004), ϱ_t, reflects the deadweight cost, associated with imperfect capital markets.[15] As can be seen from equation 18, liabilities are susceptible to foreign interest rate and exchange rate fluctuations.[16] An unanticipated increase in the foreign interest rate and/or a sudden depreciation can inflate the debt service obligations of the entrepreneur potentially deteriorating net worth, thus increasing the risk premium. This increases the opportunity cost of investment, which hampers capital accumulation thus exacerbating the severity of the potential recession.

---

[13] Capital is assumed to depreciate completely in production.

[14] 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 16.

[15] Actually, it is the cost associated with monitoring and is an increasing function of the risk premium. See Cespedes, Chang, and Velasco (2004) which is what our presentation is based on. Furthermore, Gerter, Gilchrist, and Natalucci (2003) provides additional details as well as novel extensions. Finally, see Bernanke, Gertler, and Gilchrist (1999) for the full exposition.

[16] Notice that $(1-\delta_t)$ represents that bankruptcy rate of the entrepreneurs, in the log-linearized version of the model δ_t takes the following form:

$$\delta_t = \delta + \varepsilon_{\delta t}$$

where $\varepsilon_{\delta t}$ could be interpreted as a shock to the discount rate of the entrepreneurs. In the text we refer to this as a bankruptcy shock, see Gertler, Gilchrist, and Natalucci (2003) for an analogous interpretation.

E. The Central Bank

In order to conduct more realistic monetary policy analysis, we include a central bank in our model. The central bank implements a general interest rate rule in order to achieve specific policy objectives. The interest rate rule takes the following form:

$$i_t = \bar{i} + \zeta_i i_{t-1} + \zeta_\pi \widehat{\pi}_t + \zeta_S(\widehat{s}_t - \widehat{s}_{t-1}) + \zeta_Y \left(\widehat{Y}_t - Y\right) + \varepsilon_{it} \tag{19}$$

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 periods interest rate. Notice that when $\zeta_\pi \to \infty$ the central bank is implementing strict inflation targeting and when $\zeta_s \to \infty$ the bank is implementing a fixed exchange rate regime. If, ζ_π, is finite and $\zeta_S > 0$ a managed float is being implemented. Finally, if $\zeta_i > 0$, then the central bank is engaging in interest rate smoothing. We denote by ε_{it} a domestic monetary policy shock

F. Market Clearing

Domestic expenditure on home goods is a fraction γ of final expenditures. The home good is also sold to foreigners, demand of which is assumed to be an exogenous process X_t. The market clearing condition for home goods is thus:

$$p_t Y_t = \gamma q_t \left[K_{t+1} + C_t\right] + s_t X_t \tag{20}$$

We close the model by imposing a market clearing condition for domestic bonds and specification for the exogenous variables.[17]

Assuming all firms behave symmetrically, the stationary rational expectations

---

[17] The domestic bond clearing condition implies:

$$b_t = 0$$

The exogenous variables are all assumed to be $AR(1)$ processes:

$$\begin{aligned} i_t^* - i^* &= \phi_{i^*}\left(i_{t-1}^* - i^*\right) + \varepsilon_{i^*t} \\ A_t &= \phi_A A_{t-1} + \varepsilon_{At} \\ X_t &= \phi_X X_{t-1} + \varepsilon_{Xt} \\ \eta_t &= \phi_\eta \eta_{t-1} + \varepsilon_{\eta t} \\ cp_t &= \phi_{cp} cp_{t-1} + \varepsilon_{cpt} \end{aligned}$$

Finally, to close the model, we need to explicitly define the inflation rate:

$$\pi_t = \frac{q_t}{q_{t-1}}.$$

equilibrium is a set of stationary stochastic processes $\{q_t, p_t, s_t, b_t, F_t^*, D_t^*, w_t, L_t, r_t, i_t, i_t^*, Y_t, K_t, N_t, mc_t, \pi_t, A_t, X_t\}_{t=0}^{\infty}$ satisfying equations (3), (5), (7)-(10), (12)-(13), (16)-(19) and the market clearing conditions along with initial values b_0, F_0^*, and D_0^*.

