Investment During the Korean Financial Crisis: the Role of Foreign-Denominated Debt[1]

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Abstract

Without capital market imperfections, the capital structure of a firm, including the size, the maturity and the currency compostion 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 the effect of foreign denominated debt for investment spending may account for up to 50% of the drop in investment during the crisis period.

[1] **We thank seminar participants at the Board of Governors of the Federal Reserve, Brandeis University, the NBER Summer Institute, and and the University of Pennsylvania**

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. Since 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.

In this paper, we test for the existence of a finance channel in the propagation of the Korean financial crisis. In addition, we provide 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. Our results also imply that the presence of foreign-denominated debt explains 50% of the investment decline during the crisis period.

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

[2]Frankel (2003) provides a recent discussion of the relative benefits of fixed versus flexible exchange rates. Aghion et al (2001) and Cepedes et al (2002) consider these issues in the context of a small open economy framework with dollar denominated debt. Gertler et al (2003) also consider the role of foreign denominated debt in the context of a GE framework explicitly calibrated to the Korean experience.

[3]If the production of capital uses foreign investment goods, the devaluation may also affect investment by changing the price index of investment goods. As we show below, this appears not to have been the case in the Korean episode.

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.

The effect of the devaluation on investment through the balance sheet is theoretically ambiguous and depends on the extent to which a firm's debt is denominated in foreign currency and the extent to which a firm's earnings are export dependent. The devaluation raises the value of existing debt in direct proportion to the share of debt that is denominated in foreign. Thus, foreign-denominated debt will unambiguously depress investment in the presence of financial market imperfections. For a firm that exports however, the devaluation improves its ability to pay back its debt. We expect the investment spending of firms with low exports and high levels of foreign-denominated debt to be the most adversely affected by the devaluation.

Understanding the effect of foreign-denominated debt for investment spending requires 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. This data set is unique in a number of ways. It provides detailed firm-level data on non-financial variables such as sales, profits, investment and capital; it provides financial data such as debt and equity; it is comprehensive, covering all publicly-traded as well as many non-publicly-traded Korean firms over the period 1993-2002 and thus accounts for a large fraction of overall Korean business fixed investment spending; most 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. We view the exchange-rate crisis and ensuing devaluation as a natural experiment with which we 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. By controlling for foreign exports directly, we cleanly identify the effect of foreign-denominated debt on investment spending.

While such an analysis is informative, it does not provide a complete quantitative assessment. In the second part of the paper, we adopt a structural approach. We specify a dynamic optimization problem of a firm which produces for both domestic and foreign markets and both domestic and foreign-denominated debt. The firm operates under a set of financial and non-financial constraints. We use this dynamic program to estimate the structural relationship characterizing investment, profitability and financial conditions.

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 foreign denominated debt. Because these authors do not have separate information on the export status of firms, they are unable to separate balance-sheet effects from competitiveness effects however. Aguiar (2004) examines the investment behavior of Mexican firms during the 1994 pesos devaluation, and finds a negative effect of foreign-denominated debt that is distinct from the competitiveness effect. These papers adopt a reduced-form approach and therefore cannot formally quantify the effect that foreign-denominated debt exerts on investment. Pratap and Urrutia (2003). consider a structural model of investment with financial frictions which is calibrated to the Mexican firm-level data. This paper emphasizes the role that the devaluation played on the balance sheet during the Mexican currency crisis but makes no attempt at formal estimation.

Our paper is also related to the extensive literature on firm level investment and capital market imperfections.[4]. 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), Kashyap, Hoshi 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

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

(1997), Gomes (1999) Abel and Eberly (2002)).

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), Hennessy and Whited (2004)).

A major limitation of current structural estimates is the focus on a single shock which 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. For example, in such single shock environments, Cooper and Ejarque document that cash flow–investment correlations can be explained by non-constant returns to scale of the profit function without relying on capital market imperfections. In addition to focussing on a single shock environment, these models abstract from adjustment costs, so that absent capital market imperfections, capital accumulation is frictionless.[5] Because capital market imperfections limit investment spending, such estimation procedures may not be robust to the alternative hypothesis that capital accumulation responds to profits owing to sluggish adjustment on the real side. By combining real side frictions with financial frictions, and identification through balance sheet shocks we avoid such potential pitfalls.

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 foreign-denominated debt played in the propagation of the crisis.

[5] Cooper and Ejarque (2003) is a notable exception.

Figure 1: Investment, sales and debt during financial crisis.

2 Overview of Korean Financial Crisis

In this section, we provide 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] We plot the average ratios of investment, sales and debt relative to total assets. For comparison purposes, we also plot 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 sometime 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

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

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 shows a sharp increase at the onset of the crisis, reflecting the stress on balance sheets caused by the currency depreciation. Over time, debt falls relative to assets, returning to a level somewhat below its pre-crisis value.

To investigate how the investment rate differed in response based on the degree of a firm's foreign exchange rate exposure, we divide our sample into firms with high versus low levels of exports, and high versus low levels of foreign-denominated debt. To classify firms according to export status, we compute the pre-crisis average export to total sales ratio for each firm in our sample. We then 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. We again use the pre-crisis median value as our cutoff. The average investment rates for high versus low foreign-denominated debt and high versus low export firms are plotted in the upper two panels of figure 2. We also consider the four way interaction obtained by classifying firms according to the median categorization of both high versus low exports and high versus low foreign-denominated debt. These four way classifications are plotted in the lower 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-foreign debt. We find little difference in the investment rate of firms with high levels of exports relative to firms with low levels of exports. As we discuss further below, there is a positive correlation between foreign-debt exposure and foreign-sales exposure. Thus, 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

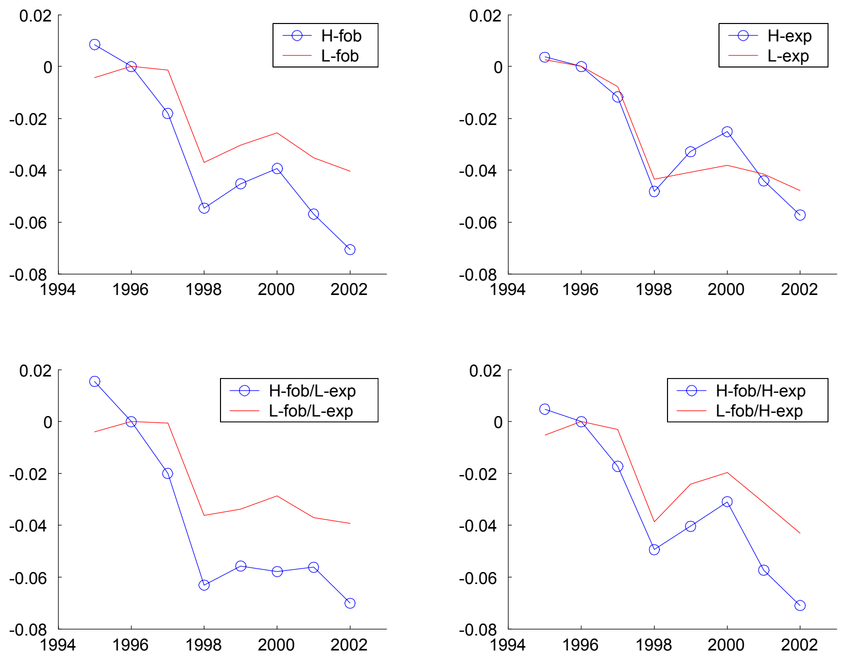

Figure 2: Investment rates.

