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Investment During the Korean Financial Crisis: A Structural Econometric Approach[1]

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Abstract

Without capital market imperfections, the capital structure of a firm, including the size, the maturity and the currency composition of debts, should not matter for investment decisions. The Asian financial crises provide a good opportunity to test this hypothesis. We approach the problem in two ways: First, we apply a conventional reduced-form analysis to a panel data of Korean manufacturing firms, argueing that the devaluation that occurred during the crisis provides a natural experiment in which to assess the effect of balance sheet shocks to investment. Second, we use indirect inference to estimate a structural dynamic programming problem of a firm with foreign debts and financial constraints. Both reduced-form evidence and structural parameter estimates imply an important role for finance in investment at the firm level. Counterfactual simulations imply that balance sheet effects may account for 50% to 80% of the drop in investment during the crisis period. Although our estimates suggest that foreign-denominated debt had relatively little effect on aggregate investment spending for the Korean economy during this crisis episode, counterfactual experiments imply sizeable contractions in investment through this mechanism for economies that are more heavily dependent on foreign-denominated debt.

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1 Introduction

Without capital market imperfections, the capital structure of a firm, including the size, the maturity, and the currency composition of debts, should not matter for investment decisions. The Asian financial crises provide a good opportunity to test this hypothesis, i.e., the irrelevance of finance in investment decisions. The devaluations that occurred during these crises abruptly and massively altered the debt burdens of firms with foreign-denominated debts. Because devaluations are exogenous events, at least from the perspectives of individual firms, such episodes make it easier to identify a distinct role for financial factors in investment decisions during financial crises.

This paper tests for the existence of a finance channel in the propagation of the Korean financial crisis. It also provides a quantitative assessment of the effect of foreign-denominated debt on investment. This analysis provides a useful perspective on the likely benefits to fixed versus flexible exchange rates during a financial crisis. A primary argument for maintaining a fixed exchange rate is that a devaluation may adversely affect balance sheets owing to the presence of foreign-denominated debt.[2] Our results imply that foreign-denominated debt plays an important role in explaining heterogenous outcomes across firms during the crisis period. The presence of foreign-denominated debt explains only a small fraction of the aggregate investment decline that occurred during the crisis period however. Although foreign-denominated debt is not an important factor for Korean firms during this time period, the financial crisis does appear to work through the balance sheet. In particular, high interest rates combined with low profits weakened firm balance sheets and exacerbated the crisis. Overall, financial frictions can account for 50 to 80% of the observed drop in investment during this episode.

Theoretically, a devaluation can affect investment through two distinct channels. First, the devaluation increases competitiveness and raises the marginal profitability of capital for firms that export. This increase in the marginal profitability of capital stimulates the investment of export-oriented firms.[3] Second, the devaluation influences the debt burden of firms – the value of debt relative to a firm's ability to repay the debt. In the presence of financial market imperfections, an increase in the debt burden causes a deterioration of the balance sheet and an increase in the cost of external finance. As external finance becomes more costly, firms reduce their investment.

Understanding the effect of foreign-denominated debt for investment spending requires

[2] Frankel (2003) summarizes the relative benefits of fixed versus flexible exchange rates. Aghion et al (2001), Cepedes et al (2002) and Gertler et al. (2007) analyze these issues using small open-economy models.

[3] The devaluation may also affect investment by changing the price index of investment goods. The increase in the domestic price of foreign investment goods was offset by a decline in the price of domestic investment goods so that the relative price of investment goods stayed relatively constant throughout this period.

firm-level data. We use a newly available panel-data set of Korean manufacturing firms to assess the strength of the finance channel discussed above. Importantly, the data set provides detailed information on the foreign exchange-rate exposure of the firm, both in terms of the amount of exports and in terms of the amount of foreign-denominated debt.

We begin with a reduced-form regression analysis. The exchange-rate crisis and ensuing devaluation provide a natural experiment with which one can measure the combined effect of the devaluation on firm-level investment spending. A key point to this identification strategy is that firms should respond differently to the devaluation depending on both the level of foreign sales and the amount of foreign-denominated debt. Following the devaluation, firms with high levels of foreign sales should increase their investment relative to other firms, while firms with high levels of foreign-denominated debt should decrease their investment relative to other firms. Controlling for exports allows us to cleanly identify the effect of foreign-denominated debt on investment spending.

The second part of the paper adopts a structural approach. It specifies a dynamic optimization problem of a firm which produces for both domestic and foreign markets and holds both domestic and foreign-denominated debt. The firm operates under a set of financial and non-financial constraints. The dynamic program is used to estimate the structural relationship characterizing investment, profitability and financial conditions. We then conduct counterfactual simulations to understand the role that foreign-denominated debt and financial frictions played during the crisis period.

Several recent papers estimate the effect of foreign-denominated debt on firm-level investment during currency devaluations. Using a sample of Latin American firms over the 1990's, Bleakley and Cowan (2002) find that the net effect of the devaluation was likely positive for firms with high levels of foreign-denominated debt. Because these authors do not have separate information on the export status of firms, they are unable to distinguish balance-sheet effects from competitiveness effects however. Aguiar (2004) examines the investment behavior of Mexican firms during the 1994 pesos devaluation.[4] By controlling for export status, this study finds a negative effect of foreign-denominated debt that is distinct from the competitiveness effect. Similarly, Desai, Foley, and Forbes (2007) document a negative response of firm-level investment to devaluations for local but not multi-national firms. These papers adopt a reduced-form approach and therefore cannot formally quantify the effect that foreign-denominated debt exerts on investment.

Our paper is also related to the extensive literature on firm-level investment and capital

[4] Pratap and Urrutia (2003) calibrate a model to fit this episode.

market imperfections.[5] Much of this literature focusses on the role of cash flow for investment spending. Although this literature finds strong evidence in favor of capital market imperfections (e.g. Fazzari, Peterson and Hubbard (1988), Hoshi, Kashyap and Scharfstein (1991)), these findings have been criticized for not adequately controlling for the possibility that cash flow is simply a proxy for investment opportunities or misinterpreting the relationship between investment, Q, and cash flow (Gilchrist and Himmelberg (1994), Kaplan and Zingales (1997), Gomes (1999), Abel and Eberly (2003), Eberly et al. (2006)).

A key question in this literature is how to identify the effect of balance sheet shocks that are independent of investment opportunities. Blanchard et al.(1994) and Lamont (1997) adopt a natural-experiment approach by examining the effect of shocks to cash flow that are arguably exogenous to the firm or firm segment's investment opportunities. More recent papers achieve identification by solving and estimating the dynamic program of a firm under capital market imperfections (Cooper and Ejarque (2003), Pratap and Rendon (2003), Bayraktar et al. (2005), Hennessy and Whited (2006)).

A major limitation of current structural estimates is the focus on a single shock that is perfectly correlated with profit opportunities. In such environments, one cannot separately identify the balance-sheet effect from the fundamentals effect absent strong assumptions regarding technology or market structure. In addition to focussing on a single shock environment, these models frequently abstract from adjustment costs, so that absent capital market imperfections, capital accumulation is frictionless. Because capital market imperfections limit investment spending, these estimation procedures may not be robust to the alternative hypothesis that investment responds to profits owing to sluggish adjustment on the real side. By combining real frictions with financial frictions, and identification through balance-sheet shocks we avoid such potential pitfalls.

From a methodological perspective our paper is the first structural estimation paper in the investment literature that models both adjustment costs and financial market imperfections within an economic environment that allows for firm-level heterogeneity and macroeconomic shocks as state variables. Thus, a contribution of our work is to provide a structural econometric framework that allows for meaningful heterogeneity in firm behavior in the presence of a large number of state variables. Our econometric procedure demonstrates the feasibility of estimating relatively complex models of firm behavior using an indirect-inference procedure that matches regression equations estimated with standard panel-data techniques. Our results imply that structural models can successfully replicate reduced-form regression equations obtained in the investment literature and that these reduced-form equations imply reasonable

[5] Hubbard (1998) and Stein (2003) provide recent surveys of this literature

structural parameters estimates for both adjustment costs and financial costs.

The organization of the remainder of the paper is as follows. Section 2 provides summary measures of our data. Section 3 formulates the decision problem of the firm and characterizes the efficiency conditions. Section 4 explains our reduced-form empirical strategy and reports the estimation results. Section 5 estimates the structural parameters using indirect inference; Section 5 also derives the impulse response functions of heterogenous firms and evaluates the role that financial factors and foreign-denominated debt played in the propagation of the crisis.

2 Overview of Korean Financial Crisis

This section provides an overview of the investment behavior of Korean firms during the financial crisis of 1997-1998. Figure 1 shows the impact of the crisis on our sample of manufacturing firms.[6] It plots the average ratios of investment, sales, and debt relative to total assets. For comparison purposes, it also plots the annual average real exchange rate. All variables are in logs and are normalized relative to their pre-crisis (1996) values.

The results in Figure 1 are consistent with the macroeconomic effects described elsewhere (Gertler, Gilchrist and Natalucci (2003), Kruger and Yoo (2001)). Between the onset of the crisis in 1996 and the trough of economic activity that occurred during 1998, sales fell 20% while investment fell nearly 100%.

Figure 1 also plots the debt-to-asset ratio for our sample of firms. Debt is valued in local currency and includes both the local-currency denominated debt and the foreign-currency denominated debt. The 70% depreciation of the currency implies a sharp rise in the value of foreign-denominated debt. As a result, the debt-to-asset ratio increases by 20% at the onset of the crisis, reflecting the stress on balance sheets caused by the currency depreciation.

To investigate how the investment rate differed in response based on the degree of a firm's foreign exchange-rate exposure, we divide firms by export status, and by the degree of foreign-denominated debt. Firms are classified according to export status using the pre-crisis average export to total sales ratio for each firm in our sample. We categorize firms as high-export firms if this ratio is above the pre-crisis median value. Similarly, we classify firms as high foreign-denominated debt firms based on the pre-crisis average foreign-denominated debt to total debt ratio, again using the pre-crisis median value as our cutoff. The average investment rates for high versus low foreign-denominated debt and high versus low export

[6] We defer our data description until section 4.

Figure 1: Investment, Sales and Debt during Financial Crisis.

firms are plotted in the upper two panels of figure 2.

Following the financial crisis, firms with high levels of foreign-denominated debt have low rates of investment relative to firms with low levels of foreign-denominated debt. There is little difference in the investment rate of firms with high levels of exports relative to firms with low levels of exports. Because there is a positive correlation between foreign-denominated debt exposure and foreign-sales exposure, high export firms tend to have higher foreign debt ratios which offset the beneficial effects of the exchange rate depreciation.

By considering low versus high export firms separately, the lower panels of Figure 2 help isolate the role of foreign-denominated debt on investment. For both high-export and low-export firms, foreign-denominated debt appears to depress the investment rate. The effect of foreign-denominated debt on investment is most severe for firms with the greatest mismatch between foreign sales and foreign-denominated debt exposures. Thus, the investment spending of the firms with high levels of foreign-denominated debt but little export revenue to offset the negative consequences of the devaluation appear to be the most vulnerable during the financial crisis.

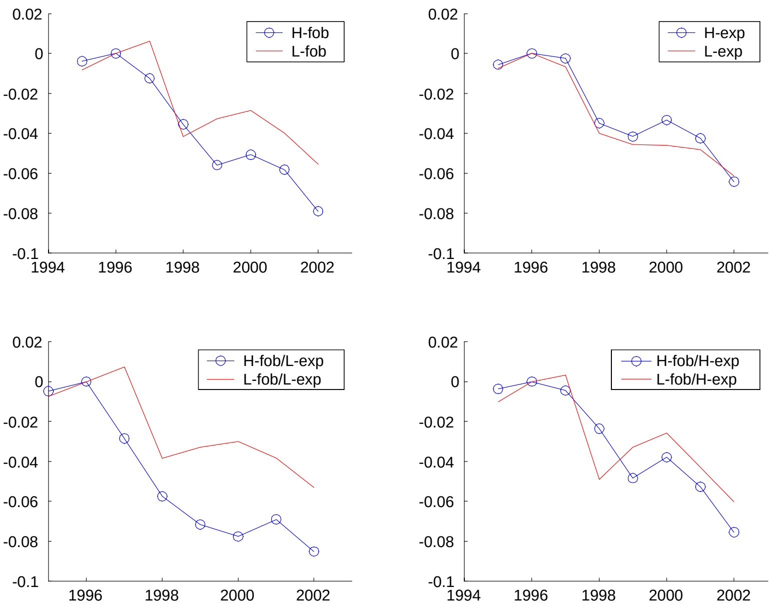

Figure 2: Investment Rates

3 The Investment Model

This section presents the structural model of investment that is estimated. The model is a standard convex-adjustment cost model of investment augmented to include financial market imperfections. The model explicitly incorporates the effect of exchange rates on investment working through the two distinct channels outlined above: the effect of exchange rates on fundamentals, and the effect of exchange rates on the firm's balance sheet.

We consider the dynamic programming problem for a firm which chooses capital, k', and foreign and domestic debt, b'_f and b'_d, to maximize the present value of dividends d subject to constraints on technology and a nonnegativity constraint on dividends. The recursive formulation of the problem is given by

$$v(k, b_d, b_f, \mathbf{z}, \mathbf{z}_{-1}) = \max_{k', b'_d, b'_f, d} \left\{ (1+\lambda)d + \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) v(k', b'_d, b'_f, \mathbf{z}', \mathbf{z}) Q(\mathbf{z}, d\mathbf{z}') \right\} \qquad (1)$$

where λ is the Lagrangian multiplier associated with the nonnegativity constraint, $\Lambda(\mathbf{z}', \mathbf{z})$ is the stochastic discount factor which the firm takes as given and μ is an exogenous survival rate introduced for technical reasons. The vector $\mathbf{z}$ contains all relevant exogenous state variables in the model, i.e., $\mathbf{z} = [r_d, e, e_{-1}, z, Z]$ where z is an idiosyncratic shock to the production function, Z is an aggregate shock which shifts the market demand, r_d is the domestic risk free rate, and e is the real exchange rate. The lagged real exchange rate is a state variable because it helps to predict the evolution of the domestic interest rate under uncovered interest parity (UIP). This point will be made clear in Section 5. $Q(\mathbf{z}, d\mathbf{z}')$ is a transition function of the vector $\mathbf{z}$. The presence of the lagged exogenous state variables $\mathbf{z}_{-1}$ as arguments of the value function is due to the dependence of the agency cost on the lagged variables.

The dividend of the firm is defined as the sum of profits net of investment costs plus net debt issuance:

$$d = \left(\sum_{i=d,f} \phi_i (\mathbf{z}) k^{\gamma_i} - \sigma \right) - i - c(i,k) - R_d b_d - eR_f b_f + b_d' + eb_f' \tag{2}$$

The profit function is the sum of domestic and foreign profits minus fixed costs to production σ. Here γ_d and γ_f denote the elasticity of profit with respect to capital in domestic and foreign markets, dictated by the degree of market powers in each market. Profits in each market depends on the exogenous profitability indices $\phi_i(\mathbf{z})$ which in turn depend on the exogenous shocks to the firm.[7] The price of investment goods is normalized to unity. Adjustment costs are convex and constant-returns-to-scale, i.e., $c_1 > 0$, $c_{11} > 0$ and $c(\alpha i, \alpha k) = \alpha c(i,k)$; capital accumulates subject to the exponential depreciation rate δ.

Domestic debt is measured in local currency units and foreign-denominated debt is measured in foreign currency units. R_d and R_f denote the gross interest rate on domestic and foreign bonds respectively where R_f is also measured in foreign currency units. The vector z denotes the set of all relevant exogenous state variables.