III. Estimation Methodology

We estimate the model using Bayesian methods based on the influential work of Shorfheide (2000). Papers using a Bayesian approach in the estimation of open economy dynamic stochastic general equilibrium (DSGE) models include Lubik and Shorfheide (2003) and Justiniano and Preston (2004). There are several advantages of using Bayesian methods for inference in estimating macroeconomic models. For our purposes, we highlight that since Bayesian methods seek to characterize the posterior distribution of the parameters, they facilitate an accurate assessment 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.[18]

We briefly sketch our approach to inference and the reader is referred to the above references for further details. Defining Θ 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, θ, is given by

$$p\left(\theta|Y^T\right) = \frac{\mathcal{L}\left(\theta|Y^T\right)p(\theta)}{\displaystyle\int \mathcal{L}\left(\theta|Y^T\right)p(\theta)d\theta}$$

where $\mathcal{L}\left(\theta|Y^T\right)$ 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.[19]

Our goal is to therefore characterize the posterior density of the parameters. In order to do so we follow a two-step approach. In the first step a numerical algorithm is used to *approximate* the posterior mode by combining the likelihood $\mathcal{L}(Y^T|\theta)$ with the prior. The posterior mode obtained from this first step is used as the starting value (θ^0) of a 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 used to draw a new candidate parameter θ^* is a $N(\theta^i, c\Sigma)$. The new draw is then accepted with probability

[18] There are also clear advantages when it comes to model comparisons, since 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 which explicitly incorporate model uncertainty.

[19] The Sims (2003b) solution method is used for the DSGE model.

$$\alpha = \min(1, \frac{\mathcal{L}(Y^T|\theta^*)\pi(\theta^*)}{\mathcal{L}(Y^T|\theta^i)\pi(\theta^i)})$$

If accepted, $\theta_k^{i+1} = \theta_k^*$; otherwise, $\theta_k^{i+1} = \theta_k^i$. We generate chains of 70,000 draws in this manner discarding the first 10,000 iterations.

Following Gelman, Carlin, Stern, and Rubin (1995) the variance covariance matrix Σ is updated after a couple of thousand draws and the scaling constant c adjusted to obtain an acceptance rate of between 25 to 35 percent. With the generated draws, point estimates of θ can be obtained from the simulated values by using various location measures, such as mean or medians. Similarly, measures of uncertainty follow from computing the percentiles of the draws.

A. Data

To estimate the model we use five key macroeconomic time series for South Korea: real GDP, inflation, hours worked, the nominal interest rate, and the real CPI based exchange rate.[20] The average of Korean money market annualized interest rate and the annualized log percentage change in the GDP deflator correspond to the interest and inflation rate, respectively. Meanwhile, real GDP, the real exchange rate, and hours worked are expressed in log-deviations from a linear time trend. We fit these series to the seven shocks given by technology, foreign demand, foreign interest rate, labor disutility, cost-push, monetary policy, and the bankruptcy shock. The first five shocks exhibit autoregressive dynamics, while the remaining disturbances are white noise processes.

The sample runs from the first quarter of 1990 to third quarter of 2003. Even though longer time series for the five observable variables were available, the sample is restricted, given that prior to this starting date a fixed exchange rate regime was implemented.[21] Throughout the 1980s Korea pegged the nominal exchange rate to a basket of major currencies. In March 1990, Korea switched to a managed float, with bands of 2.25 percent around the won to dollar exchange rate. 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 initially focus on the sample period during which Korea was pursuing a managed or a free float.[22]

The most striking feature of our data sample is the turbulence generated by the Korean crisis. As in other capital account crises, the Korean data display standard features associated with such

---

20 All series are extracted from Datastream international.

21 Note, nonetheless, that we use observations from 1988-1989 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. Future extensions to this paper will consider a longer sample.

22 It might be argued that we should have restricted ourselves to only either a period of pure float or a managed float. Markov switching methods might allow for incorporating the transition to an alternative exchange rate policy; however, our limited sample prevents us from considering the estimation of the model under these two scenarios. Furthermore, this would require a nonlinear model for estimation. See Schorfheide (2003) for further details.

economic meltdowns including a dramatic real exchange rate depreciation, a spike in interest rates and spreads, and a severe recession accompanied by a huge decline in consumption growth. 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. Moreover, we investigate whether our model can generate reasonable estimates of the risk premium.

B. Prior Distribution of the Parameters

In this section we review the assumptions that underpin our prior distributions for the relevant parameters. However, not all parameters will be estimated. We calibrate three parameters: the discount rate, β, was set at 0.98; the share of capital in the production function, α, was calibrated to 0.37; and the adjustment cost parameter for the accumulation of foreign bonds, ϕ_{F^*}, was set at 0.01. 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 in the estimation process through the definition of the steady state. In addition, the adjustment cost parameter is usually calibrated to a small value since its main goal is to render a stationary steady state, overcoming the unit-root problem in open economy models.

Table 1 reports the type of density, its mean as well as the percentiles of our prior distributions. It is noteworthy to emphasize that our priors are fairly flat and cover a wide range of parameter values. With ν, we denote the elasticity of the risk premium to changes in the capital-to-net worth ratio k. Kawai, Hahm, and Iarossi (1999) 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 is about 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, that incorporates these bounds, as depicted in Table 1. Lacking a reliable benchmark in specifying our prior for ν we choose a beta distribution for this parameter with a mean of 0.2 and standard deviation of 0.1.[23] Taken together these prior densities imply a wide range of plausible values for the endogenous risk premium.[24]

The parameters for openness, $(1-\gamma)$, and the price resetting probability of the firms, κ, are by definition constrained to the unit interval. Therefore, we specify their prior densities as belonging to the beta family with means of 0.6 and standard deviations of 0.1.