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.

3 The Investment Model

In this section we present the structural model of investment that we estimate. 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 choose 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 finance:

$$v(k, b_d, b_f, \mathbf{z}) = \max_{k', b'_d, b'_f, d} \{d + \beta E[v(k', b'_d, b'_f, \mathbf{z}')|\mathbf{z}]\} \tag{1}$$

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

$$d = \left(\sum_{j=d,f} \pi_j(a, e, z) k^{\alpha_j} - \sigma \right) - i - c(i, k) + b'_d + eb'_f - R_d b_d - eR_f b_f \tag{2}$$

The profit function is the sum of domestic and foreign profits and fixed costs to production σ. Here α_d and α_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 $\pi_j(a, e, z)$ which in turn depend on e, the real exchange rate, a an aggregate shock and z an iid idiosyncratic shock.[7] We normalize the price of investment goods to unity.[8] We assume a convex constant-returns-to-scale capital adjustment cost, 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 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 $\mathbf{z}$ denotes the set of all relevant exogenous state variables.

[7] The underling assumptions regarding market structure and production technology can be found in section 5.

[8] Production in a typical small open economy is substantially dependent on imported capital goods. This suggests that the capital goods price $p^i(e_t)$ should be modeled as an explicit function of the exchange rate. In practice however, the exchange rate devaluation influenced domestic prices and foreign prices in such a way as to not have had a significant effect on the overall relative price of investment goods.

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

$$d \geq 0.$$

and let λ denote the Lagrangian multiplier associated with the constraint. We also assume that the total borrowing cost can be decomposed into a risk-free interest rate and an external-finance premium,

$$R_f = (1 + r_f)(1 + \eta) \tag{3}$$

and

$$R_d = (1 + r_d)(1 + \eta) \tag{4}$$

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. We thus follow Gomes(1999) and assume 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 debt allows us to simplify the model and eliminate one state variable. Let

$$\omega \equiv \frac{e_{-1} b_f}{b_d + e b_f} \tag{5}$$

the ratio of foreign debt to total debt in local currency units. Uncovered interest parity implies that the firm cares about the total debt obligation $b = b_d^{'} + e b_f^{'}$ 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 and 3 is then equivalent to the following program with smaller dimension

$$v(k, b, \mathbf{z};\omega) = \max_{k',b',d}\{(1+\lambda)\, d + \beta E\left[v(k', b', \mathbf{z}';\omega)|\mathbf{z}\right]\} \tag{6}$$

where the dividend is redefined as

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

Although we assume that the currency composition is 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 an exchange rate devaluation on the balance sheet.

[9] Strictly speaking, our formulation 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.

The dividend can now be simplified to

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

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]

We specify the external-finance premium parsimoniously as

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

where

$$x \equiv x\left(k, b, \mathbf{z}_{-1}\right) \tag{11}$$

As for the $x\left(\cdot\right)$ function, we choose a functional form with the following properties: i) an increase in capital reduces the external finance premium, ii) an increase in debt increases the

[10] It is worth emphasizing that only unanticipated movements in the exchange rate causes changes in the value of debt outstanding. If the devaluation is expected, the UIP conditions guarantee that such anticipated effects are already built into the relevant risk free rate.

premium, iii) any exogenous state variable that predicts an increase in profitability in the next period also reduces the external finance premium. A simple functional form satisfying these properties is

$$x\left(k, b, \mathbf{z}_{-1}\right)=\frac{b}{\sum_{j=d,f} \pi_j(\mathbf{z}_{-1}) k} \tag{12}$$

Under this specification, the external finance premium is a function of the state variables when issued, i.e.,

$$\eta^{\prime}=\eta\left(k^{\prime}, b^{\prime}, \mathbf{z}\right) \tag{13}$$

Note that the function η is strictly convex with curvature $x\eta''(x)/\eta'(x)=x$. Thus the slope of the premium rises more rapidly as leverage increases.

Given the premium on external funds, firms 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. If $\mu=1$ and $\beta^{-1}-1$ is equal to 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. In section 5 we discuss how the fixed cost parameter and survival probability can be determined from the data.

We assume that the aggregate state variables $[e, r_d, a]$ follow first-order Markov processes which we specify to match the macroeconomic environment during the crisis period. We also allow for interdependence between the domestic real interest rate and the exchange rate by assuming that

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

where ε' may be interpreted as a shock to the country-risk premium. We specify the form of $f()$ in section 5. Because the idiosyncratic shock to profitability, z, is an iid random variable, the information contained in the vector $[b, \eta, z]$ can be summarized in a single state variable, net worth, i.e.,

$$n \equiv \left(\sum_{j=d,f} \pi_j(a, e, z) k^{\alpha_j} - \sigma \right) + (1 - \delta) k - \Omega(e, e_{-1}; \omega) R_d(r_d, \eta) b \tag{14}$$

so that the value function may be defined as

$$\begin{aligned} v(k, n, \mathbf{z}) &= \max_{k', b', d} \{ (1 + \lambda) \, d + (\beta\mu) \, E[v(k', n', \mathbf{z}') | \mathbf{z}] \} \\ & \quad \text{s.t. } 14 \end{aligned}$$

where $\mathbf{z} = [a, z, e, e_{-1}, r_d]'$.

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

$$\frac{1}{\mu} = \beta E \left[\frac{1 + \lambda'}{1 + \lambda} \Omega(e', e; \omega) \left(1 + r'_d\right) \left(1 + \eta' + \frac{\partial \eta'}{\partial b'} b' \right) | \mathbf{z} \right] \tag{15}$$

Owing to the survival probability μ, the marginal benefit from issuing new debt is greater than one when evaluated at β, the market's discount rate.

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

$$\begin{aligned} 1 + \frac{\partial c}{\partial i} (i, k) &= (\beta\mu) \, E \left\{ \frac{1 + \lambda'}{1 + \lambda} \left[\frac{d}{dk'} d' \right. \right. \\ & \quad \left. \left. + (1 - \delta) \left(1 + \frac{\partial c}{\partial i'} (i', k') \right) \right] | \mathbf{z} \right\} \end{aligned} \tag{16}$$

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

$$\frac{d}{dk'}d' = \sum_{j=d,f} \alpha_j \pi_j(\mathbf{z}')k'^{\alpha_j - 1} - \frac{\partial c}{\partial k'}(i', k') - \Omega(e', e; \omega)(1 + r'_d)\frac{\partial \eta'}{\partial k'}b'. \tag{17}$$

The model cannot be solved analytically, we therefore use numerical methods to obtain an approximation to the solution. In particular, we adopt a version of Chebyshev projection methods (Judd(1992)) to approximate the solution of the model. Owing to the presence of occasionally binding constraints, we approximate the conditional expectations of the model first and then reconstruct 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).

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.

We now formally assess 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. We begin with a description of the empirical methodology, we then describe the data, after which we discuss the estimation results.