To introduce financial frictions, we impose a zero dividend constraint on the dynamic programming problem, i.e.,

$$d \geq 0. \tag{3}$$

We also assume that the total borrowing cost can be decomposed into a risk-free interest rate and an external-finance premium,

$$R_f = (1 + r_f)(1 + \eta) \tag{4}$$

[7] Details regarding market structure, production technology, and the stochastic discount factor are discussed in section 5.

and

$$R_d = (1 + r_d)(1 + \eta) \tag{5}$$

where r_f and r_d denote the risk free rate on foreign and domestic bonds, respectively and η denotes the common external-finance premium. Following Gilchrist and Himmelberg (1998) and Gomes(1999), this formulation assumes that financial constraints are summarized in a single reduced-form external finance premium, η, combined with a dividend constraint that limits new equity issuance.

The assumption of a common external-finance premium on domestic and foreign-denominated debt allows us to simplify the model and eliminate one endogenous state variable.[8] Let

$$\omega \equiv \frac{e_{-1}b_f}{b_d + e_{-1}b_f}$$

denote the ratio of foreign-denominated debt to total debt in local currency units. Uncovered interest parity implies that the firm cares about the total debt obligation $b = b_d + eb_f$ but is indifferent exante between the currency composition of debt. Because the firm is exante indifferent to the currency composition of debt, the foreign-denominated debt ratio may be taken as a fixed parameter for each firm rather than a choice variable.[9] The programming problem given by equations 1-4 is then equivalent to the following program with smaller dimension

$$v(k, b, \mathbf{z}, \mathbf{z}_{-1};\omega) = \max_{k',b'_d,b'_f,d} \left\{ (1 + \lambda)d + \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z})v(k', b', \mathbf{z}', \mathbf{z};\omega)Q(\mathbf{z}, d\mathbf{z}') \right\} \tag{6}$$

[8] By allowing for both adjustment costs and financial frictions, our model has two endogenous state variables combined with three exogenous state variables that determine the macroeconomic environment. Adding a third endogenous state variable increases substantially the complexity of the numerical procedure used to solve the value function, making structural estimation infeasible.

[9] In a more general model that allows for expost default and exit, the non-linearity in the firm's payoff structure created by default generates a hedging demand even when both the lender and debtor are risk neutral (see Albuquerque (2006)). In our formulation, firms and lenders are risk neutral. We also do not explicitly model expost default and exit owing to the complexity of the contracting framework implied by a model that allows for both adjustment costs and financial frictions. Under these assumptions, the appendix shows that the UIP no-arbitrage condition implies that firms are indifferent exante between foreign and domestic debt so that fixing the currency composition does not reduce the expected firm value. Fixing the debt ratio is analytically convenient but not necessary for our results since what matters to the firm is the effect of the unanticipated devaluation on the balance sheet conditional on the existing debt structure. In addition, a stable foreign debt ratio is empirically justified: the firm-level correlation between ω_t and ω_{t-1} is greater than 0.9 in annual data. We also find no evidence to suggest that in the year prior to the crisis, firms changed the currency composition of their debt owing to increased anticipation of the devaluation.

where the dividend is redefined as

$$d = \left(\sum_{i=d,f} \phi_i \left(\mathbf{z} \right) k^{\gamma_i} - \sigma \right) - i - c(i,k) - \left(\frac{e}{e_{-1}} \right) R_f \omega b - R_d \left(1 - \omega \right) b + b' \tag{7}$$

Although the currency composition of debt is assumed to be fixed over time for each firm, our empirical work allows it to vary cross-sectionally in a manner consistent with the empirical relationship between the currency composition of debt and other key features of the firm such as the export-to-sales ratio and leverage ratio. Thus while we recognize that at a deeper level, hedging motives combined with market access are important determinants of the currency composition of debt, we view the effect of such motives on the investment policy during the crisis period as second order relative to the direct effect of the large (seventy-percent) exchange rate devaluation on the balance sheet.

By imposing the UIP condition $(1 + r_d) = (1 + r_f)E(e/e_{-1}|\mathbf{z}_{-1})$, the dividend can be simplified to

$$d = \left(\sum_{i=d,f} \phi_i \left(\mathbf{z} \right) k^{\gamma_i} - \sigma \right) - i - c(i,k) - \Omega \left(e, e_{-1}; \omega \right) R_d b + b' \tag{8}$$

where

$$\Omega \left(e, e_{-1}; \omega \right) \equiv \left[\omega \frac{e/e_{-1}}{E(e/e_{-1}|\mathbf{z}_{-1})} + (1 - \omega) \right] \tag{9}$$

and $\frac{e/e_{-1}}{E(e/e_{-1}|\mathbf{z}_{-1})}$ denotes the surprise to the exchange rate. The term $\Omega \left(e, e_{-1}; \omega \right)$ is a pricing function which translates the current value of debt outstanding into local currency units conditional on the currency composition ω. An unanticipated devaluation causes an unanticipated increases in the local currency value of debt outstanding in direct proportion to the share of foreign-denominated debt. Thus, if $\omega = 0$, the exchange rate devaluation has no impact on current debt obligations, whereas if $\omega = 1$, the exchange rate devaluation causes a one-for-one increase in the value of current debt outstanding.[10]

The external-finance premium is parsimoniously specified as

$$\eta(x) \equiv \kappa \left[\exp\left(x \right) - 1 \right] \tag{10}$$

[10] Only unanticipated movements in the exchange rate causes changes in the value of debt outstanding. The UIP conditions guarantee that anticipated effects are already built into the relevant risk free rate.

where

$$x \equiv x(k, b, \mathbf{z}_{-1})$$

and κ is a parameter that governs the strength of the financial friction. The $x(\cdot)$ function is chosen to satisfy the following properties: i) an increase in capital reduces the external finance premium, ii) an increase in debt increases the premium, iii) any exogenous state variable that predicts an increase in profitability in the next period also reduces the external finance premium. We consider the following functional form for $x(\cdot)$:

$$x(k, b, \mathbf{z}_{-1}) = \frac{b}{\sum_{i=d,f} \phi_i(\mathbf{z})\, k} \tag{11}$$

Normalizing the leverage ratio by the profit factor $\sum_{i=d,f} \phi_i(\mathbf{z})$ insures that a firm whose profits and debt are fully denominated in foreign currency does not experience a balance sheet shock owing to an exchange rate movement.[11] Under this specification, the external finance premium is a function of the state variables when the debt instrument issued, i.e.,

$$\eta' = \eta(k', b', \mathbf{z}).$$

Because the function η is a strictly convex function of x, the slope of the premium rises more rapidly as leverage increases.

Given the premium on external funds, firms may have an incentive to accumulate savings and grow their way out of the financial constraint. To rule out this possibility, we introduce a survival probability, μ. In steady-state, if $\mu = 1$ and Λ is determined by the steady state risk free rate, optimal leverage is indeterminate and the Modigliani-Miller theorem applies. If $\mu < 1$, the survival probability works as an additional discount factor and the firm holds a positive amount of debt in the steady state. The appendix shows how the fixed cost parameter and survival probability can be determined from the data.

[11] We have also considered the specification $x = b / \left(\sum_{i=d,f} \phi_i(\mathbf{z})\, k^{\gamma_i} + (1-\delta)k\right)$. The denominator of this expression is profits plus undepreciated capital. Without adjustment costs, this variable represents net cash available inside the firm. In this case, the premium function is simply an approximation to the cost of external finance obtained in a one period debt contracting framework. With adjustment costs, the liquidation value of the firm is more complex however, and there is no longer such a one-to-one mapping. Our formulation preserves the intuition that an optimal debt contract implies a relationship between the cost of external finance and leverage, while recognizing that, with adjustment costs, the value of capital in place depends on the profitability of the firm.

The aggregate state variables $[e, r_d, Z]$ follow Markov processes which are specified to match the macroeconomic environment during the crisis period. These processes allow for interdependence between the domestic real interest rate and the exchange rate by assuming that

$$r_d^{'} = f(r_d, e, e_{-1}) + \varepsilon'$$

where ε' may be interpreted as a shock to the country-risk premium. The form of $f()$ is determined by imposing uncovered interest parity, further details are provided in section 5.

Because the idiosyncratic shock to profitability, z, is assumed to be an iid random variable[12], the information contained in the vector $[b, \eta(k, b, \mathbf{z}_{-1}), \mathbf{z}]$ can be summarized in a single state variable, net worth, i.e.,

$$n \equiv \left(\sum_{i=d,f} \phi_i(\mathbf{z}) k^{\gamma_i} - \sigma \right) + (1-\delta)k - \Omega(e, e_{-1}; \omega) R_d(r_d, \eta) b \tag{12}$$

so that the value function may be defined as

$$\begin{aligned} v(k, n, \mathbf{z}; \omega) \quad &= \quad \max_{k', b', d} \left\{ (1+\lambda)\, d + \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) v(k', n', \mathbf{z}'; \omega) Q(\mathbf{z}, d\mathbf{z}') \right\} \\ & \quad\quad \text{s.t. } 12 \end{aligned}$$

In the above formulation, the vector, $\mathbf{z}$ depends only on the vector of aggregate state variables.

The asset pricing formula implied by the efficiency condition for b' is given by

$$\frac{1}{\mu} = \int_{\mathbf{z}'} \left(\frac{1+\lambda'}{1+\lambda} \right) \Lambda(\mathbf{z}', \mathbf{z}) \left[\Omega(e', e; \omega)\,(1 + r_d') \left(1 + \eta' + \frac{\partial \eta'}{\partial b'} b' \right) \right] Q(\mathbf{z}, d\mathbf{z}') \tag{13}$$

Owing to the survival probability μ, the marginal benefit from issuing new debt is greater than one when evaluated at $\Lambda(\mathbf{z}', \mathbf{z})$, the market's discount rate.

Similarly, the efficiency condition for k' implies the asset pricing formula

$$1 + \frac{\partial c}{\partial i}(i, k) = \mu \int_{\mathbf{z}'} \left(\frac{1+\lambda'}{1+\lambda} \right) \Lambda(\mathbf{z}', \mathbf{z}) \left[\frac{\partial d'}{\partial k'} + (1-\delta) \left(1 + \frac{\partial c}{\partial i'}(i', k') \right) \right] Q(\mathbf{z}, d\mathbf{z}') \tag{14}$$

[12] The iid assumption avoids adding an additional state variable. The estimated autocorrelation of the idiosyncratic shock to profits is on the order of 0.2~0.3, which is sufficiently low to justify this assumption.

where the effect of capital on next period's dividend is given by

$$\frac{\partial d'}{\partial k'} = \sum_{i=d,f} \gamma_i \phi_i(\mathbf{z}) k'^{\gamma_i - 1} - \frac{\partial c}{\partial k'}(i', k') - \Omega(e', e; \omega)(1 + r'_d) \frac{\partial \eta'}{\partial k'} b'.$$

Owing to the presence of financial market imperfections, the ratio of Lagrange multipliers $(1 + \lambda') / (1 + \lambda)$ acts like an additional discount factor that influences the firm's optimal choices of debt and investment.

The model cannot be solved analytically, we therefore use Chebyshev projection methods (Judd(1992)) to obtain a numerical approximation. Owing to the presence of occasionally binding constraints, the solution approximates the conditional expectations of the model first and then reconstructs the policy and the multiplier variables using the approximated conditional expectations following Wright and Williams(1982), den Haan and Marcet(1990) and Christiano and Fisher(2000). Relative to Christiano and Fisher (2000), the solution method is complicated by the fact that the model has two endogenous state variables, debt and capital. Details of this method are described in the appendix.

To understand the basic mechanism at work in the model, consider the effect on an unanticipated devaluation. The devaluation causes an increase in the current debt obligation and a reduction in net worth. The increased debt obligation raises the shadow value of internal funds. The firm may respond by either reducing the dividend, increasing debt or cutting back on investment. The firm decides on how much external finance to raise by equating the shadow value of internal funds with the marginal cost of debt according to the efficiency condition for new debt issuance. Simultaneously, the firm chooses its investment policy to equate the cost of investment today relative to the benefit tomorrow where tomorrow's benefit is evaluated at the firm's internal shadow value of funds. As a result, the unanticipated devaluation causes an increase in the premium on external funds, a reduction in new debt issuance and a fall in investment.

4 Regression Analysis.

This section formally assesses the role of foreign-denominated debt on investment spending using a panel-data regression framework. The regressions reported in this section serve two purposes: i) to assess the effect of balance-sheet shocks on investment using a reduced-form regression analysis and ii) to provide an empirical regression that can be used to estimate the structural model parameters using indirect inference.

4.1 An Empirical Investment Equation

The empirical investment equation requires measures of investment fundamentals and the balance sheet. Fundamentals are proxied by the firm's sales-to-capital ratio. This is consistent with the assumption that firms face monopolistic competition and that the production function is Cobb-Douglas in factor inputs. If producers have market power owing to monopolistic competition, firms may set different markups in the domestic market relative to the foreign market. As shown in the appendix, the marginal profitability of capital can then be decomposed into a weighted average of the domestic sales-to-capital ratio and the exports-to-capital ratio, where the relative weights depend on the degree of market power in each market. The regression analysis includes both of these variables separately. This effectively allows the response of investment to fundamentals to differ based on the source of profitability (foreign versus domestic).

To measure the effect of the exchange rate through the balance sheet we follow our model and construct a proxy for $\Omega_{jt}b_{jt}/a_{jt}$. When constructing this proxy, we are careful to use only exante information however. Let b_{jt} denote the total debt of the firm at the beginning of the period, denominated in local currency terms. Let a_{jt} denote a measure of the beginning-of-period value of total assets (again denominated in local currency terms). The ratio of debt to assets b_{jt}/a_{jt} provides a measure of the balance sheet of the firm.[13]

To construct our measure of Ω_{jt} we first measure the pre-crisis(1994∼1996) sample mean of each firm's foreign debt ratio, i.e.,

$$\hat{\omega}_j = 1/T_j^{pc} \sum \left(b_{f,j,t}/b_{j,t}\right)$$

where $b_{f,j,t}$ is the real foreign debt in domestic currency units and T_j^{pc} is the number of nonmissing observations of firm j, during the pre-crisis period. Given $\hat{\omega}_j$, the effect of an exchange rate movement on the value of debt can be measured as

$$\hat{\Omega}_{jt} = 1 - \hat{\omega}_j + \hat{\omega}_j \left(e_t/e_{t-1}\right) \tag{15}$$

[13] We use assets rather than capital in place since the former controls for cash on hand and inventory stocks.

where e_t denotes the real exchange rate.[14] If the real exchange rate is constant, $\hat{\Omega}_{jt}$ is equal to unity for all firms. In periods when the exchange rate depreciates, e_t/e_{t-1} rises and $\hat{\Omega}_{jt}$ rises with the depreciation in proportion to the firm's foreign-denominated debt share.

Movements in the balance sheet occur for one of two reasons, a rise in the overall level of indebtedness $b_{j,t}/a_{j,t}$ or an increase in the value of debt outstanding through changes in the exchange rate variable $\hat{\Omega}_{jt}$. Because $b_{j,t}/a_{j,t}$ is measured at the beginning of the period, within-period movements in $\hat{\Omega}_{jt}$ $(b_{j,t}/a_{j,t})$ are entirely attributable to movements in the exchange rate. Because the foreign-denominated debt ratio is firm specific, such variation has firm-specific effects, causing a greater deterioration of the balance sheet for firms who rely relatively more on foreign-denominated debt.

In addition to our measures of the balance sheet and fundamentals, the regression controls for firm and time fixed effects. Time dummies capture a common investment component owing to macroeconomic influences working through either output or prices. Firm fixed effects are included to control for firm-level heterogeneity in the average investment rate of firms. Such heterogeneity may arise either because the mean level of fundamentals differs, or the cost of investing differs across firms in some systematic way. Finally, for the sake of robustness, the regressions also allow for serial correlation in the investment process by including lagged investment on the right hand side of the regression.