For the remaining three structural parameters, the priors are specified as gamma distributions. For the intertemporal elasticity of substitution, σ, and the elasticity of labor supply, ψ, 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, λ, has a gamma prior with mean 8 and a standard deviation of 3. Note that this allows for a wide range of markups, ranging from 6

[23] Below—to assess robustness—we consider the case of a prior with values even closer to zero.

[24] If we consider the one standard deviation bands for both k and ν, these priors would imply that the risk premium is approximately between 5 and 28 percent.

to 59 percent.

The parameters describing monetary policy are based on a Taylor rule that allows for interest rate inertia, expanded to include responses to the real exchange rate, in addition to output and inflation. Consistent with prior beliefs that central banks smooth interest rate adjustments, the prior for ζ_i, is beta density with center at 0.8 and standard deviation 0.2. Priors for the coefficients on inflation, output gap, and real exchange rate parameters–ζ_π, ζ_Y, and ζ_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. Note that these priors encompass a wide range of responses of the monetary authority.

Finally, the priors for the variances of the exogenous stochastic processes correspond to an inverse Wishart distribution with a mean of 1 and standard deviation of 0.75. The five shocks to technology, foreign demand, foreign interest rate, labor disutility, and cost-push, are all $AR(1)$ processes. For their autoregressive coefficients we specify a Beta distribution with mean 0.5 and a standard deviation of 0.25, which yields a fairly flat density.

It is worth reiterating that, as opposed to other papers that use Bayesian techniques for estimating dynamic stochastic general equilibrium models, we choose prior distributions that are very flexible and permit a broad array of possible values.

C. Results

Table 2 presents the main results from the estimation of the model using Bayesian methods. The medians of the parameters and the 10th and 90th percentiles are provided, which quantify the uncertainty characterizing these estimates. Overall, the main structural parameters are fairly tightly estimated, despite our loose priors.

The estimate of the elasticity of the risk premium, ν, has a median value of 0.066, with the 10th and 90th percentiles of 0.031 and 0.11, respectively. This parameter could be interpreted as a summary statistic indicating how vulnerable the economy is to shocks affecting aggregate balance sheets. Notice from Figure 1 that the posterior is sharply peaked, relative to our prior distribution, suggesting that the data are informative about ν. As ν is estimated away from zero, this indicates that the financial accelerator and balance sheet vulnerabilities matter. Even though the literature does not provide a very solid justification for the calibration of this elasticity, it is encouraging that our estimates are in the range of values used previously by calibration-based studies of the financial accelerator, which typically lie between 0.01 and 0.02.[25]

The implied median risk premium we obtain is 3.27 percent per quarter, with the 10th and 90th percentiles corresponding to 1.93 and 4.87 percent, respectively. This implies an annualized endogenous risk premium of about 13 percent. Although the annualized risk premium is relatively high–most likely reflecting the impact of the East Asian crisis–using the 10th percentile implies

---

[25] See for example, Devereux, Lane, and Xu (2004).

an annual risk premium of 7.95 percent, which is much more in line with the data.[26] This not only reiterates the importance of potential balance sheet vulnerabilities, but also sheds light on the severity of the recession, especially in the context of a large contraction in investment spending.[27]

The parameters of the Taylor rule are estimated fairly well. The median value of the coefficient on inflation is 2.23, which suggests an aggressive response to inflationary pressures. The median response of the interest rate to the output gap is fairly small and equal to 0.02. Notice from Figure 1 that the posterior density for this parameter concentrates at the lower bound of zero, despite our flexible prior. There is a considerable degree of interest rate smoothing, which is very tightly estimated–judging from posterior percentiles. The median coefficient on the lagged interest rate is inferred to be 0.66, which suggests the authorities' preference to smooth interest rate fluctuations.

Our estimates also suggest that in formulating policy the monetary authorities also took into consideration movements in the exchange rate. The estimated coefficient is 0.05, with 10th and 90th percentiles corresponding to 0.02 and 0.09, respectively. This result suggests that the Korean government implemented a monetary regime that targeted the exchange rate. In this regard, our estimates could be interpreted as providing some evidence supporting the "fear of floating" hypothesis.[28] Relatedly, these estimates echo the results of Elekdag and Tchakarov (2004), who show that when the debt-to-net worth ratio exceeds a certain threshold, the welfare costs associated with a flexible exchange rate regime could exceed those of a peg.