4.1 An Empirical Investment Equation

To measure fundamentals, we rely on 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 we show 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. In our regression analysis, we include 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.[11]

To construct our measure of Ω_{jt} we first measure the pre-crisis(1994∼1996) sample mean

[11]We use assets rather than capital in place since the former controls for cash on hand and inventory stocks, neither of which are formally controlled for in our model. An alternative would be to normalize using the capital stock in the denominator but subtract cash on hand and inventories from debt obligations in the numerator – this normalization provides very similar empirical results.

of each firm's foreign debt ratio, i.e.,

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

where $b_{j,t}^f$ 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.[12] Given $\hat{\omega}_j$, the effect of an exchange rate movement on the value of debt can be measured as

$$\hat{\Omega}_{jt} = 1 - \hat{\omega}_j + \hat{\omega}_j \left(e_t / e_{t-1} \right) \tag{18}$$

where e_t denotes the real exchange rate.[13] 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 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 $\hat{\Omega}_{jt} \left(b_{j,t}/a_{j,t} \right)$ are entirely attributable to movements in the exchange rate. Because the foreign-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 debt

[12] In our model firms are indifferent exante in terms of the currency composition of their debt. In reality, firms may wish to rebalance their foreign debt ratios. To avoid any endogeneity issues associated with such rebalancing we use the pre-crisis mean rather than the current period ratio or the full sample mean. Our empirical results are robust to either of these alternative formulations however.

[13] In our model formulation $\Omega\left(e, e_{-1}; \omega\right) \equiv \left[\omega \frac{e/e_{-1}}{E(e/e_{-1}|\mathbf{z}_{-1})} + (1-\omega) \right]$ and thus depends on the innovation in the exchange rate rather than the ratio e/e_{-1} as in equation 18. For our empirical regressions we use the latter formulation since it does not require us to take a stand on the dynamic process governing the exchange rate. In our structural estimation, we specifiy the model using $\Omega\left(e, e_{-1}; \omega\right)$ but estimate the simulated regression using the model's version of Ω_{jt} so that simulated and actual regressions are correctly specified to match each other. If the exchange rate is a random walk these formulations are identical.

sources.

In addition to our measures of the balance sheet and fundamentals, we control for firm and time fixed effects in our regression analysis. 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, we also allow for serial correlation in the investment process by including lagged investment on the right hand side of the regression.[14]

Our empirical investment equation is

$$(i/k)_{j,t} = c + c_j + \rho(i/k)_{j,t-1} + \boldsymbol{\alpha}'(\mathbf{s}/\mathbf{k})_{j,t} + \beta(\hat{\Omega}b/a)_{j,t} + \delta_t + \epsilon_{j,t} \tag{19}$$

where $(i/k)_{j,t}$ is investment normalized by the tangible capital stock, $(\mathbf{s}/\mathbf{k})_{j,t}$ is a vector of domestic and foreign sales normalized by the tangible capital stock, $[(s^d/k)_{jt}\ (s^f/k)_{jt}]$, $\boldsymbol{\alpha} = [\alpha^d\ \alpha^f]$ 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 where we consider separately the effects of the devaluation given the average foreign debt ratio and the overall beginning of period leverage ratio:

$$(i/k)_{j,t} = c + c_j + \rho(i/k)_{j,t-1} + \boldsymbol{\alpha}'(\mathbf{s}/\mathbf{k})_{j,t} + \beta\hat{\omega}_j(e_t/e_{t-1}) + \gamma(b/a)_{j,t} + \delta_t + \epsilon_{j,t} \tag{20}$$

[14] The lagged dependent variable can be justified if there is a distinction between measured and actual investment because of timing distinctions between reported and actual expenditures. Alternatively, serial correlation in unobservable investment cost shocks would also justify the use of a lagged dependent variable. In the empirical section, we consider regressions with and without the lagged dependent variable. Although it is significant, the regressions with and without the lagged dependent variable provide similar implications regarding the role of fundamentals and the balance sheet variable for investment.

In this regression, we isolate 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 $\mathbf{s}/\mathbf{k}_{jt}$ properly measures fundamentals. In general, fundamentals depend on the entire present discounted value of future profit rates. If $\mathbf{s}/\mathbf{k}_{jt}$ follows an AR1 process, then the present value $\mathbf{s}/\mathbf{k}_{jt}$ is proportional to the current value $\mathbf{s}/\mathbf{k}_{jt}$, we properly measure fundamentals. If $\mathbf{s}/\mathbf{k}_{jt}$ follows a richer stochastic process, then we have introduced 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 high foreign debt firms than low foreign-debt firms. Thus, if fundamentals are measured with error, our estimation procedure is biased against finding a negative effect of the balance sheet working through the exchange rate mechanism on investment.

4.2 Econometric Methodology

To estimate equations 19 and 20, we consider two estimators: an IV version of a fixed effect estimator and a panel-data GMM estimator. We use instrumental variables to control for the endogeneity that may exist between current sales and current investment.[15] The IV estimator is a standard 2SLS estimator that controls for fixed effects by removing group means. We adopt this estimator in part for its simplicity. It controls for firm-level heterogeneity and provides a reasonable summary of the data without applying complicated instruments sets or weighting matrices. This estimator thus has the virtue that it is easy to apply when estimating the structural model through indirect inference.

[15] If time to build for investment is less than one year, current sales and current investment suffer from simultaneity bias. By adopting an IV estimator with lagged values of sales as instruments we control for this possibility.

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 instrument choice in our panel data framework.

After taking first differences, equation 19 may be expressed as

$$\begin{aligned} \Delta(i/k)_{j,t} &= \rho\Delta(i/k)_{j,t-1} + \boldsymbol{\alpha}'\Delta(\mathbf{s}/\mathbf{k})_{j,t} + \beta\Delta(\hat{\Omega}b/a)_{j,t} + \delta_t + v_{j,t} \qquad (21) \\ v_{j,t} &= \epsilon_{j,t} - \epsilon_{j,t-1} \end{aligned}$$

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 $(\mathbf{s}/\mathbf{k})_{j,t-s}$ 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 $(\mathbf{s}/\mathbf{k})_{j,t-s}$ for $s \geq 1$ and $(\hat{\Omega}b/a)_{j,t-s}$ for $s \geq 1$ as instruments. When estimating equation 20 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. The data is

Table 1: Summary Statistics

	Full Sample Mean	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
$\left(s^f/s\right)_{j,t}$	0.284	0.279	0.251	0.307
$\left(b^f/b\right)_{j,t}$	0.140	0.189	0.140	0.140
$corr\left(s^e/s, b^e/b\right)$	0.1669		0.251	0.120

comparable to Compustat, the standard data set used for U.S. firm-level investment studies, in terms of the information provided. Unlike Compustat however, our data set covers both publicly traded and non-publicly traded firms. Unlike Compustat data, it also provides 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 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 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 our structural estimation described below. The median firm in our sample has export/sales ratio of 15 percent while

Table 2: Quantile 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)	

nearly 25% of the firms have almost no exports. Likewise, the median firm in our sample has a foreign-debt to total 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, we briefly consider the determinants of foreign-denominated debt. Specifically, we regress, 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 level 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 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