Our empirical investment equation is

$$(i/k)_{j,t} = c + c_j + \rho(i/k)_{j,t-1} + \boldsymbol{\alpha}'(\mathbf{s}/\mathbf{k})_{j,t} + \beta(\hat{\Omega}b/a)_{j,t} + \delta_t + \epsilon_{j,t} \tag{16}$$

where $(i/k)_{j,t}$ is investment normalized by the tangible capital stock, $(s/k)_{j,t}$ is a vector of domestic and foreign sales normalized by the tangible capital stock, $[(s_d/k)_{j,t}\ (s_f/k)_{j,t}]$, $\alpha = [\alpha_d\ \alpha_f]$ is a vector of coefficients measuring the effect of fundamentals on investment, δ_t is a time dummy and c_j is the firm-specific fixed-effect.

As a robustness check, we also estimate another version of the empirical investment equation which considers separately the effects of the devaluation given the average foreign debt ratio and the overall beginning of period leverage ratio:

$$(i/k)_{j,t} = c + c_j + \rho(i/k)_{j,t-1} + \boldsymbol{\alpha}'(\mathbf{s}/\mathbf{k})_{j,t} + \beta\hat{\omega}_j(e_t/e_{t-1}) + \gamma(b/a)_{j,t} + \delta_t + \epsilon_{j,t} \tag{17}$$

[14] In the model $\Omega(e, e_{-1}; \omega)$ depends on the innovation in the exchange rate rather than the ratio e/e_{-1} as in equation 15. For our reduced-form analysis we use the latter formulation since it does not require us to compute expectations. In our structural estimation, we estimate the auxilliary regression using the model's version of Ω_{jt} so that simulated and actual regressions are correctly specified to match each other.

This regression isolates the heterogenous effect that the exchange rate has on firm-level investment owing to differences in firms' pre-crisis foreign-denominated debt ratio.

In the absence of capital market imperfections, standard adjustment cost theory predicts that $\beta = \gamma = 0$ under the assumption that $(\mathbf{s}/\mathbf{k})_{j,t}$ properly measures fundamentals. In general, fundamentals depend on the entire present discounted value of future profit rates. If $(\mathbf{s}/\mathbf{k})_{j,t}$ follows an AR1 process, then the present value $(\mathbf{s}/\mathbf{k})_{j,t}$ is proportional to the current value $(\mathbf{s}/\mathbf{k})_{j,t}$, and fundamentals are properly measured. If $(\mathbf{s}/\mathbf{k})_{j,t}$ follows a richer stochastic process, our proxy for fundamentals introduces measurement error into the equation however.

A frequent concern in the investment literature is that balance sheet measures may enter investment equations significantly because the regression does not properly measure fundamentals. Firms in our data set that hold greater levels of foreign-denominated debt tend to have higher ratios of exports to total sales. In the absence of financial frictions, an exchange rate depreciation is more likely to be a positive shock to fundamentals for firms with high foreign-denominated debt ratios. Thus, if fundamentals are measured with error, the estimation procedure is biased against finding a negative effect of the balance sheet working through the exchange rate mechanism on investment.[15]

4.2 Econometric Methodology

To estimate equations 16 and 17, we consider two estimators: an IV version of a fixed-effect estimator and a panel-data GMM estimator. Instrumental variables control for the endogeneity that may exist between current sales and current investment.[16] The IV estimator is a standard 2SLS estimator that controls for fixed effects by removing group means. This estimator is adopted in part for its simplicity. It controls for firm-level heterogeneity and provides a reasonable summary of the data without applying complicated instruments sets

[15] Additional biases may occur owing to the following reasons: membership in conglomerates (Chaebol); survivorship bias; and firm-specific imported material inputs. Firms that export are more likely to be members of industrial groups that provide partial insurance through cross-firm financial arrangements. Such insurance may mitigate the effect of the individual firm's balance sheet on investment. Although the firm-specific fixed factor partially controls for such differences, by omitting specific controls for the balance sheets of the conglomerate, our estimates would understate the strength of financial frictions. Similarly, by confining our attention to a balanced panel, our estimates may display survivorship bias which also biases the effect of the balance sheet to zero. Because imported material shares may be positively correlated with export status, mismeasurement of this share would produce bias in the oppositie direction however, i.e. the foreign debt ratio may proxy for the imported materials share. To control for this possibility, we use the ratio of foreign-denominated trade payables to total trade payables as a proxy for this share. Adding this variable to the regression does not change our coefficient estimates for β or γ.

[16] The IV estimator using lagged sales as an instrument controls for simultaneity bias between current sales and current investment, in the event that time to build is less than one year.

or weighting matrices. The IV estimator thus has the virtue that it is easy to apply when estimating the structural model through indirect inference.

The IV estimator has some limitations for pure regression analysis however. In particular, in the presence of lagged dependent variables, such estimators are inconsistent. We therefore also consider the more general GMM panel-data estimation procedure proposed by Arellano and Bond (1991). This estimator uses first differences to eliminate the fixed effect. First differencing introduces serial correlation in the error term which can be controlled for through the appropriate choice of instruments.

After taking first differences, equation 16 may be expressed as

$$\begin{aligned} \Delta(i/k)_{j,t} &= \rho\Delta(i/k)_{j,t-1} + \boldsymbol{\alpha}'\Delta(\mathbf{s}/\mathbf{k})_{j,t} + \beta\Delta(\hat{\Omega}b/a)_{j,t} + \delta_t + v_{j,t} \\ v_{j,t} &= \epsilon_{j,t} - \epsilon_{j,t-1} \end{aligned} \tag{18}$$

Since the sales variables are treated as endogenous and the lagged dependent variable, $\Delta(i/k)_{j,t-1}$ is correlated with the error term, $v_{j,t} = \epsilon_{j,t} - \epsilon_{j,t-1}$, by construction, $(i/k)_{j,t-s}$ and $(s/k)_{j,t-s}$ are valid instruments for $s \geq 2$. The balance-sheet variable is treated as a predetermined variable and therefore, $(\hat{\Omega}b/a)_{j,t-s}$ are valid instruments for $s \geq 1$. We use the two-step version of Arellano and Bond(1991) GMM estimator where the residuals of the first-step estimation are used to construct the optimal weighting matrix for the second-step estimator. We also provide the results of overidentifying restriction tests in the tables. For the fixed-effect IV estimator, we use $(s/k)_{j,t-s}$ for $s \geq 1$ and $(\hat{\Omega}b/a)_{j,t-s}$ for $s \geq 1$ as instruments. When estimating equation 17 which considers the separate effects of $\hat{\Omega}_{jt}$ and $b/a_{j,t-s}$, we use lags of $\hat{\Omega}_{j,t-s}$ and $b/a_{j,t-s}$ as separate instruments in both the IV fixed-effect estimator and the GMM estimator.

4.3 Data

Our data set is a unique, proprietary data set of Korean manufacturing firms. The data set is provided by KIS (Korea Information System). It provides income-statement and balance sheet data for all listed manufacturing companies over the period 1993 to 2002. Unlike Compustat data, the standard data set used for U.S. firm-level investment studies, the KIS data provide distinct information on the value of foreign versus domestically denominated debt, and foreign versus domestic sales.

Table 1 provides summary statistics, constructed for the full sample, and before and after

Table 1: Summary Statistics

	Full Sample Mean	Full Sample Std. Dev.	Pre-Crisis Mean	Post-Crisis Mean
$(i/k)_{j,t}$	0.169	0.244	0.230	0.136
$(s/k)_{j,t}$	3.756	3.195	3.939	3.657
$(\pi/k)_{j,t}$	0.764	0.866	0.785	0.753
$(b/a)_{j,t}$	0.371	0.211	0.392	0.363
$(s^f/s)_{j,t}$	0.284	0.279	0.251	0.307
$(b^f/b)_{j,t}$	0.140	0.189	0.140	0.140
$corr\ (s^e/s, b^e/b)$	0.1669		0.251	0.120

the onset of the crisis. The mean rate of investment fell from 23 percent pre-crisis to 13.6 percent post-crisis. Exports as a fraction of total sales rose form 25 percent pre-crisis to 30.7 percent post-crisis while overall profitability and overall sales fell slightly during the post-crisis period. These numbers are consistent with the figures displayed above. The last row of table 1 provides information on the correlation between foreign exchange earnings and foreign-denominated debt. The correlation is 0.17 over the entire sample period, and somewhat higher than that during the pre-crisis period (0.25). Thus, firms who access foreign-denominated debt markets are more likely to be export-oriented firms.

Table 2 provides information on the quantile distribution of firms's pre-crisis averages of export-sales ratios, leverage ratios and foreign-denominated debt ratios. This information is explicitly used to calculate a distribution of firm types embedded in the structural estimation described below. The median firm in the sample has an export/sales ratio of 15 percent while nearly 25% of the firms have almost no exports. Likewise, the median firm in the sample has a foreign-denominated debt ratio of eight percent. Importantly, there is considerable variation in the foreign-denominated debt ratio, the key variable measuring the heterogeneity in the balance sheet effect of the devaluation across firms.

To complete our summary statistics, Table 3 considers the determinants of foreign-denominated debt. This table reports results from a regression of ω_j, the foreign-denominated debt ratio on the export-sales ratio (s^f/s), the debt-to-asset ratio (b/a), and the log of assets, $\log(a)$ as a proxy for firm size. All variables are computed as firm-specific means. In the first regression, these means are computed over the pre-crisis period. In the second regression we compute the means using the full sample.

Table 2: Quartile Distribution of Pre-Crisis Firm Means

	0%	25%	50%	75%	100%	mean
$(b/a)_j$	0.000	0.261	0.399	0.504	1.632	0.391
$\left(s^f/s\right)_j$	0.000	0.034	0.158	0.419	0.983	0.255
$\left(b^f/b\right)_j$	0.000	0.024	0.081	0.185	1.000	0.141

Table 3: Determinants of Foreign-Debt Ratio

	s^f/s	b/a	$log(a)$	R^2
Pre-Crisis:	0.154	-0.276	0.017	0.15
	(0.028)	(0.044)	(0.006)	
Full-Sample:	0.124	-0.182	0.038	0.10
	(0.026)	(0.039)	(0.029)	

Table 3 highlights the finding that firms with high foreign-denominated debt ratios are firms who have a higher propensity to export. Such firms also tend to have stronger balance sheets as measured by the debt-to-asset ratio. Finally, the data show a modest size effect – controlling for exports and leverage, firms with high levels of foreign-denominated debt tend to be larger firms. The non-randomness in foreign-denominated debt ratios justifies explicitly controlling for firm factors through fixed effects in our reduced-form investment regression. It also motivates our firm-specific controls used in the structural estimation.

4.4 Estimation Results

Table 4 summarizes the main empirical findings using both IV fixed effects and the first-differenced GMM specification.[17] The first column of estimates reported in Table 4 include

[17] Our structural estimation reported below is conducted with a balanced panel of firms. Accordingly, we confine our attention to the balanced panel when reporting reduced form estimation results though we have estimated all regressions using both the balanced and unbalanced panels. We find little difference between these estimates – the coefficient on the balance-sheet variable is slightly smaller for the balanced panel, which is consistent with the notion that selection induced by the balanced-panel biases our estimates towards higher quality firms with less severe financial frictions.

Table 4: Investment Equation

	IV Fixed Effects		First Diff. GMM	
	$(i/k)_{j,t}$	$(i/k)_{j,t}$	$(i/k)_{j,t}$	$(i/k)_{j,t}$
$(s^d/k)_{j,t}$	0.069	0.069	0.054	0.051
	(0.007)	(0.006)	(0.006)	(0.022)
$(s^e/k)_{j,t}$	0.047	0.047	0.035	0.035
	(0.011)	(0.011)	(0.005)	(0.005)
$(\hat{\Omega}b/a)_{j,t}$	-0.208	–	-0.177	–
	(0.037)		(0.041)	
$(b/a)_{j,t}$	–	-0.194	–	-0.160
		(0.038)		(0.049)
$\hat{\omega}_j e_t$	–	-0.503	–	-0.205
		(0.124)		(0.074)
$(i/k)_{j,t-1}$	–	–	0.204	0.201
			(0.018)	(0.022)
Rsq (within)	0.19	0.20	–	–
Sargan	–	–	106.34	105.89
(p-val)	–	–	(0.39)	(0.17)
m2	–	–	-0.22	-0.29
(p-val)	–	–	(0.83)	(0.77)
No. of Obs.	2490	2490	1990	1990
No of Inds.	419	419	412	412

the sales-to-capital ratios (both domestic and foreign) along with the balance sheet variable $(\hat{\Omega}b/a)_{j,t}$. Fundamentals, as measured by the sales-to-capital ratios, have a statistically significant positive effect on investment. The coefficient on the balance-sheet variable is negative and highly statistically significant. At the mean value of the foreign-denominated debt and leverage ratios ($\omega_j = 0.14, b/a = 0.4$), the estimated coefficient on $(\hat{\Omega}b/a)_{j,t}$ suggests that the 70% devaluation reduces the investment rate by 60 basis points through the balance sheet mechanism. This mechanism can thus account for only a small fraction of the 10 percentage point reduction in the rate of investment that occurred during the crisis.

The second column of table 4, decompose the balance-sheet effect into two terms – the beginning-of-period debt-level $(b/a)_{j,t}$ and the exchange rate interacted with the pre-sample foreign-denominated debt ratio $\hat{\omega}_j e_t$. Because the regression includes a full set of time dummies, the coefficient on $\hat{\omega}_j e_t$ captures the heterogenous effect of the exchange rate on investment owing to the fact that firms face different degrees of foreign-debt exposure at the onset of the crisis. Both balance sheet variables are negative, statistically significant, and quantitatively large.

The third and fourth columns of Table 4 report the GMM estimates based on first-differencing. These estimates include the lagged dependent variable for robustness. The coefficient estimates on the balance sheet variables are again negative, quantitatively large and statistically significant. When the balance sheet is split into its two components (column 4) we again find an independent effect of the exchange rate interacted with the pre-sample foreign debt ratio. In all regressions, the coefficient on the lagged dependent variable is statistically significant though relatively small in magnitude.

In table five, the devaluation is allowed to have non-linear effects which depend on the firm's pre-crisis export and foreign-denominated debt position. To do so, the sample is divided into four sub-groups based on whether firms are high vs low export-sales ratios and high vs low foreign-denominated debt ratios. These classifications are based on the median pre-crisis averages of export-sales and foreign-denominated debt ratios. For parsimony, only the GMM estimates are reported.

Table 5: Investment Equation

First Differenced GMM by sub-groups

	H-fob/L Exp $(i/k)_{j,t}$	H-fob/H-exp $(i/k)_{j,t}$	L-Fob/L-exp $(i/k)_{j,t}$	L-Fob/H-exp $(i/k)_{j,t}$
$(s^d/k)_{j,t}$	0.060	0.082	0.041	0.058
	(0.001)	(0.004)	(0.002)	(0.002)
$(s^e/k)_{j,t}$	0.028	0.064	0.150	0.041
	(0.001)	(0.002)	(0.014)	(0.003)
$(\hat{\Omega}b/a)_{j,t}$	-0.406	-0.203	-0.197	-0.021
	(0.023)	(0.019)	(0.026)	(0.013)
$(i/k)_{j,t-1}$	0.145	0.148	0.130	0.149
	(0.004)	(0.005)	(0.011)	(0.009)
Sargan	57.13	100.28	88.91	58.97
	(0.99)	(0.56)	(0.84)	(0.99)
m2	-0.63	-0.99	0.51	-0.94
	(0.53)	(0.32)	(0.61)	(0.35)
No of Obs.	349	640	686	315
No of Inds.	70	137	136	69

Firms who are most vulnerable to the exchange rate shock – firms with low exports and high foreign debt – exhibit the greatest sensitivity of investment to the balance sheet variable. The coefficient on the balance sheet is -0.401 and highly significant. Firms who are least vulnerable – firms with high exports and low foreign-denominated debt ratios exhibit essentially no response of investment to the balance sheet. As expected, the other two categories, low foreign debt/high exports and high foreign debt/low exports, exhibit responses that are between these extremes.