The intratemporal elasticity of substitution between varieties of domestically produced goods, λ, takes a median value of 10.39, implying a price markup of approximately 11 percent, which is remarkably close to the value commonly used in calibrations.[29] As observed from Figure 1, the posterior density for λ is concentrated in a neighborhood of 10, despite our very flat prior.

The inverse of the intertemporal elasticity of substitution, σ, has a median value of 1.67, which is also well in line with values found in the literature. The inverse of the Frisch elasticity of labor supply, ψ, is 4.15. Contrary to other studies–including Smets and Wouters (2003)–who have reported difficulties in pining down this parameter, we find fairly tight posterior probability bands around this estimate, as seen in Figure 1. Also, our estimate is much closer to 3, the preferred value in calibration exercises and is much larger than estimates from open economy DSGE models in other papers for other countries (see Justiniano and Preston, 2004). A 90 percent posterior density interval for the Calvo price resetting probability, κ, lies between 0.48 and 0.53, with a median value of 0.51. This implies that the average duration of a contract is around two quarters,

---

[26] Relatedly, work is under way on incorporating stochastic volatility in the estimation of DSGE models.

[27] In fact, the correlation of real gross fixed capital formation growth and the EMBI spread for Korea was -0.37 during the sample spanning the fourth quarter of 1997 to the fourth quarter of 2000. Recall that the EMBI spread for Korea was not calculated before the fourth quarter of 1997.

[28] We defer however formally testing the hypothesis that $\zeta_S > 0$ for future work. This is the approach in Lubik and Shorfheide (2003) who investigate, using a small open economy model, whether or not central banks target exchange rates.

[29] See for instance the calibration in Chari, Kehoe, and McGrattan (2002).

reflecting a relatively higher average inflation rate in our sample period.

The estimate for the parameter representing the degree of openness $(1 - \gamma)$ is about 0.08. This estimate suggests a fairly low share of foreign goods in the the Korean economy. National account data reveal that the share of imports in GDP $(1 - \gamma)$ for the sample is equal to 0.34, reflecting a high degree of openness. Our low estimates could well be an artifact of our choice of specification for the foreign block. In staying close to the original model of Cespedes, Chang, and Velasco (2004), we have abstracted from providing an explicit characterization of the foreign economy and have limited ourselves to just positing an ad hoc functional form for the export demand function. Consequently, we hypothesize—and leave for future research—that if the foreign block were to be modeled in greater depth, this estimate would be more in line with the one obtained from national accounts.

Turning to the persistence of the exogenous disturbances, the estimates for the autoregressive coefficients are reasonable and imply substantial inertia in the shocks. The only exception is the export disturbance, which as seen in Figure 1, yields a fairly flat posterior density. This suggests, once again, that our choice of modeling for the foreign block may not allow to adequately capture the influence of foreign disturbances. The technology and the cost-push are the most persistent. It is noteworthy that the estimate for the cost-push shock is close the boundary of one. We view this result as indicating that the addition of price indexation mechanisms could capture endogenous inflation persistence more accurately and may be a worthy extension of the model that we intend to pursue in the near future.

Overall, we are able to obtain very reasonable and tight estimates of most parameters. More importantly, and to the heart of this paper, the model delivers plausible estimates of the endogenous risk premium that coincide with the actual values for this variable.

D. Robustness Analysis

We check the robustness of our main result—the structural estimate of the risk premium—by re-estimating the model with a different prior for the capital to net worth ratio, k, and the elasticity of the risk premium, ν. The prior for ν in the baseline estimates was centered at 0.2, and delivered a median estimate of 0.066. Even though the posterior percentiles are bounded away from zero, it is important to recognize that when ν it is exactly equal to zero, the financial accelerator mechanism seizes to exist. Entrepreneurs will then borrow from abroad in foreign currency, but the cost associated with this source of financing will be given by the foreign interest rate and will not be augmented by a risk premium.

To explore the behavior of the model when we allow a priori for values of this elasticity closer to zero, we consider an alternative prior specification that is centered at 0.07 and has a standard deviation of 0.06. This alternative prior ordinate has 1st, 5th, and 10th percentiles corresponding to 0.001, 0.005, and 0.01, meaning that we allow ν to take values in a tight neighborhood around zero. We keep all other priors unchanged and repeat the estimation procedure.

The results are presented in Table 3. The values for some of the structural parameters change somewhat relative to the baseline estimates, but the differences are minor. More importantly, the median estimate for ν is now 0.04, with a 90 percent posterior probability band in the range of 0.014 and 0.051. Once again, despite a prior allowing for very small values of v ,the posterior density is bounded away from zero**.** Crucially, the estimated level of the endogenous risk premium is around 11 percent. This value of the premium is very close to our previous estimate of 13 percent in the baseline model and represents an even more accurate description of the average risk premium evident in our sample. Together with the plots of prior and posterior discussed earlier, this observation confirms that the data is informative about ν and that our estimate of the endogenous risk premium is not an artifact of the choice of priors.