We now turn to our estimation results. Table 4 summarizes the main findings using both IV fixed effects and the first-differenced GMM specification.[16] The first column of estimates reported in Table 4 include 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-debt to total-debt ratio ($\omega_j = 0.14$), the estimated coefficient on $(\hat{\Omega}b/a)_{j,t}$ of -0.208 implies that the 70% devaluation would reduce the investment rate by five percentage points and can thus account for slightly less than half of the reduction in 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-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. Here we have included the lagged dependent variable for robustness. Again, we

[16] 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

find a statistically significant role for fundamentals as measured by the ratios of domestic sales and foreign sales to capital ratios. The coefficient estimates on the balance sheet variables are again negative, quantitatively large and statistically significant. When the balance sheet is broken out into its two components, beginning of period debt and the term $\omega_j \Delta e_t$ we again find an independent effect of the exchange rate interacted with the pre-sample foreign debt ratio. This coefficient is somewhat smaller in magnitude than the coefficient obtained with the IV estimator but is still larger than the coefficient on the debt-to-asset ratio. In all regressions, the coefficient on the lagged dependent variable is statistically significant though relatively small in magnitude.

In table five, we allow the devaluation to have non-linear effects which depend on the overall export and foreign debt position. To do so, we divide our sample between four sub-groups based on whether they are high vs low export firms and high vs low foreign-denominated debt firms. These classifications are again based on the median pre-crisis averages of export-sales and foreign-denominated debt ratios. For parsimony, we report only the GMM estimates.

According to table five, 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 -0.56 and highly significant. Firms who are least vulnerable – firms with high exports and low foreign debt actually exhibit a small positive response of investment to the balance sheet – the coefficient is 0.1. 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.

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
	(0.004)	(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)
$(\hat{\Omega}b/a)_{j,t}$	-0.401	-0.203	-0.197	-0.021
	(0.022)	(0.019)	(0.026)	(0.000)
$(i/k)_{j,t-1}$	0.145	0.148	0.130	0.209
	(0.004)	(0.005)	(0.011)	(0.000)
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

5 Structural Estimation

Structural estimation proceeds in two stages. In the first stage, we derive a parametric form of the profit function and apply conventional panel-data econometric techniques to identify relevant structural parameters. In the second stage, using the estimated parameters from the profit function, we solve the dynamic program numerically and simulate a complete set of panel data using our parametric policy functions. We calculate moments which summarize the actual panel data and the simulated panel data and use indirect inference to estimate the structural parameters of the model. We then use the estimated structural parameters to evaluate the role that foreign-denominated debt plays in propagating the financial crisis through investment spending.

When identifying the role of foreign-denominated debt on investment, we explicitly recognize 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, our structural estimation explicitly accounts for firm-level heterogeneity observed in the data. In particular, our estimation strategy conditions on the underlying distribution of export composition, foreign-denominated debt ratios and leverage.

We first consider the explicit functional form assumptions underlying the profit function and estimate relevant parameters. We then discuss our parameterization of the macroeconomic environment used in our structural estimation procedure, after which we discuss our estimation method based on indirect inference.

5.1 Production Technology, Market Structure and Profitability

To derive a closed-form profit function, we assume that firm j produces $y_t(j)$, a 2×1 vector composed of two differentiated goods, $y_{d,t}(j)$ and $y_{f,t}(j)$ with a constant-returns-to-scale Cobb-Douglas technology. Although the firm produces two differentiated goods, it employs only one type of capital, $k_t(j)$ and the production processes of both goods are subject to the same iid productivity shock, $a_t(j)$. The production technology also allows for both domestic variable inputs such as labor and foreign variable inputs such as imported materials:

$$y_t(j) = \begin{bmatrix} y_{d,t}(j) \\ y_{f,t}(j) \end{bmatrix} = \exp\left[a_t(j)\right] k_t(j)^{\alpha} \begin{bmatrix} (m_{d,t}(j)^{\sigma} n_{d,t}(j)^{1-\sigma})^{1-\alpha} \\ (m_{f,t}(j)^{\sigma} n_{f,t}(j)^{1-\sigma})^{1-\alpha} \end{bmatrix} \tag{22}$$

where $m_{d,t}(j)$, $n_{d,t}(j)$ are imported materials and labor inputs employed for the production of the domestic goods, and $m_{f,t}(j)$ and $n_{f,t}(j)$ are imported intermediate materials and labor inputs employed for the production of the foreign goods. Finally, α is the income share of the capital, $\sigma(1-\alpha)$ is the income share for the imported materials and $(1-\sigma)(1-\alpha)$ is the income share of labor.

In this framework, a firm with a given level of technology a and capital k must choose how to allocate variable inputs across the domestic and foreign markets to maximize profits. The firm faces monopolistic competition in both markets. We assume an iso-elastic demand curve and allow the elasticities to differ across the domestic and foreign markets, ε_i for $i = d, f$:

$$y_{i,t}(j) = \theta_i(j)\left[p_{i,t}(j)\right]^{-\varepsilon_i} Z_{i,t} \quad \text{for} \quad i = d, f \tag{23}$$

where $y_{i,t}(j)$ is the demand for the firm j's output in market i, $p_{i,t}(j)$ is the real price of the product in market i, $Z_{i,t}$ is an aggregate shock common to all firms in market i. The term $\theta_i(j)$ can be interpreted as a firm-specific constant term in the log-linear demand function of firm j in market i. In the long-run, firm size is determined by the firm-specific demand shifter $\theta_i(j)$. In the short-run, we allow for idiosyncratic shocks to production and aggregate shocks to market demands with the aggregate shocks being determined by a combination of exchange rate dynamics and demand shifters. The values of the aggregate shocks are normalized to one in the steady state.

The closed-form profit function of a firm can be written as a weighted average of sales in each market

$$\pi_t(j) = \sum_{i=d,f} \Gamma_i s_{i,t}(j) \tag{24}$$

where sales for market i satisfy

$$s_{i,t}(j) = \theta_i(j)^{\varsigma_i} \Xi_{i,t} \exp\left[a_t(j)\right]^{\vartheta_i} e_t^{\xi_i} k_t(j)^{\gamma_i} \quad \text{for } i = d, f \tag{25}$$

and the mark-up ratios in each market are given by

$$\Gamma_i = 1 - \chi_i \left(1 - \alpha\right). \tag{26}$$

Note that the mark-up ratios for both markets are constants determined by two parameters, the inverse of market power, χ_i and the production share of capital, α. $\Xi_{i,t}$ is a function of aggregate state variables, more specifically, a decreasing function of variable factor prices and an increasing function of demand shifter.