In summary, the response of investment to the exchange rate devaluation is consistent with the notion that credit frictions working through the balance sheet were a determining factor. The devaluation depressed investment for firms whose financial position was most exposed to exchange rate shocks. In particular, the balance sheet mechanism is strongest for firms with a significant currency mismatch between export exposure and debt exposure.

5 Structural Estimation

Structural estimation proceeds in two stages. The first stage derives a parametric form of the profit function and applies conventional panel-data econometric techniques to identify relevant structural parameters. It also determines the forcing processes for the macroeconomic variables. The parameters of these forcing processes determine future expectations. The second stage uses indirect inference to estimate the structural parameters that determine adjustment costs and financial frictions. The estimated structural parameters are then used to evaluate the role that financial factors and foreign-denominated debt play in propagating the financial crisis through investment spending.

When identifying the role of foreign-denominated debt on investment, the estimation procedure recognizes that firms who issue foreign-denominated debt are non-representative. In particular, such firms often issue foreign-denominated debt to hedge against foreign earnings and are thus more likely to be exporters than other firms. To allow for this possibilities, the structural estimation explicitly accounts for firm-level heterogeneity observed in the data. In particular, the estimation strategy conditions on the underlying distribution of export composition, foreign-denominated debt ratios and leverage.

5.1 Production Technology, Market Structure and Profitability

To derive a closed-form profit function, firm j is assumed to produces two differentiated goods – domestic and foreign – with a constant-returns-to-scale Cobb-Douglas technology. Although the firm produces two differentiated goods, it employs only one type of capital, $k_{j,t}$ and the production processes of both goods are subject to the same iid productivity shock. The production technology also allows for both domestic variable inputs such as labor and foreign variable inputs such as imported materials. In this framework, a firm with a given level of technology and capital chooses how to allocate variable inputs across the domestic and foreign markets to maximize profits. The firm faces monopolistic competition in both markets. Demand is assumed to be iso-elastic demand and the demand elasticities are allowed to differ across the domestic and foreign markets.

Under these assumptions, the closed-form profit function of a firm can be expressed as a weighted average of sales in each market

$$\begin{aligned} \pi_{j,t} &= \sum_{i=d,f} \phi_{i,j}(\mathbf{z}_t) k_{j,t}^{\gamma_i} - \sigma_j \\ &= \sum_{i=d,f} \Gamma_i s_{i,j,t} - \sigma_j \end{aligned} \tag{19}$$

where the weights are determined by the mark-up $1/\chi_i$ in each market combined with the production share of capital $v\alpha$.

$$\Gamma_i = 1 - \chi_i(1 - \nu\alpha). \tag{20}$$

Because the profit function and the sales function are identical up to a scaler, Γ_i, the structural parameters of the profit function can be identified by estimating distinct sales equations for the domestic and foreign markets.[18] By maximizing over the variable factor inputs, sales in each market may be expressed as a log-linear function of a fixed firm factor $\theta_{i,j}$, the individual firm's capital k_{jt}, the exchange rate e_t, a common (aggregate) demand component y_{it} and an idiosyncratic error $v_{i,t}$. For estimation purposes, we allow the idiosyncratic error to follow an AR1 process, in which case we obtain the following fixed-effect regression

[18] Separate accounting data are available for domestic and foreign sales but not earnings.

Table 6: Profit Function: Export vs Domestic Sales

	$\log e_t$	$\log k_{j,t}$	$\log a_{d,t}$	$\log a_{f,t}$	rho_v	R^2	Obs/Inds
$\log s_{f,t}$	0.360	0.545	–	5.355	0.325	0.41	2544
	(0.086)	(0.038)	–	(1.76)			416
$\log s_{d,t}$	-0.120	0.412	1.479	–	0.223	0.62	2847
	(0.052)	(0.024)	(0.198)	–			441

specification with AR(1) error term, developed by Baltagi and Wu(1999) and Baltagi(2000),

$$\begin{aligned} \ln s_{i,j,t} &= \varsigma_i \ln \theta_{i,j} + \gamma_i \ln k_{j,t} + \xi_i \ln e_t + \alpha_i \ln y_{i,t} + v_{i,t} \\ v_{j,t} &= \rho_v v_{j,t-1} + u_{j,t}, \qquad u_{j,t} \sim \text{iid } N(0, \sigma_u^2) \end{aligned} \tag{21}$$

for $i = d, f$. The elasticity of sales with respect to capital in equation 21 provide estimates of the implied markups in the domestic and foreign markets through the relationship

$$\gamma_i = \frac{\nu \alpha \chi_i}{1 - \chi_i(1 - \nu\alpha)}.$$

To estimate the sales equations, all variables are expressed as real quantity values deflated by appropriate price indices. The domestic demand shifter $y_{d,t}$ is the HP-detrended Korean log-GDP. The foreign demand shifter $y_{f,t}$ is the HP-detrended index of world income obtained from the World Economic Outlook (WEO) data base obtained from the IMF.

Table 6 reports the estimation results for the profit function. The coefficient estimates for γ_i, the elasticity of sales with respect to capital, are significant and in line with other estimates obtained in the literature. Using an estimate of the production share of capital $v\alpha = 0.225$, the implied mark-ups are moderate and somewhat stronger in the domestic market (1.32) than in the foreign market (1.18).[19]

The estimated exchange rate coefficients indicate that domestic sales respond negatively

[19] The choice of $v\alpha$ allows us to infer an implied markup consistent with the curvature estimates γ_i. Setting $v\alpha = 0.225$ is consistent with the profit to sales ratio observed in the data given our estimate of fixed costs. This choice is also roughly consistent with estimates provided by Kim and Park (2000) and Park (1999).

to an exchange rate devaluation ($\hat{\xi}_d = -0.12$) while exports respond positively to the devaluation ($\hat{\xi}_f = 0.36$). The negative response of domestic sales is consistent with the reliance of domestic sales on foreign inputs. These estimates also imply a threshold value, 0.25, above which a firm's profit is increasing in the real exchange rate.[20] This threshold value is a greater than the median export-sales ratio(0.203) and smaller than the mean export-sales ratio(0.284) in the sample. Thus, on average, movements in the real exchange rate do not exert a strong influence on competitiveness in the Korean manufacturing sector.

Given a firm's state variables $[k_{j,t}, e_t, y_{d,t}, y_{f,t}]$, the elasticities $[\gamma_i, \xi_i, \alpha_i]$ combined with the estimates of the idiosyncratic forcing process $[\rho_v, \sigma_u^2]$ are sufficient to characterize the sales process in each market. To reconstruct the profit function we also need the weights Γ_i, an estimate of the fixed cost σ_j, and a methodology to determine the productivity factors $\theta_{i,j}$. The weights Γ_i are determined directly from equation 20. The fixed cost σ_j is assumed to be proportional to the steady-state value of sales for firm j. The fixed cost may then be estimated as a fraction of sales using the average overhead costs to sales ratio reported in the data. Under this formulation, one may define an appropriate normalizing factor to render the firm problem scale invariant. The ratio of exports to total sales determines the relative productivity of exports and hence the ratio $\theta_{f,j}/\theta_{d,j}$. The productivity levels $\theta_{f,j}$ and $\theta_{d,j}$ relate to the steady-state size of the firm and are determined by the normalizing factor. The appendix provides further details of these steady-state calculations.

5.2 Macroeconomic Shock Processes

To specify a stochastic process for the real interest rate, we decompose the domestic risk free rate into subcomponents

$$\begin{aligned} 1 + r_d &= (1 + r_f) E\left(e/e_{-1} | \mathbf{z}_{-1}\right) \\ &= (1 + \bar{r})(1 + \xi) E\left(e/e_{-1} | \mathbf{z}_{-1}\right) \end{aligned} \quad (22)$$

where $1 + r_f$ is the risk free rate on foreign bonds which has two components, the foreign interest rate, $\bar{r}$, which we take as a constant, and the country risk premium, ξ.[21] The exchange

[20] The elasticity of gross profits with respect to the exchange rate may be approximated by $\left(1 - \zeta_{f,j}\right) \hat{\xi}_d + \zeta_{f,j} \hat{\xi}_f$ where $\zeta_{f,j}$ denotes the export to total sales ratio.

[21] One can think of $\bar{r}$ as the real US Treasury Bond rate and ξ as the spread on the emerging market government bond. To construct the real interest rate data, we use an inflation forecast based on AR(1) process.

rate is assumed to follow an AR(1) process with persistence parameter ρ. The country risk premium is specified as an AR(1) process in logs, i.e., $\xi = \bar{\xi}^{1-\varphi}\exp(\varepsilon)\xi_{-1}^{\varphi}$, where $\bar{\xi}$ is the normal level of the country risk premium.

In log deviations, equation 22 implies

$$r_d = \bar{r} + (\rho - 1)\log e_{-1} + \xi.$$

Substituting the data generating process for the country risk premium:

$$r_d = \bar{r} + (\rho - 1)\log e_{-1} + \bar{\xi}\,(1 - \varphi) + \varphi\xi_{-1} + \varepsilon.$$

Lagging this equation one period and solving for ξ_{-1}, we have:

$$r_d = (1 - \varphi)\left(\bar{r} + \bar{\xi}\right) + (\rho - 1)\left(\log e_{-1} - \varphi \log e_{-2}\right) + \varphi r_{d-1} + \varepsilon. \tag{23}$$

Equation 23 implies the following time-series model for the real interest rate is

$$r_d = a_1 + a_2 \log e_{-1} + a_3 \log e_{-2} + a_4 r_{d-1} + \varepsilon \tag{24}$$

where $a_1 \equiv (1 - \varphi)\,\bar{r}$, $a_2 \equiv (\rho - 1)$, $a_3 \equiv -\varphi\,(\rho - 1)$, and $a_4 = \varphi$.

OLS estimates for the exchange rate process over the pre-crisis period 1966-1997 imply $\hat{\rho} = 0.804$, while estimates over the period 1966-2002 with a crisis dummy included imply $\hat{\rho} = 0.89$. If the UIP condition holds, the persistence parameter estimated from the exchange rate process, $\hat{\rho}$ must be closed to $1 + \hat{a}_2$. Also, $\hat{a}_3$ must be close to $-\hat{a}_4\,(\hat{\rho} - 1)$. Under UIP, $\hat{a}_1/(1 - \hat{a}_4)$ may be interpreted as the real interest rate in the foreign country plus the normal level of country risk premium.

Table 7 provides estimation results for equation 24. It reports both OLS and MLE estimates and include a crisis dummy, though estimation results for the interest rate process are not particular sensitive to the inclusion of this variable. Given a persistence parameter for the real exchange rate, $\hat{\rho}$,between 0.8 to 0.89,. $\hat{a}_2$ should range between -0.1 to -0.19. The actual estimates fall within this range.[22] The coefficient a_4 measures the degree of persistence

[22]Using pre-crisis data, the persistence parameter is estimated to be 0.801. Using the full-sample combined

Table 7: Interest Rate Process

	a_1	a_2	a_3	a_4	$1(t = 1998)$
MLE (1966-2002)					
Estimate	0.042	-0.118	-0.049	0.665	0.018
std errors	(0.016)	(0.063)	(0.024)	(0.312)	(0.402)
OLS (1966-2002)					
Estimate	0.015	-0.136	0.035	0.560	0.015
std errors	(0.010)	(0.023)	(0.040)	(0.196)	(0.015)

in the country-risk premium. The estimates vary between 0.56 and 0.665. Finally, estimation results imply a long-run real interest rate of 3% to 12% depending on the specification.

Based on these values, we set the steady-state risk-free rate $r_d = 5\%$, the degree of persistence for the country risk premium $\varphi = 0.6$, and choose $\rho = 0.85$ as our baseline estimate for the degree of persistence in the exchange rate process. Estimation and simulation results are highly robust to reasonable variation in the choice of the steady-state risk free rate and the degree of persistence in the country-risk premium. Because the degree of persistence in the exchange rate process plays a key role in the dynamics of the expected future interest rate path under UIP, we conduct a variety of robustness checks for both structural estimation and model simulation to this parameter.

In addition to the exchange rate and interest rate process, our model requires us to specify a stochastic process for the aggregate demand shifter in the sales equation. These shifters include Korean real GDP for domestic sales and World GDP for foreign sales. Estimating an AR1 process for the HP detrended log of domestic GDP over the sample period, 1990-1997, implies $\rho_A = 0.7$.[23]. Because world output shows only small variation over this period, it is assumed to be fixed at a constant value.

5.3 Indirect Inference

This section applies indirect inference to estimate the two structural parameters of the model that govern the investment process, one for the capital adjustment cost and the other for the

with a dummy variable for the crisis year, the persistence parameter is estimated to be 0.897.

[23] Estimating the Korean GDP process for 1970-2002 gives an implied persistence $\rho_A = 0.3$ which is extremely low for the detrended log-level of GDP. Nonetheless, setting $\rho_A = 0.3$ provides structural parameter estimates that are very close to those obtained when setting $\rho_A = 0.7$.

agency cost, κ. The adjustment cost function is specified as

$$c(i,k) = \frac{\psi}{2}\left(\frac{i}{k} - \delta\right)^2 k.$$

The agency cost function is specified as

$$\kappa\left[\exp\left(x\right) - 1\right]$$

where x is a measure for the firm's financial burden properly normalized by firm assets, namely the leverage ratio defined in equation 11. Under the null hypothesis of no financial market frictions, the estimated value of κ should be close to zero.

Indirect inference uses a criterion function derived from an auxiliary statistical model which may be estimated from both the actual data and the simulated data obtained from the structural model. The structural parameter vector $\boldsymbol{\theta} = [\psi, \theta]'$ is chosen so that the auxiliary model's parameter estimates obtained from the simulated data are close to the parameter estimates obtained from the actual data.