E. 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 fully fledged microfounded model can describe macroeconomic aggregates as well as, if not better then a BVAR, which has led to an increased interest in the role of models for policy making (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 (see Schorfheide, 2000 for details).

Model comparisons in a Bayesian setting are achieved by computing the posterior model probabilities, that, in the case of equal prior probabilities across models, reduce to the ratio of the model marginal likelihoods. 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 (1999), that can easily be done given the draws that are generated for the estimation of the model. Our estimate of the marginal density is presented in the top row of Table 4 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.

Closed form solutions for the marginal likelihood of a BVAR are available. Moreover, the marginal likelihood permits comparing the BVARs at various lag lengths, which we allow to vary between 1 and 4. Following Sims and Zha (2004) we use a symmetrized version of the Minnesota prior and include dummy priors for the own persistence of the data. The marginal densities obtained in this manner are displayed in Table 4 for the different lag lengths.

As expected, our model does not compare favorably to the preferred BVAR (1). This result is not surprising considering that our baseline model lacks any source of endogenous persistence such as habit formation or price indexation. The recent literature on DSGE modeling in closed economy contexts has emphasized the importance of such mechanisms in fitting the properties of the data.

Consequently, it is becoming increasingly standard to introduce habit formation and wage and price indexation in estimated DSGE models in order to impart further endogenous inertia to the model.[30] Therefore, it is reasonable to expect that the addition of some of these mechanisms will permit a more favorable comparison of this model relative to a BVAR.

IV. Concluding Remarks and Future Research

We use Bayesian techniques to estimate a small open economy model with the financial accelerator mechanism and the balance sheet channel. To our knowledge, this is the first paper that bridges the gap between theory and empirical research in the context of the financial accelerator with the presence of a balance sheet channel.

Our results can be summarized as follows. We find evidence of a sizable risk premium, which indicates that the financial accelerator mechanism is supported by the data. This highlights the latent balance sheet vulnerabilities that are believed to have exacerbated the Korean crisis. Also, it seems that the data do indeed support the argument that emerging market countries (EMCs) only have limited access to international capital markets.

In the context of Taylor rules, we find evidence for exchange rate smoothing. The explicit inclusion of the balance sheet channel and the financial accelerator could be interpreted as a possible rationale explaining why many EMCs seem to heavily manage their exchange rates, even when their respective governments have clearly announced a break from the regime of fixed exchange rates. Moreover, the coefficient on the lagged interest rate is significant, revealing the authorities' preference to smooth interest rate fluctuations.

It is worth mentioning that contrary to the burgeoning literature on the Bayesian estimation of DSGE models, the priors that we employ are fairly loose, thus allowing more freedom in estimating the parameters of interest. Despite this fact, the vast majority of the structural parameters are tightly estimated, which is satisfactory and attests to the fact that the data are quite informative.

Although the estimation of our simple model has yielded significant insights on the dynamics of the financial accelerator mechanism and potential balance sheet vulnerabilities, future research will include augmenting the model to include richer dynamics. We thus hope to refine our parameter estimates, especially with regard to the risk premium by incorporating endogenous persistence by including habit formation as well as price and wage indexation. We also plan to enrich the current framework by refining investment dynamics and including a more developed foreign block that could further refine our parameter estimates. Finally, we would like to use longer time series and consider estimating the model using data from other EMCs.

---

[30] Smets and Wouters (2003) as well as Juillard, Karam, Laxton, and Pesenti (2004) provide clear examples of how to enhance the endogenous persistence of the models when taking them to the data.