The elasticities in the sales equation are

$$\varsigma_i = \frac{1 - \chi_i}{1 - \chi_i(1 - \alpha)} \tag{27}$$

$$\vartheta_i = \frac{\chi_i}{1 - \chi_i(1 - \alpha)} \tag{28}$$

$$\gamma_i = \frac{\chi_i \alpha}{1 - \chi_i(1 - \alpha)} \tag{29}$$

$$\xi_i = \mathbf{1}(i = f) + \frac{\chi_i \left[\mathbf{1}(i = f) - \sigma\right]}{1 - \chi_i(1 - \alpha)} \tag{30}$$

for $i = d, f$. $\mathbf{1}(i = f)$ is an indicator function which takes unity when $i = f$ and zero otherwise. The elasticity with respect to capital is greater than zero but less than unity owing to market power($\chi_i < 1$). The elasticity with respect to the real exchange rate is negative for the domestic market owing to the dependence of production on imported materials. The elasticity for the foreign market is positive and bounded by $(1 + \chi_i \alpha) / [1 - \chi_i(1 - \alpha)]$, which is the case of an imported input ratio, $\sigma = 0$.

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 the sales function.[17] By taking logs of equation 25, we obtain the following fixed-effect regression specification with AR(1) error term, developed by Baltagi and Wu(1999) and Baltagi(2000),

$$\begin{aligned} \log s_{i,t}(j) &= \varsigma_i \log \theta_i(j) + \xi_i \log e_t + \gamma_i \log k_t(j) + \log \Xi_{i,t} + v_{i,t} \\ v_{i,t}(j) &= \rho_v v_{i,t}(j) + u_t(j), \quad u_t(j) \sim iidN(0, \sigma_u^2) \\ v_{i,t}(j) &\equiv \vartheta_i a_t(j) \end{aligned} \tag{31}$$

for $i = d, f$. All variables are real quantity values deflated by appropriate price indices. The regression includes real GDP indices to control for the influences of demand shifters. The domestic sales regression includes the log-differenced real GDP for Korea. The export sales regression includes the log-differenced index of world income obtained from the World Economic Outlook (WEO) data base obtained from the IMF.

Table 7 reports the results from estimating this equation. The estimated elasticity for foreign sales with respect to the real exchange rate is smaller than predicted by theory. Theoretically it must be greater than 1, as shown by the theoretical coefficient, $1 + \chi_i (1 - \sigma) / [1 - \chi_i(1 - \alpha)] > 1$. This might be the result of abstracting from pass-through

[17] Separate accounting data are available for domestic and foreign sales but not domestic and foreign 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

phenomena or pricing to market behavior in our theoretical model. The estimated coefficients for capital suggest a substantial degree of market power in both the domestic and foreign market. where the capital share in the production function, α is calibrated as 0.45 according to recent Bank of Korea(1995) estimates. The mark-up $\frac{1}{\hat{\chi}_d} = \frac{\alpha+\hat{\gamma}_d(1-\alpha)}{\hat{\gamma}_d}$ implied by the capital coefficients is stronger in the domestic market (1.642) than in the foreign market (1.376)

The estimated exchange rate coefficients imply a threshold value, 0.25, above which a firm's profit is increasing in the real exchange rate. In other words, if a firm's steady state export-sales ratio is greater than 0.25, then profits are increasing in the real exchange rate.[18] 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. This result implies that, on average, movements in the real exchange rate do not influence competitiveness in the Korean manufacturing

[18] If we approximate this arithmetic average form of the profit function using a geometric average, then the real exchange rate elasticity can be written as

$$\begin{aligned}\xi(j) &= \zeta_f(j)\xi_f + \zeta_d(j)\xi_d \\ &= \left(\xi_f - \xi_d\right)\zeta_f(j) + \xi_d\end{aligned}$$

where the last equality was from $\zeta_d(j) = 1 - \zeta_f(j)$. The firm specific weight, $\zeta_f(j)$ relfects the steady state export- sales ratio of firm j, so that firm j's profit is increasing in the real exchange rate only if the steady state export sales ratio is greater than the ratio

$$-\hat{\xi}_d / \left(\hat{\xi}_f - \hat{\xi}_d\right) = 0.25$$

sector.

5.2 Macroeconomic Shock Processes

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

$$\begin{aligned} 1+r_d &= (1+r_f)E\left(e/e_{-1}|\mathbf{z}_{-1}\right) \\ &= (1+\bar{r})\left(1+\xi\right)E\left(e/e_{-1}|\mathbf{z}_{-1}\right) \end{aligned} \tag{32}$$

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.[19] The exchange rate is assumed to follow an AR(1) process with persistence parameter ρ. We specify the data generating process for the country risk premium 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, UIP then implies:

$$r_d = \bar{r} + (\rho-1)\log e_{-1} + \xi. \tag{33}$$

Substituting the data generating process for the country risk premium:

$$r_d = \bar{r} + (\rho-1)\log e_{-1} + \bar{\xi}\left(1-\varphi\right) + \varphi\xi_{-1} + \varepsilon. \tag{34}$$

[19] One can think of $\bar{r}$ as the real US Treasury Bond rate and ξ as the spread on the emerging market government bond, for instance, the EMBI of J.P. Morgan.

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

$$r_d = (1-\varphi)\left(\bar{r}+\bar{\xi}\right) + (\rho-1)\left(\log e_{-1} - \varphi \log e_{-2}\right) + \varphi r_{d-1} + \varepsilon \tag{35}$$

Equation 35 implies the following time-series model for the real interest rate is

$$r_d = a_1 + a_2 \log e_{-1} + a_3 \log e_{-2} + a_4 r_{d-1} + \varepsilon \tag{36}$$

where $a_1 \equiv (1-\varphi)\,\bar{r}$, $a_2 \equiv (\rho-1)$, $a_3 \equiv -\varphi\,(\rho-1)$, and $a_4 = \varphi$. Notice that in the steady state, $\log e_{-1} = \log e_{-2} = 0$, and $r_d = r_{d-1} = \bar{r}+\bar{\xi}$.

We estimate two separate time series equations, a univariate AR(1) model of the real exchange rate, and the stochastic process for the domestic interest rate specified in equation 36. 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 $-\hat{a}_4\,(\hat{\rho}-1)$. Notice that 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. The persistence parameter for the real exchange rate, ρ, is estimated to be 0.80. This implies that $\hat{a}_2$ must be close to -0.2, which is close to the actual estimate.[20] The estimate, $\hat{a}_3$ is also close to $-\hat{a}_4\,(\hat{\rho}-1) = 0.108$. Finally, the estimation results imply a long-run real interest rate of 0.069.

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. We estimate

[20] Using pre-crisis data, the persistence parameter is estimated to be 0.801. If we include post-crisis data in our sample, the persistence parameter is estimated to be 0.596. This is primarily due to the fact that after the huge shock of devaluation in 1998, the real exchange rate appreciated substantially in the following year. If we use a dummy variable for 1998, then the parameter is estimated as 0.897, which is closer to a random walk. In addition to the persistence of shocks, we also condition on appropriate choices of the variance of shocks.