Denote the criterion function for the auxiliary model applied to the real data by Q. The estimate of the auxiliary model can be defined as

$$\hat{\beta} = \arg\max_{\beta} Q_T(\mathbf{x}_T; \beta)$$

where x_T is a data matrix and T is the number of observations. In the case of panel data, T implies the product of the number of time observations and the number of individuals. Following Gourieroux et al.(1993), define the binding function, $\beta = b(\boldsymbol{\theta})$ as a simulated counterpart of $\hat{\beta}$, i.e., a solution to $E_\theta\left[\partial Q(\mathbf{x}; b\left(\boldsymbol{\theta}\right))/\partial b(\boldsymbol{\theta})\right] = 0$. In actual estimation, the binding function is replaced by its empirical counterpart,

$$\hat{b}_S(\boldsymbol{\theta}) = \frac{1}{S}\sum_{s=1}^{S} \hat{\beta}_T^{(s)}(\boldsymbol{\theta})$$

where S is the number of simulations. The minimum distance estimator of the structural parameter vector, $\boldsymbol{\theta}$, is defined as

$$\hat{\boldsymbol{\theta}}_{MD}^{S} = \arg\min \left[\hat{\beta} - \hat{b}_S(\boldsymbol{\theta})\right]' W \left[\hat{\beta} - \hat{b}_S(\boldsymbol{\theta})\right]$$

where W is a positive-definite matrix. As the sample size goes to infinity, the indirect inference estimator $\hat{\boldsymbol{\theta}}_{MD}^{S}$ is consistent and asymptotically normal for any fixed S. The asymptotically optimal weighting matrix is

$$W_0 = A_0 B_0^{-1} A_0$$

where

$$A_0 = \lim_{T \longrightarrow \infty} E\{\partial^2 Q(\mathbf{x}; \beta)/\partial\beta_0 \partial\beta_0'\}$$

and

$$I_0 = \lim_{T \longrightarrow \infty} var\{\sqrt{T}\partial Q(\mathbf{x}; \partial\beta)/\partial\beta_0 - E[\sqrt{T}\partial Q(\mathbf{x}; \partial\beta)/\partial\beta_0 | \mathbf{x}]\}.$$

With this choice of the weighting matrix, the asymptotic distribution of the indirect inference estimator satisfies

$$\sqrt{T}(\hat{\boldsymbol{\theta}}_{MD}^{S} - \theta_0) \xrightarrow{d} N(0, avar(\hat{\boldsymbol{\theta}}_{MD}^{S}))$$

where $avar(\hat{\boldsymbol{\theta}}_{MD}^{S}) = (1 + 1/S)[\partial b(\boldsymbol{\theta}_0)/\partial\boldsymbol{\theta} W_0 \partial b(\boldsymbol{\theta}_0)/\partial\boldsymbol{\theta}']^{-1}$

The asymptotic efficiency of the estimator depends on how well the auxiliary model captures the properties of the original structural model. In our case, the auxiliary model should reflect two fundamental aspects, namely the influences of both the investment fundamentals and the financial frictions, controlling for important individual characteristics. The reduced form regression used in section 4,

$$(i/k)_{j,t} = c_j + \beta^d (s_d/k)_{j,t} + \beta^f (s_f/k)_{j,t} + \beta^{fd} (\hat{\Omega} b/a)_{j,t} + \delta_t + \varepsilon_{j,t}$$

is well suited for these requirements. The sales-to-capital ratios and the balance-sheet term control for fundamentals and financial conditions in a parsimonious way, while the fixed-effect allows for heterogeneity in investment rates across firms that may be correlated with either profitability or financial factors.[24]

When generating the simulated data used to estimate the structural model, we also wish

[24]When matching the model to the data, we assume that total assets in the model are equal to current profits plus undepreciated capital, i.e. $a = \sum_{i=d,f} \phi_i(\mathbf{z}_{-1}) k^{\gamma_i} + (1 - \delta)k$.

to control for firm-level heterogeneity. To do so in a model consistent manner, we specify a firm-specific vector of individual characteristics, h_j. The vector h_j measures the firm-specific steady-state values of the foreign-denominated debt ratio and the export-sales ratio. The export-sales ratio may me mapped into the firm-specific structural parameters that determine the relative productivity of exports θ_f/θ_d. These firm-specific ratios are estimated using pre-crisis sample means. The dynamic programming problem of each individual in the simulation stage is then a function of this individual characteristics vector, h_j. Firms are also allowed to differ in their initial debt to capital or leverage ratios. While these differences do not affect the model solution, they are relevant when simulating the data for estimation purposes. In summary, individual firms are characterized by a vector, $h_j = [\omega_j, \theta_f/\theta_d, (b/k)_j]$ which is predetermined at the onset of the crisis.

The distributions of these individual characteristics are nondegenerate and chosen to replicate the distributions observed in the data prior to the onset of the financial crisis. The realized paths of the exchange rate, the country-risk premium and the macroeconomic shock are chosen to match the actual aggregate realizations on an annual basis. Table 11 in the appendix summarizes these variables.

The simulated panel data has the same number of time observations for each individual. Since we do not model exit behavior, the panel is balanced in both the simulated data and the actual data. For variance reduction, we compute $S = 100$ simulations. In other words, $\hat{b}_S(\boldsymbol{\theta})$ is an average of 100 IV Fixed Effect estimates.

Ideally, to completely control for firm-level heterogeneity, one would solve the value function and simulate the data for each firm in our sample. Because the data contain over 400 individual firms, it is a computationally formidable task to generate a simulated panel with the same number of individuals as the data however. To reduce the computational burden, our estimation procedure creates a simulated panel with a smaller number of individuals, but which replicates the distributions of individual characteristics in the data. This is done in the following way: i) Estimate the joint empirical distribution function for the three individual characteristics describe above. Since we rely on the quartile distribution, this procedure generates a panel with $4^3 = 64$ individuals. ii) Numerically solve the value function for the sixteen types (4^2) characterized by the quartiles of the distribution for ω_j and θ_f/θ_d. iii) For each of these sixteen model solutions, simulate the model for four separate initial leverage positions $(b/k)_j$. This procedure generates 64 time series for each simulation. iv) Apply a weighted average version of an IV Fixed Effect estimator to the simulated data. The weights are determined by the empirical probability of observing each of the 64 types. This procedure assumes that the data is well approximated by 64 individual types characterized by the individual characteristics described above. By relying on the joint empirical distribution to weight these types, our estimation procedure effectively controls for the fact that a firm who is a high foreign-debt type is also more likely to be a high export type in our estimation strategy.[25]

This procedure is used to estimate two structural parameters using three moments, namely, $[\hat{\beta}^d - \hat{b}^d_S(\boldsymbol{\theta}),\ \hat{\beta}^f - \hat{b}^f_S(\boldsymbol{\theta}), \hat{\beta}^f - \hat{b}^{fd}_S(\boldsymbol{\theta})]$. Consequently, the system is overidentified, and the choice of the weighting matrix matters for our estimates. The optimal weighting matrix is the inverse of variance-covariance matrix of the auxiliary parameter estimates in the real data, i.e. $\hat{W} = [T\hat{V}(\hat{\beta})]^{-1}$. This is the optimal weighting matrix under the null hypothesis that the model is correct. Because the system is over-identified, the minimized distance follows a chi-square distribution with the degree of freedom 1 and therefore provides the following Sargan test statistic of overidentifying restrictions:

$$J(\hat{\boldsymbol{\theta}}) = \frac{TS}{1+S}\left[\hat{\beta} - \hat{b}_S(\hat{\boldsymbol{\theta}})\right]' \hat{W} \left[\hat{\beta} - \hat{b}_S(\hat{\boldsymbol{\theta}})\right] \sim \chi^2(1).$$

5.4 Structural Estimation Results

We now report parameter estimates obtained from our indirect inference procedure. Because counterfactual simulations are sensitive to the assumptions made regarding both the persistence of the exchange rate process and the choice of the household discount factor $\Lambda(\mathbf{z}', \mathbf{z})$, the estimation procedure considers four distinct macroeconomic environments. These alternative parameter estimates are considered for robustness. In particular, although time effects are removed from both the model and the data when matching moments, structural estimates may be sensitive to the specification of the macroeconomic environment owing to nonlinearities inherent to the structural model. Also, model simulations consider alternative but plausible assumptions regarding agents' expectations of the macroeconomic processes.

The first three estimates assume that the firm discounts future profits using the domestic interest rate

$$\Lambda(\mathbf{z}', \mathbf{z}) = \frac{\beta}{1 + r_d}$$

[25] An early version of this paper assumed that firms are heterogeneous in their steady-state leverage positions owing to differential access to capital markets. This can be motivated by allowing the survival probability μ to be heterogenous across firms. This assumption requires one to solve the value function separately for 64 types rather than sixteen which implies a four-fold increase in computational costs. Because this dimension of heterogeneity made very little difference in our estimation results, we adopted the simplifying assumption that all firms have the same long-run access to capital markets but differ in their initial leverage positions.

Table 8: Auxilliary Parameter Estimates

	$(s^d/k)_{jt}$	$(s^e/k)_{jt}$	$(\hat{\Omega}b/a)_{jt}$
Data Moments	0.0692	0.0465	−0.2075
Simulated Moments			
1. $\Lambda(\mathbf{z}_{t+1}, \mathbf{z}_t) = (1 + r_{t+1})^{-1}$			
$\rho_e = 0.85$	0.0689	0.0464	−0.2060
$\rho_e = 0.90$	0.0687	0.0468	−0.2006
$\rho_e = 0.95$	0.0683	0.0405	−0.1931
2. $\Lambda(\mathbf{z}_{t+1}, \mathbf{z}_t) = \beta\,(C_{t+1}/C_t)^{-1}$			
$\rho_e = 0.85$	0.0648	0.0422	−0.1850

The degree of persistence in the real exchange rate is then allowed to vary: $\rho = 0.85, 0.9, 0.95$. By imposing uncovered interest parity, the persistence of the real exchange rate process influences firms' beliefs regarding the expected future interest rate path for a given degree of persistence in the country-risk premium.

Imposing UIP is a strong assumption however. Therefore, the model is also estimated under the assumption that households have log-utility and the firm discounts future profits using the household discount factor:

$$\Lambda(\mathbf{z}', \mathbf{z}) = \beta \frac{C_t}{C_{t+1}}$$

Using the household discount factor avoids imposing uncovered interest rate parity. To avoid adding an additional state variable, C_t is assumed to move one for one with the (HP filtered) Korean real GDP which is already included as a state variable in the model.

Table 8 reports the auxiliary regression coefficients obtained from both the model and the data. For all four estimates, the model successfully matches the auxiliary coefficients obtained from the IV fixed effect regression in the data. Let the case $\Lambda(\mathbf{z}', \mathbf{z}) = \frac{\beta}{1+r_d}$ and $\rho = 0.85$ denote the baseline case. In this case, the coefficients for the domestic and foreign sales to capital ratios obtained from the model are 0.0689 and 0.0464. The coefficients obtained from the data are 0.0692 and 0.0456. The model does an equally successful job matching the coefficient on the balance-sheet variable — (-0.2060) in the model versus (-0.2075) in the data. The other cases considered provide only slight differences in the estimates for the auxiliary coefficients.

Table 9: Estimates of Structural Parameters

	$\hat{\psi}$ (s.e)	$\hat{\kappa}$ (s.e)	$\hat{J}$ ($p-val.$)
1. $\Lambda(\mathbf{z}_{t+1}, \mathbf{z}_t) = (1 + r_{t+1})^{-1}$			
$\rho_e = 0.85$	**0.9530**	**0.1443**	**0.0042**
	(0.0497)	(0.0395)	(0.9483)
$\rho_e = 0.90$	**0.9569**	**0.1355**	**0.0442**
	(0.0762)	(0.0220)	(0.8335)
$\rho_e = 0.95$	**1.0670**	**0.1429**	**0.6544**
	(0.1740)	(0.0116)	(0.4185)
2. $\Lambda(\mathbf{z}_{t+1}, \mathbf{z}_t) = \beta\,(C_{t+1}/C_t)^{-1}$			
$\rho_e = 0.85$	**1.0066**	**0.1284**	**0.5388**
	(0.0314)	(0.0139)	(0.4629)

Table 9 reports the structural parameters obtained from this estimation procedure, along with the test of over-identifying restrictions. For the baseline case, the adjustment cost parameter is estimated to be 0.9530 with a standard error of 0.0497. This estimate is similar to the structural estimates reported in Gilchrist and Himmelberg (1998) and Eberly et al. (2006). It is also much lower than what one would obtain using a Tobin's Q-style regression framework. Varying the degree of persistence in the exchange rate process, we obtain parameter estimates of the adjustment cost coefficient that vary between 0.9530 and 1.067.

The structural coefficients imply an important role for financial market imperfections in the investment process. For the baseline case, the coefficient measuring agency costs, κ, is estimated to be 0.1443 and highly significant. The model therefore clearly rejects the null hypothesis of no financial market imperfections. At the mean value of the leverage ratio, this estimate implies that a 10 percent increase in leverage implies an 84 basis point rise in the premium on external funds. Across the four parametrizations, the estimated value of κ varies between 0.1284 and 0.1443. Roughly speaking, these estimates imply that if leverage doubles, the cost of external finance rises by ten percentage points.

Finally, Table 9 also reports the J-statistics for the over-identifying restriction. According to this J-statistic, one cannot reject the model's over-identifying restriction for any of the four estimates provided.

Table 10 verifies that the model's structural parameters are well identified by the auxiliary

Table 10: The effects of conditional variations in the structural parameters

	$(s^d/k)_{jt}$	$(s^e/k)_{jt}$	$(\hat{\Omega}b/a)_{jt}$	$\hat{J}$
$\kappa = 0.1443$				
$\psi = 0.6000$	0.0904	0.0918	−0.3412	**62.180**
$\psi = 0.8000$	0.0680	0.0743	−0.2683	**11.355**
$\psi = 0.9530$	0.0689	0.0464	−0.2060	**0.004**
$\psi = 1.0000$	0.0831	0.0130	−0.1763	**13.003**
$\psi = 1.2000$	0.0710	0.0325	−0.0989	**11.715**
$\psi = 1.4000$	0.0769	0.0391	0.0127	**37.319**
$\psi = 0.9530$				
$\kappa = 0.0700$	0.0274	0.0249	−0.1662	**61.8962**
$\kappa = 0.1000$	0.0503	0.0111	−0.2064	**27.1192**
$\kappa = 0.1300$	0.0624	0.0278	−0.2237	**5.432**
$\kappa = 0.1443$	0.0689	0.0464	−0.2060	**0.004**
$\kappa = 0.1480$	0.0257	0.0995	−0.3345	**67.093**
$\kappa = 0.1500$	−0.0049	0.1402	−0.4144	**196.491**

regression that we match. Identification requires that the loss function is well behaved and varies systematically with the structural parameter estimates. In the top panel of table 10, we fix the value for κ at its estimated value and vary the adjustment cost coefficient, while in the bottom panel, we fix ψ at its estimated value and vary κ. In both cases, substantial deviations from the estimated parameter values imply large increases in the loss function. Importantly, the model implies a strong relationship between the severity of financial frictions (higher κ) and the response of investment to the balance sheet variable.

5.5 Model Simulations

We first consider the effect of the financial crisis – the exchange rate devaluation combined with the fall in domestic GDP and the rise in the country risk premium – on firm-level investment. We then consider the aggregate implications of the financial crisis given our structural estimates.

Figure 3 plots the effect of the devaluation combined with rising interest rates and falling output on investment for firms whose export share is at the first quartile of the distribution and whose foreign-denominated debt share is at the third quartile of the distribution. Results are reported for varying degrees of balance sheet exposure as measured by firm leverage ratios.

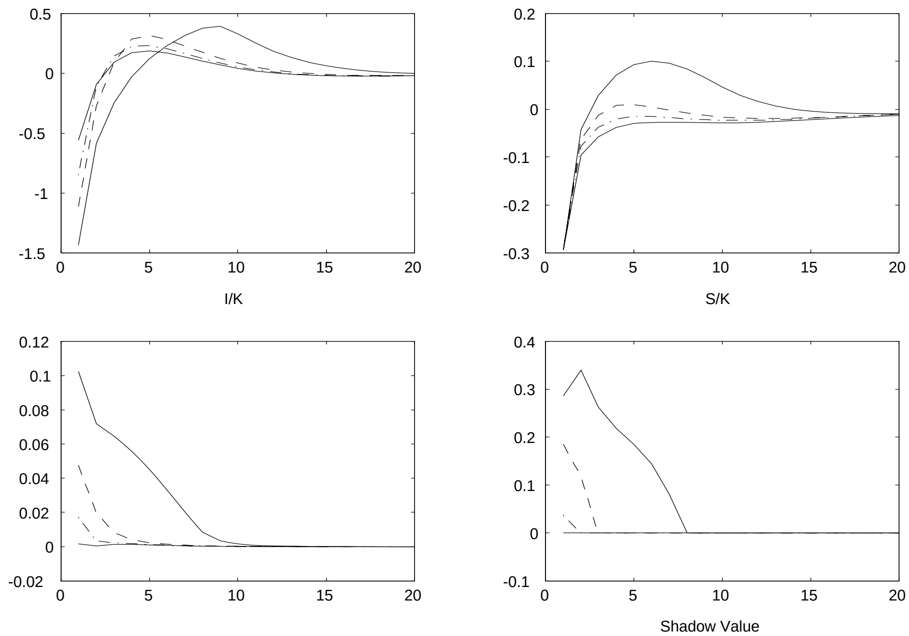

Figure 3: Investment Response: Low Export, High Foreign Debt.

The leverage ratios vary from the highest quartile (solid line) to the lowest quartile (dotted line).

For firms with low exports and high foreign-denominated debt ratios, the devaluation combined with the macroeconomic shocks implies a reduction in sales and investment. It also implies an increase in the cost of external funds that varies between one to ten percentage points. The overall spread between Korean corporate bonds and government bonds rose by 9% during this period. Thus, model estimates imply relatively conservative movements in the premium on external finance. The rise in the cost of external funds is larger for firms with high leverage ratios. As a result, the investment rate is substantially lower for such firms.