Table 1. Priors for the Baseline Model

Coefficient		Distribution	Mean	Percentiles 1%	99%
α	share of capital in production function	C	0.370	0.370	0.370
β	discount rate	C	0.980	0.980	0.980
ϕ_{F*}	adjustment cost for foreign bonds	C	0.010	0.010	0.010
γ	degree of closeness	B	0.6	0.361	0.814
κ	Calvo-pricing	B	0.6	0.361	0.814
λ	elasticity of substitution-domestic goods	G	8	2.692	16.572
ψ	inverse elasticity of labor supply	G	3	1.169	5.801
σ	inverse elasticity of substutution	G	3	1.169	5.801
ν	elsaticity of risk premium	B	0.2	0.033	0.478
k	capital-to-net worth-ratio	G	2	1.369	2.763
ζ_{π}	response on inflation, Taylor Rule	G	3	1.961	4.284
ζ_S	response on exchange rate, Taylor Rule	G	1	0.043	3.710
ζ**i**	interest rate smoothing, Taylor Rule	B	0.8	0.199	1.000
ζ_y	response on output, Taylor Rule	G	1.2	0.111	3.785
ϕ_A	AR(1), technology	B	0.5	0.033	0.967
ϕ_X	AR(1), foreign demand	B	0.5	0.033	0.967
ϕ_{i*}	AR(1), foreign interest rate	B	0.5	0.033	0.967
ϕ_{LD}	AR(1), labor disutility	B	0.5	0.033	0.967
ϕ_{CP}	AR(1), cost-push shock	B	0.5	0.033	0.967
ε_A	sd of technology shock	IG1	1	0.260	5.629
ε_X	sd of foreign demand shock	IG1	1	0.260	5.629
ε_{i*}	sd of foreign interest shock	IG1	1	0.260	5.629
ε_i	sd monetary policy shock	IG1	0.8	0.260	5.629
ε_{LD}	sd of labor disutility shock	IG1	1	0.260	5.629
ε_{CP}	sd of cost-push shock	IG1	1	0.260	5.629
ε_{NW}	sd of net worth shock	IG1	1	0.260	5.629

Distributions: G (Gamma), B (Beta), N (Normal), IG (Inverse Gamma). C means that the parameter has been calibrated. ϕ corresponds to the autoregressive coefficient of an AR(1) process. Last two columns report the inverse cumulative distribution function of each prior ordinate for the percentiles 0.01 and 0.99.

Table 2. Parameter Estimates for the Baseline Model

	Posterior Distribution		
	10%	**Median**	**90%**
α	0.370	0.370	0.370
β	0.980	0.980	0.980
ϕ_{F*}	0.010	0.010	0.010
γ	0.913	0.922	0.931
κ	0.487	0.505	0.523
λ	9.796	10.397	11.102
ψ	3.642	4.155	4.703
σ	1.579	1.678	1.773
ν	0.039	0.066	0.097
k	1.386	1.633	1.891
ζ_π	2.039	2.238	2.459
ζ_S	0.029	0.054	0.082
ζi	0.648	0.668	0.687
ζ_y	0.009	0.018	0.029
ϕ_A	0.823	0.849	0.866
ϕ_X	0.326	0.497	0.684
ϕ_{i*}	0.753	0.770	0.783
ϕ_{LD}	0.616	0.681	0.760
ϕ_{CP}	0.988	0.993	0.997
ε_A	0.967	1.056	1.158
ε_X	0.400	0.702	1.473
ε_{i*}	1.520	1.595	1.701
ε_i	0.479	0.513	0.552
ε_{LD}	2.866	3.170	3.469
ε_{CP}	1.220	1.431	1.698
ε_{NW}	0.461	0.644	0.880

Median and posterior deciles of the draws generated with a Random Walk Metropolis algorithm. Discarded the first 10,000 draws, retained the remaining 60,000 values.

Table 3. Parameter Estimates for the Alternative Model

	Posterior Distribution 10%	Median	90%
α	0.370	0.370	0.370
β	0.980	0.980	0.980
ϕ_{F^*}	0.010	0.010	0.010
γ	0.926	0.931	0.936
κ	0.480	0.495	0.510
λ	10.790	11.318	11.841
ψ	3.424	3.559	3.713
σ	1.635	1.850	2.039
ν	0.015	0.040	0.051
k	1.567	1.957	2.458
ζ_{π}	1.840	1.944	2.097
ζ_S	0.036	0.049	0.063
ζi	0.614	0.642	0.670
ζ_y	0.017	0.023	0.030
ϕ_A	0.805	0.826	0.837
ϕ_X	0.399	0.560	0.756
ϕ_{i^*}	0.795	0.808	0.819
ϕ_{LD}	0.664	0.703	0.743
ϕ_{CP}	0.992	0.995	0.997
ε_A	0.856	0.949	1.067
ε_X	0.358	0.568	0.882
ε_{i^*}	1.224	1.356	1.508
ε_i	0.465	0.505	0.546
ε_{LD}	2.891	3.063	3.187
ε_{CP}	1.228	1.364	1.496
ε_{NW}	0.385	0.678	1.224