Table 7: MLE Estimates for r_d and Implied Structural Parameters

$\hat{a}_1$	$\hat{a}_2$	$\hat{a}_3$	$\hat{a}_4$	ρ	φ	$\bar{r}+\bar{\xi}$
0.032	-0.179	0.091	0.538	0.821	0.538	0.069
(0.094)	(0.003)	(0.142)	(0.003)			

an AR1 process for the log of domestic GDP over the sample period, 1990-2003, the implied persistence is $\rho_A = 0.3$. Because world output shows only small variation over this period, we fix it 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 agency cost, κ. The adjustment cost function is specified as

$$C(I,K) = \frac{\gamma}{2}\left(\frac{I}{K} - \delta\right)^2 K.$$

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:

$$x(\mathbf{s}_t;\mathbf{h}_j) \equiv \frac{\Omega_j(e_t, e_{t-1})b_t(j)}{\Pi_j(e_t, z_t(j))k_t(j)}$$

where $\Pi_j(e_t, z_t(j)) \equiv \sum_{i=d,f} \Gamma_i \theta_i(j)^{\varsigma_i} \Xi_{i,t} \exp\left[a_t(j)\right]^{\vartheta_i} e_t^{\xi_i}$ and measures the profitability of installed capital. 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 parameters are 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

$$\hat{\beta} = \arg\max_{\beta} Q_T(\mathbf{x}_T; \beta) \tag{37}$$

where $\mathbf{x}_T$ 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, $\beta = b(\theta)$ as a simulated counterpart of $\hat{\beta}$, i.e., a solution to $E_\theta \left[\partial Q(\mathbf{x}; b(\theta))/\partial b(\theta)\right] = 0$. In actual estimation, the binding function is replaced by its empirical counterpart,

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

where S is the number of simulations. The minimum distance estimator of the structural parameter vector, θ, is defined as

$$\hat{\theta}_{MD}^{S} = \arg\min \left[\hat{\beta} - \hat{b}_S(\theta)\right]' \Omega \left[\hat{\beta} - \hat{b}_S(\theta)\right] \tag{38}$$

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

$$\Omega_0 = A_0 B_0^{-1} A_0$$

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

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

where $avar(\hat{\theta}_{MD}^{S}) = (1 + 1/S)[\partial b(\theta_0)/\partial\theta\Omega_0\partial b(\theta_0)/\partial\theta']^{-1}$

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 3,

$$(i/k)_{j,t} = c_j + \beta^d (s^d/k)_{j,t} + \beta^e (s^e/k)_{j,t} + \beta^f (\hat{\Omega} b/x)_{j,t} + \delta_t + \varepsilon_{j,t}$$

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.

When generating the simulated data used to estimate the structural model, we also wish to control for firm-level heterogeneity.[21] To do so in a model consistent manner, we specify a firm-specific vector of individual characteristics, $\mathbf{h}_j$. The vector $\mathbf{h}_j$ 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. We estimate these firm-specific ratios using pre-crisis sample means. The dynamic programming problem of each individual in the simulation stage is a function of this individual characteristics vector, $\mathbf{h}_j$, hence the notation, $v(\mathbf{s}_j; \mathbf{h}_j)$ for the firm value. We also allow firms 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, $\mathbf{h}_j = [\omega_j, \theta_f/\theta_d, (b/k)_j]$ 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. For the real exchange rate, we use the actual realizations in the simulations. We do the same for the aggregate shock to profitability – it replicates the observed average drop in sales during the crisis period. Finally, 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, $\hat{b}_S(\theta)$ is an average of 100 IV Fixed Effect estimates.

Ideally, to completely control for firm heterogeneity, we would solve the value function and simulate the data for each firm in our sample. Because our data contain over 400 individual

[21] Recent researchers using indirect inference to estimate structural investment models with financial frictions have applied indirect inference to firm-level panel data., (Cooper and Ejarque(2004) and Whited and Hennessy(2004)). The models used in these studies do not allow for individual firm characteristics however.

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, we create a simulated panel with a smaller number of individuals, but which replicates the distributions of individual characteristics in the data. This is done in a following way: i) Estimate the empirical distribution functions for the three individual characteristics describe above. The quartiles of this distribution are reported in the table 2 . ii) Using this empirical distribution, calculate a joint distribution of the three individual characteristics. Since we rely on the quartile distribution, this procedure generates a panel with $4^3 = 64$ individuals. iii) Generate 64 time series for each simulation and 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. Effectively, 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.

Using this procedure, we estimate two structural parameters using three moments, namely, $[\hat{\beta}^d - \hat{b}_S^d(\theta), \hat{\beta}^e - \hat{b}_S^e(\theta), \hat{\beta}^f - \hat{b}_S^f(\theta)]$. 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. $\hat{\Omega} = [T\hat{V}(\hat{\beta})]^{-1}$. 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 a Sargan test statistic of overidentifying restrictions. For the Sargan statistics, we use the following statistics

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

5.4 Structural Estimation Results

We now report parameter estimates obtained from our indirect inference procedure. We consider a baseline case and two alternative estimates which vary the degree of persistence of the macroeconomic processes for aggregate output and the exchange rate. The alternative parameter estimates are considered for robustness. In particular, although we remove time effects from both the model and the data when matching moments, the nonlinearities inherent to the structural model may imply that our structural estimates are sensitive to the specification of the macroeconomic environment. In addition, when simulating the model, we wish to consider alternative but plausible assumptions regarding agents expectations of the macroeconomic processes.

In the baseline case, we set $\rho_A = 0.3$ and $\rho_e = 0.8$. These numbers are obtained from estimating the Korean output and exchange rate processes over the post-war period. The first alternative assumes a much higher persistence for output, 0.7 rather than 0.3. The second alternative assumes near random-walk for the exchange rate, i.e. a persistence of 0.98.

In all three cases, we use the actual realizations of the data to generate the realized values of output and exchange rates when computing estimates and model simulations. Varying the degree of persistence in the shock alters firms expectations regarding future outcomes however. In particular, with near random walk behavior of the exchange rate, firms no longer anticipate a sharp reduction in future interest rates following the devaluation.

Table 8 reports the auxiliary regression coefficients obtained from both the model and the data. Under our baseline estimation, the model successfully matches the auxiliary coefficients obtained from the IV fixed effect regression in the data. For the foreign and domestic sales-to-capital ratios, the coefficients obtained from the model are 0.0657 and 0.0443. 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 — (-0.2148) in the model versus (-02075) in the data. The auxiliary parameter estimates are not very sensitive to increasing

Table 8: Estimates of Auxilliary Parameters

		$(s^d/k)_{jt}$	$(s^e/k)_{jt}$	$(\hat{\Omega}b/a)_{jt}$
Data Moments		0.0692	0.0465	-0.2075
Simulated Moments				
Baseline	$(\rho_a = 0.30, \rho_e = 0.80)$	0.0657	0.0443	-0.2148
Alternative 1	$(\rho_a = 0.70, \rho_e = 0.80)$	0.0694	0.0475	-0.2244
Alternative 2	$(\rho_a = 0.30, \rho_e = 0.95)$	0.0763	0.0796	-0.2641

the degree of persistence in the GDP process. We see larger differences between the baseline case and the near random walk alternative for the exchange rate. In particular, both the sales and balance sheet coefficients increase somewhat relative to the data under the alternative scenario of near random walk behavior in the exchange rate.

Table 9 reports the structural parameters obtained from this estimation procedure, along with the test of over-identifying restrictions. The adjustment cost parameter is estimated to be 0.8968 with standard error of 0.0528 in the baseline case, which is similar to estimates reported by Cummins, Hasset and Hubbard (1994) and Gilchrist and Himmelberg (1998) in alternative contexts. This estimated adjustment cost is also much lower than what one would obtain using a Tobin's Q-style regression framework.