Figure 4 plots the response of investment to the same experiment for firms with a high export share (fourth quartile) and a low foreign-denominated debt ratio (first quartile). Results are again reported for leverage ratios that vary from the first to the fourth quartile. For firms that export, the devaluation combined with macroeconomic shocks implies an increase in sales and investment. The external finance premium rises by 0 to 3 percentage points depending on their initial leverage position. Again, the firms with high leverage experience a greater increase in the premium and a lower rate of investment in the initial period relative to firms with low initial leverage.

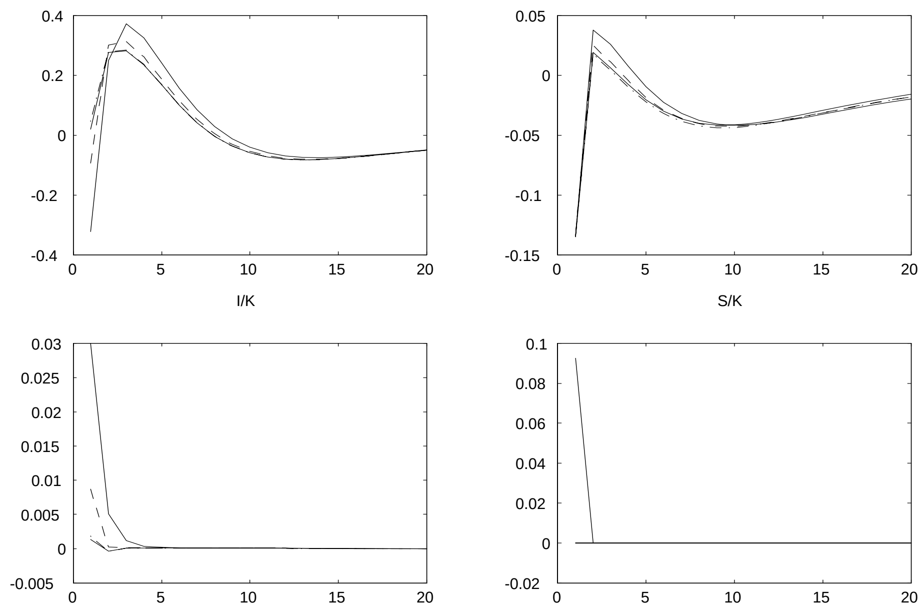

Figure 4: Investment Response: High Export, Low Foreign Debt.

We now consider the aggregate implications of our structural estimates. Figure 5 simulates the aggregate effect of the crisis, along with several counterfactual scenarios. To compute these simulations, we again feed in the macroeconomic shocks and compute the simulated path of investment for each of our firm types. We then compute the weighted average of this response, using the empirical distribution to compute the weights.[26] The resulting path for investment is plotted in the solid line in figure 5. We conduct three counterfactual experiments. First, we assume that foreign-denominated debt is zero (dot-dash line). Second, we assume that $\kappa = 0$, so that financial frictions play no role in the dynamics (dashed line). Third, we assume that all firms have a foreign-denominated debt ratio of fifty percent (dotted line). The foreign-denominated debt ratio in this last experiment are consistent with the ratios observed in Latin American economies during the 1980's and 1990s'.

Using the existing distribution of foreign-denominated debt, the simulation implies an 80% reduction in investment. This matches the observed drop in investment during the crisis.

Foreign-denominated debt plays only a small role in the model's aggregate investment dynamics – it accounts for less that 5% of the decline. This finding is consistent with the reduced form estimates which suggest that investment fell by one-half a percentage point owing

[26]Value weighted responses imply similar results across counterfactuals.

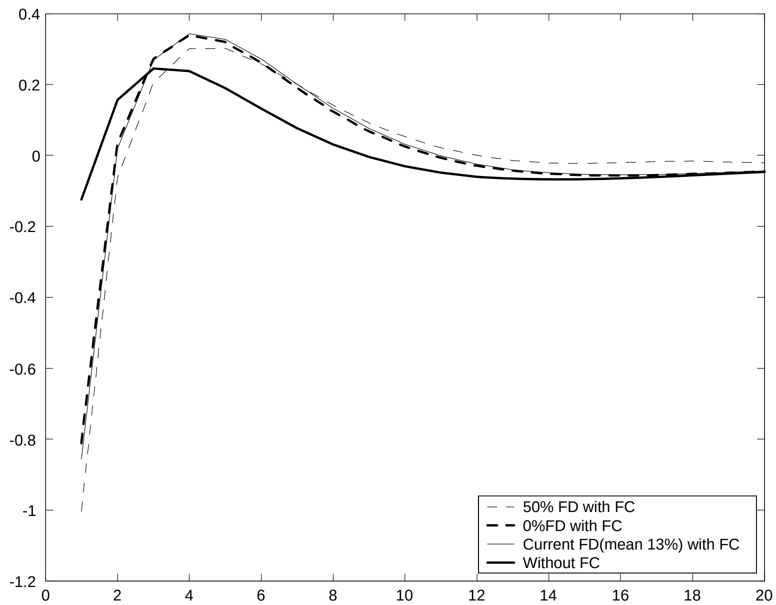

Figure 5: Aggregate Investment with Counterfactuals, $\rho_e = 0.85$.

to the presence of foreign-denominated debt. Our counterfactual simulation also considers the effect of a foreign-denominated debt ratio that is much higher than what we observed in the data. The average value in the data is 14%. The dashed line reports the effect of the devaluation under the assumption that all firms have a 50% foreign-denominated debt ratio. This counterfactual implies an additional 18% decline in investment. This is somewhat higher than the effect that one would compute using the reduced-form estimates which imply a 13% decline in investment owing to such an increase in the foreign-denominated debt ratio.[27] Intuitively, the model is non-linear in the financial mechanism, at higher levels of foreign-denominated debt, more firms are pushed into a region where the dividend constraint binds following the contraction. For such firms, the response of investment is particularly large.

Although foreign-denominated debt does not play an important role in determining the aggregate investment response, financial frictions are clearly an important determinant of investment dynamics. The increase in the domestic interest rate combined with the 20% reduction in demand cause a contraction in internal funds and therefore an increase in the

[27] The reduced form effect of increasing the foreign-denominated debt ratio is computing $\left[\hat{\beta}\ \Omega(e, e_{-1}; \omega')(b/a) - \hat{\beta}\ \Omega(e, e_{-1}; \omega)(b/a)\right]$ where $\Omega(e, e_{-1}; \omega) = (1 - \omega) + \omega e/e_{-1}$. Assuming that $\hat{\beta} = 0.21$, $b/a = 0.4$,and $e/e_{-1} = 1.7$, and letting $\omega' = 0.5$ and $\omega = 0.14$ we obtain an estimated response to the investment rate of 0.02. At a mean investment rate of 16%, this implies a 13% decline in investment.

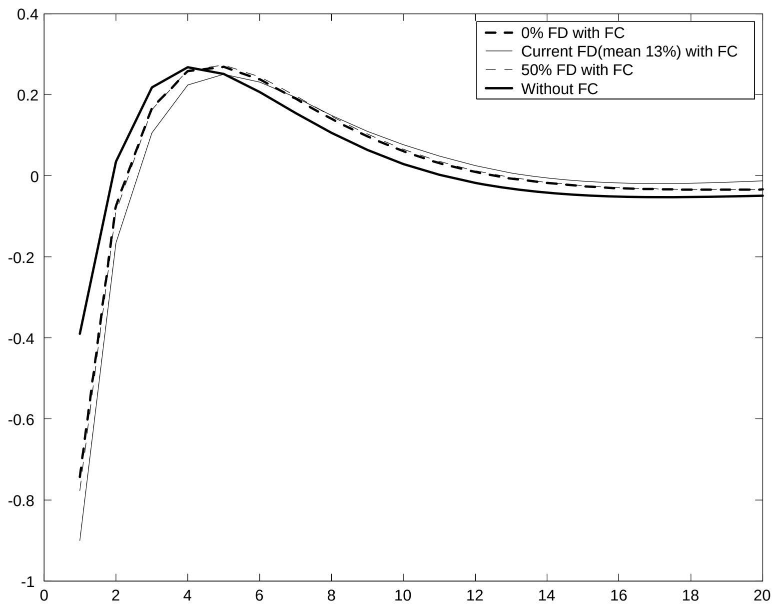

Figure 6: Aggregate Investment with Counterfactuals, $\rho_e = 0.90$.

premium on external funds. Without this financial mechanism, the negative consequences of the crisis are offset by the positive effect of the devaluation working through the competitiveness channel, and the anticipated drop in future interest rates under UIP, which in the case of $\rho = 0.85$ is substantial. As a result, with financial frictions, the model predicts an 80% fall in investment, whereas, absent financial frictions, the model predicts a 20% fall in investment.

Figure 6 considers the same experiment but increases the persistence of the exchange rate to $\rho = 0.9$. With greater exchange rate persistence, firms no longer expect future interest rates to revert as quickly under UIP. For unconstrained firms, the initial response of investment depends strongly on the future path of interest rates. As a result, the main difference between the $\rho = 0.85$ and $\rho = 0.90$ case is that the aggregate investment path for the case $\kappa = 0$ (no financial constraints) is lower when the exchange rate is perceived to be more persistent. Even in this case however, financial frictions still account for 50% of the overall drop in investment during the crisis period.

6 Conclusion:

This paper studies the effect of financial factors working through the balance sheet on investment spending during the Korean financial crisis. Our identification strategy combines reduced-form regression analysis and structural econometric estimation. We exploit firm-level heterogeneity in foreign-denominated debt ratios to identify shocks to the balance sheet that are distinct from shocks to fundamentals. By allowing for both adjustment costs and financial frictions, the structural model successfully replicates the reduced-form investment regression. The structural parameter estimates imply that the presence of foreign-denominated debt exerted a strong influence on investment at the micro-level.

Our structural parameter estimates allow us to conduct counterfactual exercises. These exercises imply that foreign-denominated debt plays an important role in explaining heterogeneous outcomes across firms. The overall effect of foreign-denominated debt was negligible during the crisis period however – accounting for at most one half a percent drop in aggregate investment. This finding is primarily due to the fact that the foreign-denominated debt ratio of the average Korean firm is relative small. Increasing the foreign-denominated debt ratio to 50% would lead to an additional 18% fall in investment. This result suggests that investment may indeed be sensitive to the presence of foreign-denominated debt in countries where the foreign-denominated debt ratios are sufficiently large.

Although foreign-denominated debt does not play an important role in investment dynamics during this time period, our structural estimates imply that financial frictions account for a large fraction of the investment decline. The rise in domestic interest rates combined with the fall in domestic GDP caused a deterioration of corporate balance sheets and a rise in the cost of external finance. According to our estimates, these factors can account for one half to three-fourths of the overall investment decline during the Korean financial crisis.

References:

Abel, Andrew and Janice Eberly (2003), "Q without adjustment costs and cash flow effects without financing constraints", University of Pennsylvania.

Eberly, Janice, Rebelo, Sergio, and Nicolas Vincent (2006), "Investment, Cash Flow and Value: A Neoclassical Benchmark", Northwestern University.

Aghion, Philippe, Philippe Bachetta and Abijit Banerjee (2000), "A simple model of monetary policy and currency crises", European Economic Review 44, 728-738.

Baltagi, Badi and Pinx X. Wu (1999), "Unequally spaced panel data with AR(1) disturbances", Econometric Journal 15, 814:823.

Bayraktar, Nihal, Plutarchos Sakellaris and Philip Vermeulen (2005), "Real versus financial frictions to capital investment", European Central Bank Working Paper No. 566.

Blanchard, Olivier Jean, Florencio Lopez-de-Silanes and Andrei Schleifer (1994), "What do firms do with cash windfalls", Journal of Financial Economics 36, 337-360.

Bleakley, Hoyt and Kevin Cowan (2002), "Corporate dollar debt and depreciations: much ado about nothing?", Working Papers 02-5, Federal Reserve Bank of Boston.

Cepedes, Luis Philippe., Roberto Chang and Andres Velasco (2000), "Balance sheets and exchange rate policy", NBER WP # 7840.

Christiano, Lawrence and Jonas Fisher (2000), "Algorithms for solving dynamic models with occasionally binding constraints", Journal of Economic Dynamics and Control 24, 1179-1232.

Cooper, Russell and Joao Ejarque (2003), "Financial frictions and investment: requiem in q", Review of Economic Dynamics 6, 710-728.

Den Haan, WJ and A. Marcet (1990), "Solving the Stochastic Growth Model by Parameterizing Expectations", Journal of Business and Economic Statistics 8, 31-34.

Desai, Mihir A., C. Fritz Foley, and Kristin J. Forbes (2007), "Financial Constraints and Growth: Multinational and Local Firm Responses to Currency Crises", Journal of International Economics, forthcoming.

Fazzari, Steven R., Glenn Hubbard and Bruce Petersen (1988), Financing constraints and corporate investment, Brookings Papers on Economic Activity 1, 144-195.

Gertler, Mark, Simon Gilchrist and Fabio M. Natalucci (2007), "External constraints on monetary policy and the financial accelerator", Journal of Money, Credit and Banking, forthcoming.

Gilchrist, Simon and Charles P. Himmelberg (1995), "The Role of Cash Flow in Reduced-Form Investment Equations" Journal of Monetary Economics Vol. 36.

Gilchrist, Simon and Charles P. Himmelberg (1998), "Investment, Fundamentals and Finance", in Ben Bernanke and Julio Rotemberg (eds.), The NBER Macroeconomics Annual.

Gomes, Joao F. (2001), "Financing Investment", American Economic Review 91, 1263-1285.

Gourieroux, Christian, and Alain Monfort (1996), Simulation Based Econometric Methods (Oxford University Press, Oxford, U.K.).

Frankel, Jeffrey A. (2003), "Experience of and lessons from exchange rate regimes in emerging economies", NBER WP # 10032.

Hennessy, Christopher A. and Toni M. Whited (2006), "How costly is external financing? Evidence from a Structural Estimation", Univ. of Wisconsin.

Hoshi, T. A., Kashyap and D. Scharfstein (1991), "Corporate structure, liquidity and investment: evidence from Japanese industrial groups", Quarterly Journal of Economics 106, 33-60.

Hubbard, R. Glenn (1998), "Capital market imperfections and investment", Journal of Economic Literature 36, 193-225.

Kim, Young-Sik and Park, Woo-Young, 2000, "Input-Output Structure and the Real Business Cycle of Korean Economy", Kyunghee University WP.

Judd, Kenneth (1992), "Projection Methods for Solving Aggregate Growth Models," Journal of Economic Theory 58, 410-425.

Kaplan, Steven N. and Luigi, Zingales (1997), "Do investment-cash flow sensitivities provide useful measures of financing constraints?", Quarterly Journal of Economics, February.

Krueger, Anne and Jae Yoo (2001), "Chaebol Capitalism and the currency-financial crisis in Korea", NBER Conference on Currency Crises Prevention.

Lamont, Owen (1997), “Cash flow and investment: evidence from internal capital markets”, Journal of Finance 52, 83-109.

Park, Hyung-Soo, (1999), "Korean Real Business Cycle", BOK WP

Pratap, Sangeeta and Silvio Rendon (2003), “Firm investment in imperfect capital markets: a structural estimation”, Review of Economics and Dynamics 63, 513-545.

Pratap, Sangeeta and Carlos Urrutia (2003), “Firm dynamics, investment and debt portfolio: Balance sheet effects of the Mexican crisis of 1994”, ITAM Working Paper.

Stein, Jeremy (2003), “Agency, Information and Corporate Investment” in G.M. Constantinides, M. Harris and R. Stulz, Handbook of the Economics of Finance, Elsevier Science.

Wright and Williams 1982, “The economic role of commodity storage”, Economic Journal, 596-614.