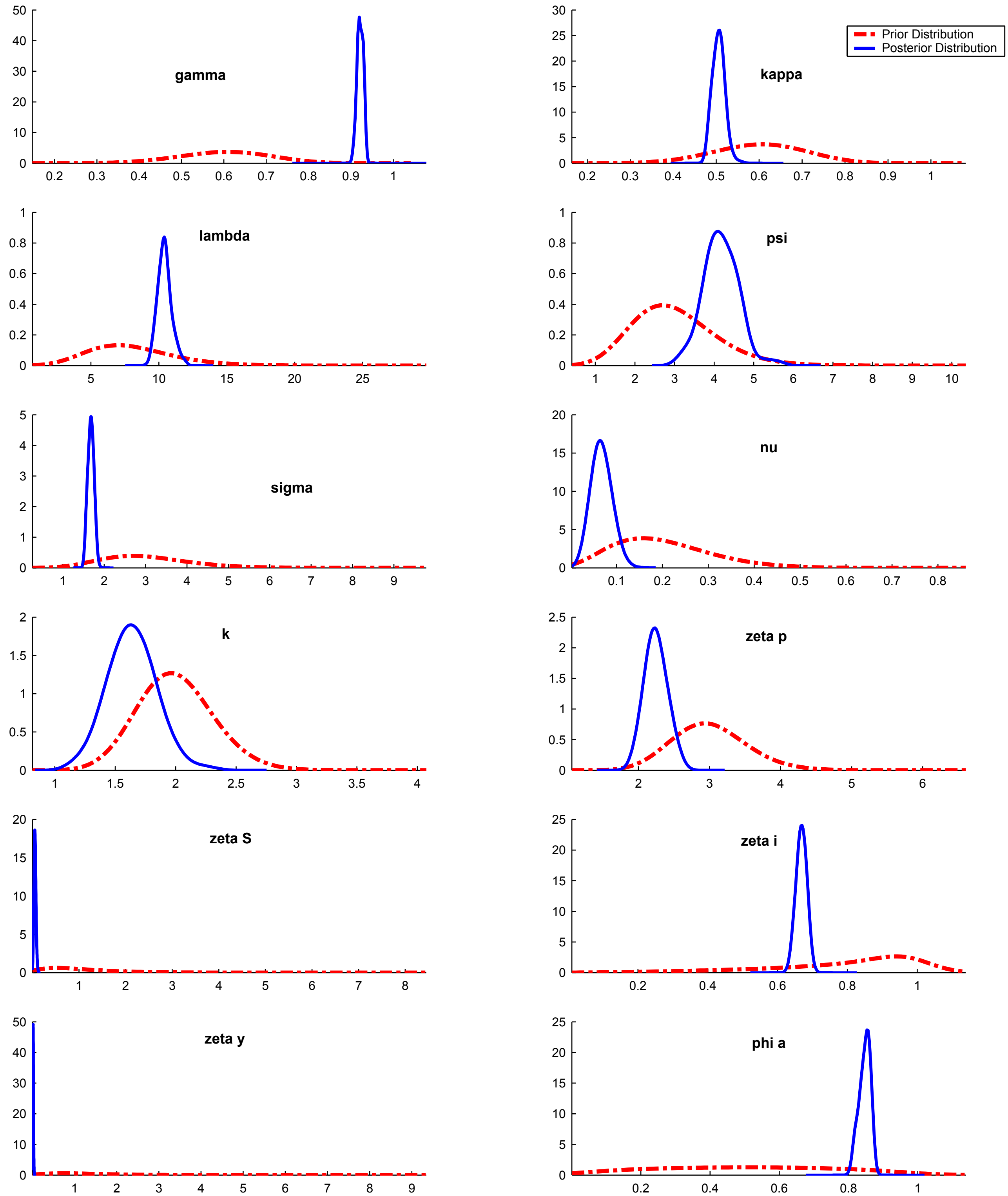
Median and posterior deciles of the draws generated with a Random Walk Metropolis algorithm. Discarded the first 10,000 draws, retained the remaining 60,000 values.

Table 4. Cross-validation

Modified Harmonic Mean	
Model	**Log marginal**
DSGE-FA Model	-680.974
BVAR(1)	-624.7053
BVAR(2)	-653.7563
BVAR(3)	-672.0307
BVAR(4)	-690.1469

Log Marginal Likelihood obtained with a Modified-Harmonic Mean estimator. Results are insensitive to the choice of cutoff point for the reciprocal importance sampling density and correspond above to the mean of a grid of values between 0.1 and 0.9. Results are also robust to whether the density point corresponds to mean, medians, or the (simulated) maximizing value of the posterior draws.

Figure 1. Prior and Posterior Distributions for Selected Parameters



REFERENCES

Aghion, P., P. Bacchetta, and A. Banerjee, 2001, "Currency Crises and Monetary Policy in an Economy with Credit Constraints," *European Economic Review,* Vol. 47, pp. 1121-50.

Bernanke, B., M. Gertler, and S. Gilchrist, 1999, "The Financial Accelerator in a Quantitative Business Cycle Framework," *Handbook of Macroeconomics* (Amsterdam: North-Holland).

Calvo, G., 1983, "Staggered Prices in a Utility-Maximizing Framework," *Journal of Monetary Economics*, Vol. 12, pp. 383-398.

———, 1999, "On Dollarization" (unpublished; College Park, Maryland: University of Maryland).