The structural coefficients imply an important role for financial market imperfections in the investment process. The coefficient measuring agency costs, κ, is estimated to be 0.085 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 a one percentage point rise in the premium on external funds. Thus, roughly speaking, if leverage doubles, the cost of external finance would rise by ten percentage points.

Table 9: Structural Parameter Estimates

	$\hat{\gamma}$	$\hat{\kappa}$	$\hat{J}$
Baseline ($\rho_a = 0.30, \rho_e = 0.80$)			
Estimates	**0.8968**	**0.0850**	**0.4071**
Standard Errors	(0.0528)	(0.0065)	(0.5234)
Alternative 1 ($\rho_a = 0.70, \rho_e = 0.80$)			
Estimates	**1.0920**	**0.0733**	**0.2474**
Standard Errors	(0.1042)	(0.0232)	(0.6189)
Alternative 2 ($\rho_a = 0.30, \rho_e = 0.95$)			
Estimates	**0.8906**	**0.0646**	**18.4534**
Standard Errors	(0.0046)	(0.0004)	(0.0000)

Finally, the baseline estimates also report the J-statistic for the over-identifying restriction. According to this J-statistic, one cannot reject the model's over-identifying restriction.

The alternative estimates imply similar results for the adjustment cost coefficient and the agency cost parameter. The over-identifying restrictions are not rejected when we allow for higher persistence to the GDP process. The over-identifying restriction is rejected for the near random walk exchange rate model however, suggesting some sensitivity of model moments to the specification of the aggregate processes despite the inclusion of time dummies in the estimation procedure.

The successful application of indirect inference relies on the model ability to provide meaningful variation in the auxiliary parameter estimates and therefore the loss function as we vary structural parameter values. To investigate the sensitivity of the auxiliary parameter estimates to structural parameters, we consider the effect of varying each structural parameter, holding the other parameter fixed at its estimated value.

These results are reported in table 10. We report both the implied J statistic as well as the auxiliary coefficients. Both the parameter estimates and the J statistic are highly

Table 10: Effects of Conditional Variations in the Structural Parameters

	$\hat{J}$	$(s^d/k)_{jt}$	$(s^e/k)_{jt}$	$(\hat{\Omega}b/a)_{jt}$
$\boldsymbol{\gamma = 0.8968}$				
$\kappa = 0.0200$	**154.71**	0.0042	0.0165	−0.1043
$\kappa = 0.0400$	**64.23**	0.0335	0.0120	−0.1562
$\kappa = 0.0800$	**7.08**	0.0550	0.0665	−0.2330
$\kappa = 0.0900$	**10.01**	0.0813	0.0512	−0.2798
$\kappa = 0.0950$	**22.18**	0.0622	0.0664	−0.3581
$\kappa = 0.0980$	**175.34**	0.0101	0.1529	−0.3953
$\kappa = 0.1000$	**389.45**	−0.0967	0.2321	−0.3345
$\boldsymbol{\kappa = 0.0850}$				
$\gamma = 0.0500$	**4816.81**	0.2905	0.5174	−0.6940
$\gamma = 0.2000$	**659.39**	0.2118	0.0968	−0.3403
$\gamma = 0.4000$	**223.97**	0.1413	0.0527	−0.4792
$\gamma = 0.6000$	**91.77**	0.1163	−0.0158	−0.3672
$\gamma = 0.8000$	**9.27**	0.0726	0.0535	−0.3053
$\gamma = 1.0000$	**128.62**	0.0873	−0.0740	−0.2630
$\gamma = 1.2000$	**155.20**	0.0922	−0.0815	−0.3719
$\boldsymbol{\kappa = 0.0000}$				
$\gamma = 0.0500$	**6077.51**	0.2904	0.4215	2.1576
$\gamma = 0.2000$	**3911.57**	0.2936	0.2370	1.7183
$\gamma = 0.4000$	**495.31**	0.1203	−0.1105	0.3348
$\gamma = 0.6000$	**229.28**	0.0568	0.0141	0.2928
$\gamma = 0.8000$	**225.11**	0.0490	0.0254	0.2858
$\gamma = 1.0000$	**202.69**	0.0296	0.0109	0.1670
$\gamma = 1.2000$	**205.11**	0.0370	0.0037	0.1863

sensitive to changes in the structural parameters. Holding the adjustment cost coefficient fixed at $\phi = 0.8968$ and varying the agency cost parameter κ over the range 0.02 to 0.1 produces clear variation in the J statistic. For values of κ outside the range 0.05 to 0.09 we see particularly large values for the loss function. Thus the agency cost parameter value is well identified by our estimation procedure. We also find that for nearly all of this parameter range, there is a monotonic increase in the absolute value of the regression coefficient for the balance sheet variable. Thus, as the size of the agency cost increases, the model implies a higher degree of sensitivity of investment to variation in the balance sheet in the reduced form regression.

Holding the agency cost parameter fixed and varying the adjustment cost also produces well-defined variation in the loss function. With low adjustment costs, investment is highly sensitive to both fundamentals and the balance sheet. Again, these estimation results imply that the adjustment cost coefficient is well identified by our estimation strategy.

Finally, we also consider the effect of assuming no financial frictions on the auxiliary parameter estimates. To do so, we fix $\kappa = 0$ and vary the adjustment cost coefficient. Strikingly, in the absence of adjustment costs, the model implies a positive relationship between investment and the balance sheet variable. This finding confirms our intuition that, in the absence of financial frictions, the balance sheet variable is positively correlated with investment fundamentals, and therefore, purely reduced form estimation procedures are likely to be biased against finding evidence that balance sheet variables influence investment behavior.

5.5 Model Simulations

We now consider the implications of our structural model for investment. We first consider the effect of the exchange rate devaluation on firm-level investment for firms with low versus high levels of foreign-denominated debt. We do this separately for firms with low exports and firms with high exports. These results are plotted in figures 3 and 4.

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.