Appendix A: Indeterminancy of foreign-denominated debt ratio under UIP

We show that the foreign debt ratio is indeterminate if the agency cost is independent of the foreign debt ratio. To see the effects on the firm value of changing foreign debt ratio, consider a situation where the firm is allowed to readjust the foreign debt ratio each time period. In this case, the dynamic programming problem of the firm is given by

$$v(k,b,\omega,\mathbf{z},\mathbf{z}_{-1}) = \max_{k',b',\omega',d} \left\{ (1+\lambda)d + \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}',\mathbf{z}) v(k,b,\omega,\mathbf{z}',\mathbf{z}) Q(\mathbf{z},d\mathbf{z}') \right\}$$

where

$$d = \left(\sum_{i=d,f} \phi_i(\mathbf{z})\, k^{\gamma_i} - \sigma \right) - i - c(i,k) - \left(\frac{e}{e_{-1}} \right) R_f \omega b' - R_d(1-\omega)b + b'$$

and $b' \equiv b'_d + eb'_f = [(1-\omega') + \omega']b'$. Notice that the foreign debt is not parametrized any more.

Using the envelope theorem, the effects on the firm value of readjusting foreign debt ratio is given by

$$\begin{aligned} & \int_{\mathbf{z}'} \Lambda(\mathbf{z}',\mathbf{z}) \frac{\partial}{\partial \omega'} v(k',b',\omega',\mathbf{z}',\mathbf{z}) Q(\mathbf{z},d\mathbf{z}') \\ = & -(1+\eta')\, b' \int_{\mathbf{z}'} \Lambda(\mathbf{z}',\mathbf{z})\,(1+\lambda') \left[\left(\frac{e'}{e} \right) (1+r'_f) - (1+r'_d) \right] Q(\mathbf{z},d\mathbf{z}') \end{aligned}$$

Indeterminacy requires that the above expression is identically zero regardless of the foreign debt ratio. Since $(1+\eta')\, b' > 0$, this implies that

$$0 = \int_{\mathbf{z}'} \left\{ (1+\lambda')\, \Lambda(\mathbf{z}',\mathbf{z}) \left[\left(\frac{e'}{e} \right) (1+r'_f) - (1+r'_d) \right] \right\} Q(\mathbf{z},d\mathbf{z}')$$

Notice that the above expression is composed of two terms: the first term is the shadow value of the internal fund tomorrow. The second term can be considered as the shock to the UIP condition, i.e., news to the foreign exchange market. To see this last aspect, we can rewrite the bracketed term as

$$(1+r'_f) \left[\left(\frac{e'}{e} \right) - E\left(\frac{e'}{e} | \mathbf{z} \right) \right]$$

Therefore, the indeterminacy requires that the product of shadow value of the internal fund tomorrow and unanticipated news to the foreign exchange market should be expected to be zero once the current information set is controlled.

To show that this is the case indeed, we consider the original form of the dynamic programing in which the firm chooses the domestic debt and the foreign debt separately each time period, i.e.,

$$v(k, b_d, b_f, \mathbf{z}, \mathbf{z}_{-1}) = \max_{k', b'_d, b'_f, d} \left\{ (1+\lambda)\, d + \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) v(k', b'_d, b'_f, \mathbf{z}', \mathbf{z}) Q(\mathbf{z}, d\mathbf{z}') \right\}$$

where

$$d = \left(\sum_{i=d,f} \phi_i(\mathbf{z})\, k^{\gamma_i} - \sigma \right) - i - c(i,k) - R_d b_d - e R_f b_f + b'_d + e b'_f$$

The FOCs for domestic and foreign debts are given by

$$1 + \lambda = \mu \int_{\mathbf{z}'} \left[\Lambda(\mathbf{z}', \mathbf{z}) (1+\lambda') \left(R'_d + \frac{\partial R'_d}{\partial b'_d} b'_d + \frac{\partial R'_f}{\partial b'_d} e' b'_f \right) \right] Q(\mathbf{z}, d\mathbf{z}')$$

and

$$1 + \lambda = \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) (1+\lambda') \left[\left(\frac{1}{e}\right) \frac{\partial R'_d}{\partial b'_f} b'_d + \left(\frac{e'}{e}\right) R'_f + \frac{\partial R'_f}{\partial b'_f} \left(\frac{e'}{e}\right) b'_f \right] Q(\mathbf{z}, d\mathbf{z}')$$

Note that since $b' = b'_d + e b'_f$, we can write $\frac{\partial R'_d}{\partial b'_d} = \frac{\partial R'_d}{\partial b'}$, $\frac{\partial R'_f}{\partial b'_d} = \frac{\partial R'_f}{\partial b'}$, $\frac{\partial R'_d}{\partial b'_f} = \frac{\partial R'_d}{\partial b'} e$ and $\frac{\partial R'_f}{\partial b'_f} = \frac{\partial R'_f}{\partial b'} e$. Substituting these expression in the FOCs, we have

$$1 + \lambda = \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) (1+\lambda') \left(R'_d + \frac{\partial R'_d}{\partial b'} b'_d + \frac{\partial R'_f}{\partial b'} e' b'_f \right) |\mathbf{z} Q(\mathbf{z}, d\mathbf{z}')$$

and

$$1 + \lambda = \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) (1+\lambda') \left[\frac{\partial R'_d}{\partial b'} b'_d + \left(\frac{e'}{e}\right) R'_f + \frac{\partial R'_f}{\partial b'} e' b'_f \right] Q(\mathbf{z}, d\mathbf{z}')$$

Subtracting the first from the second results in

$$\begin{aligned} 0 &= \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) (1+\lambda') \left[\left(\frac{e'}{e} \right) R'_f - R'_d \right] Q(\mathbf{z}, d\mathbf{z}') \\ &= \mu (1+\eta') \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) (1+\lambda') \left[\left(\frac{e'}{e} \right) (1+r'_f) - (1+r'_d) \right] Q(\mathbf{z}, d\mathbf{z}') \end{aligned}$$

Therefore, we can see that

$$\int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) \frac{\partial}{\partial \omega'} v(k', b', \omega', \mathbf{z}', \mathbf{z}) Q(\mathbf{z}, d\mathbf{z}') = 0$$

identically regardless of the foreign debt ratio.

In case where the agency cost is affected by the foreign debt ratio, the effects on the firm value is slightly modified into

$$\begin{aligned} & \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) \frac{\partial}{\partial \omega'} v(k', b', \omega', \mathbf{z}', \mathbf{z}) Q(\mathbf{z}, d\mathbf{z}') \\ = & -\left[(1+\eta') b' + \frac{\partial \eta'}{\partial \omega'} \omega b' \right] \\ & \times \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) (1+\lambda') \left[\left(\frac{e'}{e} \right) (1+r'_f) - (1+r'_d) \right] Q(\mathbf{z}, d\mathbf{z}') \\ & - \frac{\partial \eta'}{\partial \omega'} b' \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) (1+\lambda') (1+r'_d) Q(\mathbf{z}, d\mathbf{z}') \end{aligned}$$

By the same logic, we can see that the first term of the above expression is identically zero. However, the second term is not zero identically by the assumption.

If the agency cost is monotonic in the foreign debt ratio, the optimal foreign debt ratio will have a boundary solution, i.e., either 0 or 1. If the agency cost is not monotonic, then the optimal policy may have inner solutions.

Appendix B: Derivation of the Profit Function.

The production technology is specified as

$$y_{j,t} = \begin{bmatrix} y_{d,j,t} \\ y_{f,j,t} \end{bmatrix} = z_{j,t} k_{t,j}^{\nu \alpha} \begin{bmatrix} n_{d,j,t}^{\nu(1-\alpha)} m_{d,j,t}^{1-\nu} \\ n_{f,j,t}^{\nu(1-\alpha)} m_{f,j,t}^{1-\nu} \end{bmatrix}$$

The demand function in each market is given by

$$y_{i,j,t} = \theta_{i,j} p_{i,j,t}^{-\epsilon_i} Z_{i,t} \quad \text{for } i = d, f$$

The profit function is then defined by

$$\begin{aligned}\pi_{j,t} &= p_{d,j,t} y_{d,j,t} + e_t p_{f,j,t} y_{f,j,t} - \sigma \\ &\quad - w_{n,t}(n_{d,j,t} + n_{f,j,t}) - e_t w_{m,t} \left(m_{d,j,t} + m_{f,j,t}\right)\end{aligned}$$

Using the definition of market demands, the profit can be rewritten as

$$\begin{aligned}\pi_{j,t} &= (\theta_{d,j} Z_{d,t})^{1-\chi_d} y_{d,j,t}^{\chi_d} + e_t (\theta_{f,j} Z_{f,t})^{1-\chi_f} y_{f,j,t}^{\chi_f} - \sigma \\ &\quad - w_{n,t}(n_{d,j,t} + n_{f,j,t}) - e_t w_{m,t} \left(m_{d,j,t} + m_{f,j,t}\right)\end{aligned}$$

where $\chi_i \equiv (\epsilon_i - 1)/\epsilon_i$ for $i = d, f$. Static optimization with respect to variable inputs, $m_{i,j,t}$ and $n_{i,j,t}$ for $i = d, f$ leads to the following conditional variable input demand functions,

$$\begin{aligned}n_{i,j,t} &= \nu(1-\alpha)\chi_i \frac{s_{i,j,t}}{w_{n,t}} \\ m_{i,j,t} &= (1-\nu)\chi_i \frac{s_{i,j,t}}{e_t w_{m,t}}\end{aligned}$$

where

$$s_{i,j,t} \equiv e_t^{1(i=f)} (\theta_{i,j} Z_{i,t})^{1-\chi_i} z_{j,t}^{\chi_i} k_{t,j}^{\nu\alpha\chi_i} n_{i,j,t}^{\nu(1-\alpha)\chi_i} m_{i,j,t}^{(1-\nu)\chi_i}$$

the sales of the firm j in market i at time t. $1(i = f)$ is an indicator function which takes 1 when $i = f$ and 0 otherwise. Thus input demands are proportional to sales in each market.

Subsitituting the conditional input demand functions in the profit function, we have

$$\pi_{j,t} = \sum_{i=d,f} \Gamma_i s_{i,j,t} - \sigma$$

where

$$\Gamma_i = 1 - \chi_i(1 - \nu\alpha)$$

In case of perfect competition, $\chi_i = 1$ and $\Gamma_i = \nu\alpha$, which is the capital share in the production function.

To obtain the closed form profit function, subsititute the conditional demand functions in the sales functions:

$$s_{i,j,t} = \theta_{i,j}^{(1-\chi_i)/\Gamma_i} \Xi_{i,t} z_{j,t}^{\chi_i/\Gamma_i} e_t^{(1(i=f)-(1-\nu)\chi_i)/\Gamma_i} k_{j,t}^{\nu\alpha\chi_i/\Gamma_i}$$

where

$$\Xi_{i,t} \equiv \Psi_i \left[\frac{Z_{i,t}^{(1-\chi_i)}}{w_{n,t}^{\nu(1-\alpha)\chi_i} w_{m,t}^{(1-\nu)\chi_i}} \right]^{1/\Gamma_i}$$

and

$$\Psi_i \equiv \left[(\nu(1-\alpha)\chi_i)^{\nu(1-\alpha)\chi_i} \left((1-\nu)\chi_i \right)^{(1-\nu)\chi_i} \right]^{1/\Gamma_i}.$$

The term $\Xi_{i,t}$ represents the common (across firms) aggregated component of profits in the domestic and foreign market respectively. We assume that the common aggregate component can be represented as a log-linear combination of the relevant aggregate economic activity variable in each market, $y_{i,t}$ and the exchange rate

$$\log \Xi_{i,t} = \psi_i + \alpha_i \log y_{i,t} + \beta_i e_t$$

in which case the log sales equation becomes:

$$\log s_{i,j,t} = \log \psi_i + \varsigma_i \ln \theta_{i,j} + \xi_i \ln e_t + \gamma_i \ln k_{j,t} + \alpha_i \log y_{i,t} + v_{i,t},$$

where the key elasticity in the sales equation is

$$\gamma_i = \frac{\nu\alpha\chi_i}{1 - \chi_i(1-\nu\alpha)}.$$

Given an estimate for the production share of capital, $v\alpha$, and an estimate of γ_i, we can obtain an estimate of the relevant mark-up in each market $1/\chi_i$ along with the weights

$$\Gamma_i = 1 - \chi_i(1-\nu\alpha)$$

necessary to construct the profit function.

The coefficient on the exchange rate in the sales equation satisfies $\xi_i = \beta_i + \frac{1(i=f)-\chi_i(1-\nu)}{1-\chi_i(1-\nu\alpha)}$. If exchange rates only influence sales through their direct effect on profits and have no indirect effect through their impact on the macro factor $\Xi_{i,t}$, we expect $\xi_i = \frac{1(i=f)-\chi_i(1-\nu)}{1-\chi_i(1-\nu\alpha)}$ which implies a restriction on the sales equation coefficients. If the macro factor depends on the exchange rate, this restriction does not hold however. Alternatively, imperfect pass through of exchange rates to prices could also lead to deviations between the estimated parameter ξ_i and the implied structural parameter $\frac{1(i=f)-\chi_i(1-\nu)}{1-\chi_i(1-\nu\alpha)}$. In practice, our estimates imply substantial deviations from the restriction $\xi_i = \frac{1(i=f)-\chi_i(1-\nu)}{1-\chi_i(1-\nu\alpha)}$.

The market demand has a constant proportionality term, $\theta_{i,j}$. The constant term may be interpreted as the steady state size of the market for the product j in market i because the steady state implies $y_{i,j,ss} = \theta_{i,j}$ (for a symmetric equilibrium). Note that the closed-form sales functions also have proportionality factors which are determined by $\theta_{i,j}$. The sales functions are composed of a time-invariant component ($\theta_{i,j}^{(1-\chi_i)/\Gamma_i}$) and a time-varying component $\tilde{s}_{i,j,t}$ where $\tilde{s}_{i,j,t} \equiv \Xi_{i,t} z_{j,t}^{\chi_i/\Gamma_i} e_t^{(1(i=f)-(1-\nu)\chi_i)/\Gamma_i} k_{j,t}^{\nu\alpha\chi_i/\Gamma_i}$. To simplify the functional form of the sales equations, we rewrite them as

$$\pi_{j,t} = \eta_j \sum_{i=d,f} \Gamma_i \eta_{i,j} \tilde{s}_{i,j,t} - \sigma_j$$

where $\eta_j \equiv \sum_{i=d,f} \theta_{i,j}^{(1-\chi_i)/\Gamma_i}$ and $\eta_{i,j} \equiv \theta_{i,j}^{(1-\chi_i)/\Gamma_i} / \sum_{i=d,f} \theta_{i,j}^{(1-\chi_i)/\Gamma_i}$. We then normalize the term η_j to unity and assume that σ_j is also normalized by the same factor so that the fixed-cost to sales ratio is invariant across firms. We then approximate $\eta_{i,j}$ using the export-sales ratio of a firm combined with the estimated Γ_i so that $\tilde{\eta}_{i,j} \equiv \Gamma_i(s_{i,j,ss}/\sum_{i=d,f} s_{i,j,ss})$. This approximation is exact (i.e. $\tilde{\eta}_{ij} = \eta_{ij}$)when all functional assumptions are satisfied and the firm has the same market powers in all markets. Profits can now be approximated as a weighted average of the time-varying component of sales, $\tilde{s}_{i,j,t}$:

$$\begin{aligned} \pi_{j,t} &= \sum_{i=d,f} \tilde{\eta}_{i,j} \tilde{s}_{i,j,t} - \sigma_j \\ &= \sum_{i=d,f} \phi_{i,j}(\mathbf{z}_t) k_{j,t}^{\gamma_i} - \sigma_j \end{aligned}$$

where $\phi_{i,j}(\mathbf{z}_t) \equiv \tilde{\eta}_{i,j}[\Xi_{i,t} z_{j,t}^{\chi_i/\Gamma_i} e_t^{(1(i=f)-(1-\nu)\chi_i)/\Gamma_i}]$ and $\gamma_i \equiv \nu\alpha\chi_i/\Gamma_i$. Therefore, given the weights $\tilde{\eta}_{i,j}$, the profit function may be computed directly from the sales equations. To determine the weight, we compute $\Gamma_i = 1 - \chi_i(1 - \nu\alpha)$ using the estimated χ_i of the gross profit function and set $v\alpha = 0.225$. We calibrate the ratio $s_{i,j,ss} / \sum_{i=d,f} s_{i,j,ss}$ using the pre-crisis mean export-sales ratio for each firm.