———, and C. Reinhart, 2001, "Fear of Floating" (unpublished; College Park, Maryland: University of Maryland).

Cespedes, L., R. Chang, and A. Velasco, 2004, "Balance Sheet and Exchange Rate Policy," *American Economic Review*, Vol. 94, No. 4, pp. 1183-93.

Chari, V. V., P. J. Kehoe, E. R. McGratten, 2002, "Can Sticky Price Models Generate Volatile and Persistent Real Exchange Rates?" *Review of Economic Studies*, Vol. 69, pp. 533-67.

Devereux, M., P. Lane, and J. Xu, 2004, "Exchange Rates and Monetary Policy in Emerging Market Economies," IIIS Discussion Paper No. 36 (Dublin, Ireland: Trinity College).

Elekdag, S., and I. Tchakarov, 2004, "Balance Sheets, Exchange Rate Policy, and Welfare," IMF Working Paper No. 63 (Washington: International Monetary Fund).

Eichengreen, B., and R. Hausman, 1999, "Exchange Rates and Financial Fragility," NBER Working Paper No. 7418 (Cambridge, Massachusetts: National Bureau of Economic Research).

Friedman, M., 1953, *Essays in Positive Economics*, (Chicago: University of Chicago Press).

Gelman, A., J. Carlin, H. Stern, and D. Rubin, 1995, *Bayesian Data Analysis*, (New York: Chapman and Hall).

Gertler, M., S. Gilchrist, and F. Natalucci, 2003, "External Constraints on Monetary Policy and the Financial Accelerator," BIS Working Paper No. 139 (Basel: Bank for International Settlements).

Geweke, J., 1999, "Using Simulation Methods for Bayesian Econometric Models Inference, Development and Communication," Staff Report 249 (Minneapolis: Federal Reserve Bank).

Julliard, M., P. Karam, D. Laxton, and P. Pesenti, 2004, "Welfare-Based Monetary Policy Rules in an Estimated DSGE Model of the U.S. Economy"(unpublished; Washington: International Monetary Fund).

Justiniano, A., and B. Preston, 2004, "New Open Economy and Exchange Rate Pass Through: an Empirical Investigation" (unpublished; New York: Columbia University).

Kawai, M., H. J. Hahm, and G. Iarossi, 1999, "Corporate Foreign Debt in East Asia: Too Much or Too Little?" (unpublished; Washington: World Bank).

Kollmann, R., 2002, "Monetary Policy in the Open Economy: Effects on Welfare and Business Cycles," *Journal of Monetary Economics*, Vol. 49, pp. 989–1015.

Krugman, P., 1999, "Balance Sheets, the Transfer Problem, and Financial Crises," *International Tax and Public Finance*, Vol.6, pp. 459–72.

Linaa, J. G., and J. Rand, 2004, "Testing the Financial Accelerator in a Small Open Economy. Evidence from Scandinavia" (unpublished; Copenhagen, Denmark: University of Copenhagen).

Lubik, T. and F. Schorfheide, 2003, "Do Central Banks Respond to Exchange Rates? A Structural Investigation" The Johns Hopkins University Department of Economics Working Paper No. 505 (Baltimore: The Johns Hopkins University).

Meier, A., and G. Mueller, 2004, "Fleshing Out the Monetary Transmission Mechanism: Output Composition and the Role of Financial Frictions" (unpublished; Florence, Italy: European University Institute).

Schorfheide, F., 2000, "Loss Function-Based Evaluation of DSGE Models," *Journal of Applied Econometrics*, Vol. 15, No. 6, pp. 645-70.

———, 2003, "Learning and Monetary Policy Shifts" (unpublished; Philadelphia: University of Pennsylvania).

Sims, C., and T. Zha, 2004, "Were There Regime Switches in U.S. Monetary Policy?," Federal Reserve Bank of Atlanta Working Paper No. 2004-14 (Atlanta: Federal Reserve Bank of Atlanta).

Sims, C., 2003a, "Probability Models for Monetary Policy Decisions" (unpublished; Princeton: Princeton University).

———, 2003b, "Solving Linear Rational Expectations Models," *Computational Economics*, Vol. 20, No. 1, pp. 1-20.

Smets, F., and R. Wouters, 2003, "An Estimated Stochastic Dynamic General Equilibrium Model of the Euro Area," *Journal of the European Economic Association*, Vol. 1, pp. 1123-75.

Tovar, C., 2004, "Devaluations, Output, and the Balance Sheet Effect: A Structural Econometric Analysis," forthcoming, BIS Working Paper.

Yun, T., 1996, "Nominal Price Rigidity, Money Supply Endogeneity, and Business Cycles," *Journal of Monetary Economics*, Vol. 37, pp. 345-70.