Appendix C: Computational Method.

We transform the FOC into forms more convenient for computation in the following way

$$\frac{1+\lambda}{b'} = \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) \frac{1+\lambda'}{b''} \frac{b''}{b'} \Omega(e', e; \omega) \left(1 + r_d'\right) \left(1 + \eta' + \frac{\partial \eta'}{\partial b'} b'\right) Q(\mathbf{z}, d\mathbf{z}')$$

and

$$\left(1 + \frac{\partial c}{\partial i}(i, k)\right) \frac{1+\lambda}{b'} = \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}', \mathbf{z}) \frac{1+\lambda'}{b''} \frac{b''}{b'} \left[\frac{\partial d'}{\partial k'} + (1-\delta)\left(1 + \frac{\partial c}{\partial i'}(i', k')\right)\right] Q(\mathbf{z}, d\mathbf{z}')$$

where

$$\frac{\partial d'}{\partial k'} = \sum_{i=d,f} \alpha_j \phi_i(\mathbf{z}) k^{\alpha_j - 1} - \frac{\partial c}{\partial k'}(i', k') - \Omega(e', e; \omega)\left(1 + r_d'\right) \frac{\partial \eta'}{\partial k'} b'$$

We adopt a version of Chebyshev projection method(Judd(1992) and Christiano and Fisher(2000)). We approximate the conditional expectations of the model using orthogonal polynomials, i.e.,

$$\exp(h_b(\mathbf{u})) \simeq \int_{\mathbf{z}'} \left[\Lambda(\mathbf{z}', \mathbf{z}) \frac{1+\lambda'}{b''} \frac{b''}{b'} \Omega(e', e; \omega)\left(1 + r_d'\right)\left(1 + \eta' + \frac{\partial \eta'}{\partial b'} b'\right)\right] Q(\mathbf{z}, d\mathbf{z}')$$

and

$$\exp(h_k(\mathbf{u})) \simeq \int_{\mathbf{z}'} \left\{\Lambda(\mathbf{z}', \mathbf{z}) \frac{1+\lambda'}{b''} \frac{b''}{b'} \left[\frac{\partial d'}{\partial k'} + (1-\delta)\left(1 + \frac{\partial c}{\partial i'}(i', k')\right)\right]\right\} Q(\mathbf{z}, d\mathbf{z}')$$

where $h_b(\cdot)$ and $h_k(\cdot)$ are Chebyshev polynomials and u is a vector of logged state variables. The choice for the conditional expectations rather than the policy and the multiplier functions as the objects of approximation is due to relative smoothness of the conditional expectation functions(See Christiano and Fisher(2000)).

Assuming that the approximating functions are close enough to the actual conditional expectations, we can reconstruct the system of the equations using those approximating functions in the following way. Dividing the FOC for k' by the FOC for b', we can derive an expression for Tobin's q in terms of the approximating functions, i.e.,

$$q(i/k) \equiv 1 + \frac{\partial c}{\partial i}(i,k) \simeq \frac{\exp(h_k(\mathbf{u}))}{\exp(h_b(\mathbf{u}))}$$

Under the functional assumptions we adopt regarding the capital adjustment cost, Tobin's q is invertible for investment ratio and therefore

$$i(\mathbf{u}) \simeq q^{-1}\left[\frac{\exp(h_k(\mathbf{u}))}{\exp(h_b(\mathbf{u}))}\right]k$$

and

$$k'(\mathbf{u}) \simeq \left\{q^{-1}\left[\frac{\exp(h_k(\mathbf{u}))}{\exp(h_b(\mathbf{u}))}\right] + (1-\delta)\right\}k$$

The optimal debt policy can be computed using this investment policy. The debt policy is given by

$$b'(\mathbf{u}) \simeq \begin{cases} [\mu \exp(h_b(\mathbf{u}))]^{-1} & \text{if} \quad [\mu \exp(h_b(\mathbf{u}))]^{-1} \geq \bar{b}'(\mathbf{u}) \\ \bar{b}'(\mathbf{u}) & \text{if} \quad [\mu \exp(h_b(\mathbf{u}))]^{-1} < \bar{b}'(\mathbf{u}) \end{cases}$$

or more simply

$$b'(\mathbf{u}) \simeq \max\left\{\bar{b}'(\mathbf{u}), [\mu \exp(h_b(\mathbf{u}))]^{-1}\right\}$$

where $\bar{b}'(\mathbf{u})$ is the minimum level of debt finance satisfying the dividend constraint, i.e.,

$$\bar{b}'(\mathbf{u}) \equiv \sigma + i(\mathbf{u}) + c(i(\mathbf{u}), k) + \tilde{b} - \sum_{i=d,f} \phi_i(\mathbf{z})\, k^{\gamma_i}$$

Note that the max function in the debt policy is replaced with a smooth max function in the actual computation.[28] Since the investment and debt policies are constructed, the dividend

[28] For instance, $\max(x,0) \simeq x + \frac{1}{\alpha}\log(1+\exp(-x))$. The greater α results in the less smooth the approximation.

policy is simply given by the definition, i.e.,

$$d(\mathbf{u}) \simeq \left(\sum_{i=d,f} \phi_i(\mathbf{z})\, k^{\gamma_i} - \sigma\right) - i(\mathbf{u}) - c(i(\mathbf{u}), k) - \tilde{b} + b'(\mathbf{u})$$

Finally, the shadow value of the internal fund can be computed as

$$\lambda(\mathbf{u}) = \frac{b'(\mathbf{u})}{[\mu \exp(h_b(\mathbf{u}))]^{-1}} - 1$$

Notice that the shadow value is computed as the vertical distance between the constrained policy, $b'(\mathbf{u})$ and the unconstrained policy, $[(\beta\mu)\exp(h_b(\mathbf{u}))]^{-1}$. Since the constrained policy cannot be less than the unconstrained policy, the Lagrange multiplier cannot take a negative value regardless of correctness of the approximated functions.[29]

The recursive nature of the functional equations can be seen in the following.

$$\begin{aligned}\exp(h_b(\mathbf{u})) \simeq\ & \mu \int_{\mathbf{z}'} \Lambda(\mathbf{z}',\mathbf{z}) \exp(h_b(\mathbf{u}')) \frac{b''(\mathbf{u}')}{b'(\mathbf{u})} \\ & \times \Omega(e', e; \omega)(1 + r_d') \left(1 + \eta'(\mathbf{u}) + \frac{\partial \eta'}{\partial b'}(\mathbf{u}) b'(\mathbf{u})\right) Q(\mathbf{z}, d\mathbf{z}')\end{aligned}$$

[29] There could be other transformations to enable us to identify the policy variable. However, some of them do not satisfy the Kuhn-Tucker condition. For instance, a possible invertible form might be

$$b'(1+\lambda) = \mu E\left[\Lambda(\mathbf{z}',\mathbf{z}) b''(1+\lambda') \frac{b'}{b''} \Omega(e', e; \omega)(1 + r_d') \left(1 + \eta' + \frac{\partial \eta'}{\partial b'} b'\right) |\mathbf{z}\right]$$

In this case we parameterize the conditional exectation to get

$$b'(1+\lambda) = \mu \exp(h_b(\mathbf{u}))$$

If the constraint is nonbinding, the debt policy is given by $b'(u) = \beta\mu \exp(h_b(u))$. If the constraint is binding the policy is given by $b'(u) = \bar{b}'(u) \equiv \sigma k + i(u) + c(i(u), k) - \sum_i \phi_i(z) k^{\gamma_i}$. Finally the Lagrangian multiplier is calculated as

$$\lambda = \frac{\mu \exp(h_b(\mathbf{u}))}{b'(\mathbf{u})} - 1$$

If the constraint is nonbinding, this formula correctly gives $\lambda = 0$. However, if the constraint is binding where the unconstrained policy is lower than the constrained policy, this formula returns a negative value for the multiplier.

and

$$
\begin{aligned}
\exp(h_k(\mathbf{u})) \quad &\simeq \quad (\beta\mu)\int_{\mathbf{z}'}\Lambda(\mathbf{z}',\mathbf{z})\exp(h_b(\mathbf{u}'))\frac{b''(\mathbf{u}')}{b'(\mathbf{u})} \\
&\times\left[\frac{\partial}{\partial k'}d'(\mathbf{u}')+(1-\delta)\frac{\exp(h_k(\mathbf{u}'))}{\exp(h_b(\mathbf{u}'))}\right]Q(\mathbf{z},d\mathbf{z}')
\end{aligned}
$$

Since these are not contraction mappings, a nonlinear numerical equation solver must be adopted for the solutions. The integrations over the future uncertainties are replaced by Gauss-Hermite quadratures in the actual computations. When applying this solution method, we use a complete basis computed from a second-order chebyshev polynomials in each of the state variables. The resulting second-order approximation in logs is sufficiently non-linear to capture higher order effects when computing conditional expectations.

Appendix D: The Determination of Steady State and Fixed Cost

In the steady state, the Euler equations may be expressed as

$$
1=\mu\beta\left[(1+r_{d,ss})\left(1+\eta_{ss}+\frac{\partial\eta_{ss}}{\partial b_{ss}}b_{ss}\right)\right]
$$

and

$$
1=\mu\beta\left[\sum_{i=d,f}\alpha_j\phi_i\left(\mathbf{z}_{ss}\right)k_{ss}^{\alpha_j-1}-(1+r_{d,ss})\frac{\partial\eta_{ss}}{\partial k_{ss}}b_{ss}+(1-\delta)\right]
$$

where we use $\beta\Lambda(\mathbf{z}_{ss},\mathbf{z}_{ss})=\beta$, $\lambda'_{ss}=\lambda_{ss}$, $e_{ss}=1$ and $\Omega(e_{ss},e_{ss};\omega)=1$. Under the functional form assumptions, it is straight forward to show that η_{ss} and $\frac{\partial\eta_{ss}}{\partial b_{ss}}b_{ss}$ are solely determined by the steady-state leverage ratio, b_{ss}/k_{ss}. We assume that all firms have an identical long run leverage ratio which is calibrated from the post-crisis mean level of leverage in the data. An implicit assumption behind the use of post crisis mean is that the pre crisis mean level of leverage ratio was higher than the long run level. Although this assumption does not seriously affect the estimation results, it captures the realistic notion that firm balance sheets were extended at the onset of the crisis.

For a given parameter estimate κ, the first Euler equation may be used to back out the survival probability μ that is consistent with this long-run leverage ratio. The information of the leverage ratio and the survival probability can then be used in the second Euler equation to determine the steady-state level of capital. Note that the steady-state level of capital differs across heterogeneous firms to the extent that different firms have different steady-state export-sales ratios owing to the nonhomogeneous curvatures of the profit functions for domestic and foreign markets. Finally, the long-run leverage ratio and capital then determines the long-run level of debt.

Once the steady state capital stock is determined, the fixed cost can be calibrated from the data. In case we assume that the fixed cost is proportional to the steady state level of capital, we write $\sigma = c_F k_{ss}$ where c_F is the proportionality factor. The operating income to capital ratio in the data is given by $OIK = (P - F)/K = (S - C - F)/K$ where P, S, C, F and K are accounting data on profit, sales, cost, fixed cost(the item, Sales and General Management Cost) and capital. The same ratio in the model is given by

$$OIK = \sum_{i=d,f} \phi_i(\mathbf{z}_{ss}) k_{ss}^{\gamma_i - 1} - \sigma/k_{ss} = \sum_{i=d,f} \phi_i(\mathbf{z}_{ss}) k_{ss}^{\gamma_i - 1} - c_F$$

where $\phi_i(\mathbf{z}_{ss}) k_{ss}^{\gamma_i - 1} = \Gamma_i (s/k)_i$ are determined by the estimates of Γ_i in conjunction with the mean sales-to-capital ratio in each market. The proportionality factor c_F may then be computed as

$$c_F = \sum_{i=d,f} \phi_i(\mathbf{z}_{ss}) k_{ss}^{\gamma_i - 1} - OIK = \sum_{i=d,f} \phi_i(\mathbf{z}_{ss}) k_{ss}^{\gamma_i - 1} - 0.241$$

We use the pre-crisis mean of the operating income ratio(0.241) to determine c_F.

If we assume that the fixed cost is proportional to the steady-state sales, we write $\sigma = c_F \sum_{i=d,f} \phi_i(\mathbf{z}_{ss}) k_{ss}^{\gamma_i}$. The operating income ratio in the model is then given by

$$OIK = (1 - c_F) \sum_{i=d,f} \phi_i(\mathbf{z}_{ss}) k_{ss}^{\gamma_i - 1}$$

We can then determine the fixed cost parameter c_F as

$$c_F = 1 - \frac{OIK}{\sum_{i=d,f} \phi_i(\mathbf{z}_{ss}) k_{ss}^{\gamma_i - 1}} = 1 - \frac{0.241}{\sum_{i=d,f} \phi_i(\mathbf{z}_{ss}) k_{ss}^{\gamma_i - 1}}$$

These procedures provides identical results since they are just different methods of applying the same constraint.

Appendix E: Data Construction.

We construct standard ratios for investment and sales relative to capital. All variables

Table 11: Macroeconomic Variables

	1995	1996	1997	1998	1999	2000	2001	2002
Z	2.295	3.771	3.852	−7.268	−2.552	0.912	−0.149	1.753
e	0.760	0.743	0.799	1.3799	1.044	1.008	1.123	1.154
r_d	7.996	7.542	8.966	7.826	4.119	2.816	0.632	1.526
ξ	1.879	1.086	3.600	10.65	2.765	0.936	0.372	1.675

are deflated by the appropriate price indices. Investment spending is deflated by the capital goods price index from the producer price index; domestic sales, total debt and total assets are deflated by the producer price index for manufacturing; and foreign sales are deflated by the export price index. Investment data are constructed as the difference between the Increase in Tangible Asset and the Decrease in Tangible Asset variables from the Cash Flow Statement. All other variables in the regression are extracted from either the Balance Sheet or Income Statement.

The real capital stock data is constructed according to the perpetual inventory method, i.e.,

$$k_{j,t+1} = (1-\delta)k_{j,t} + \frac{I_{j,t}}{P_{k,t}} \tag{25}$$

where $I_{j,t}$ is nominal investment spending of firm j and $P_{k,t}$ is the capital goods price index. This way of constructing of the real capital stock requires an information for initial value, $k_{j,0} \equiv K_{j,0}/\tilde{P}_{k,0}$ where $\tilde{P}_{k,0}$ is the price index for installed capital at time 0. Since this price level is not available, we deflate the initial nominal capital stock by the capital price index, $P_{k,0}$. To exclude the influences of extreme observations, our sample is constructed using a cut-off rule which drops outliers defined as observations in the lowest and the highest 0.5% of the sample.

Table 11 reports the actual values of the macroeconomic variables that are used in the estimation and simulation of the structural model. These macroeconomic variables are computed on an annual basis. Table 11 reports the values for the demand shifter Z (HP-filtered real GDP), the real exchange rate e, the domestic real rate r_d, and the implied country-risk premium, ξ, which may be obtained backed out of the UIP condition given the data for the exchange rate and the domestic real rate. On an annual basis, the crisis (1998) generated a 10% drop in the demand factor Z, a 70% devaluation and a seven percentage point rise in the country-risk premium.