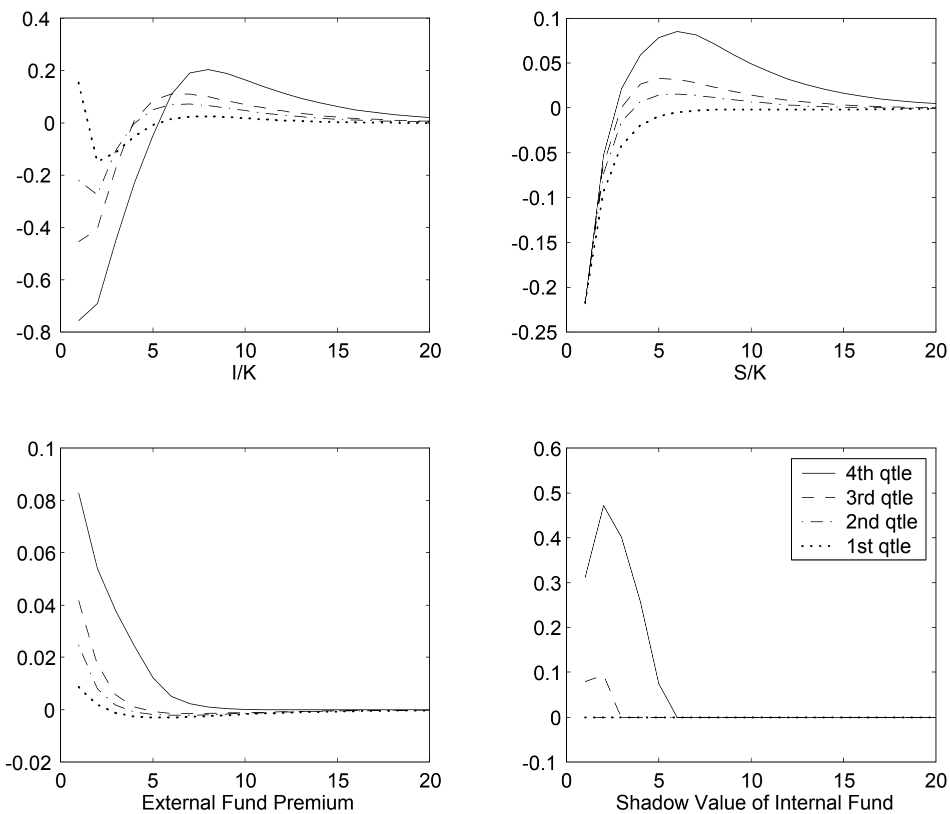

Figure 3: Impulse Responses of Low Export and High Foreign Debt Group

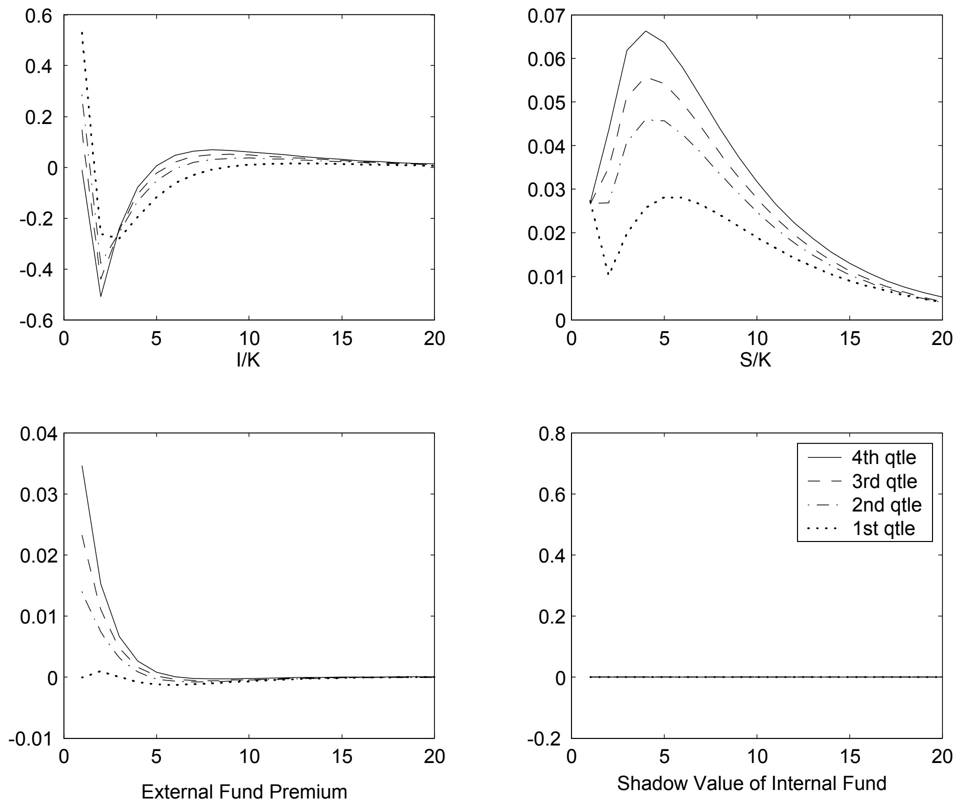

Figure 4: Impulse Responses of High Export and Low Foreign Debt Group

We plot these results for firms with foreign denominated debt ratios that vary from the highest quartile (solid line) to the lowest quartile (dotted line). Overall leverage is set at the median value. For firms with low exports, 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 eight percentage points. The overall corporate bond spread rose by 9% during this period. Thus, our model provides relatively conservative movements in the premium on external finance. The rise in the cost of external funds is larger for the firm with high foreign-denominated debt, as a result, the investment rate is substantially lower for such firms.

Figure 4 plots the same experiment for firms with a high export share (the third quartile).

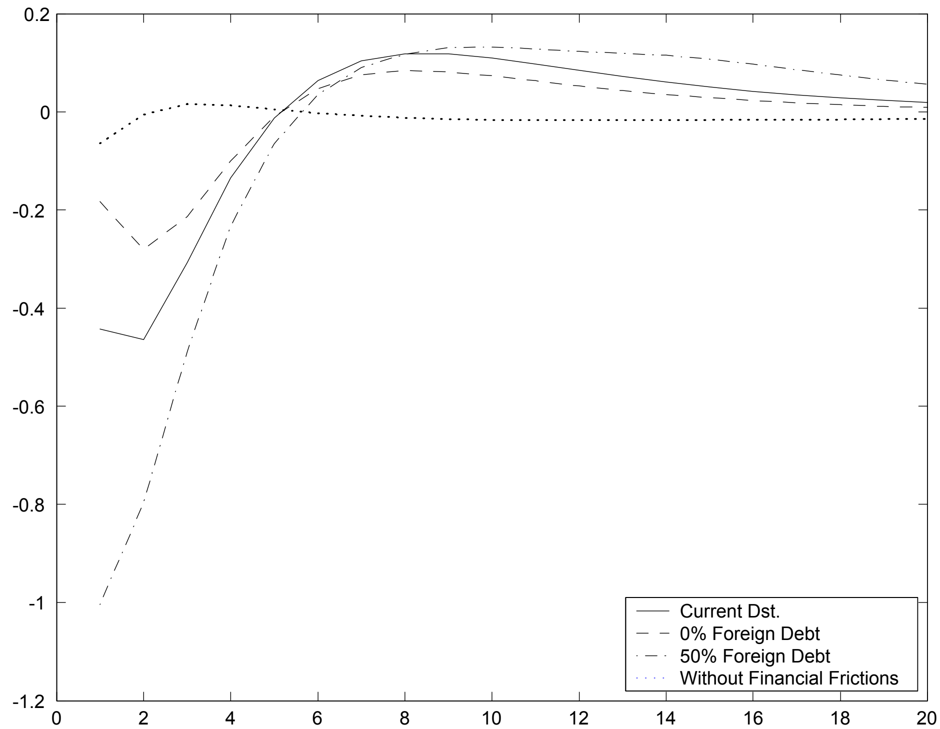

Figure 5: Aggregate Investment Under UIP Condition

Again we consider firms low versus high foreign denominated debt. For firms who export, the devaluation combined with macroeconomic shocks implies an increase in sales and investment. The external finance premium rises by 0 to 3.5 percentage points depending on the degree of foreign denominated debt exposure. Again, the increase in the premium is larger for firms with high foreign-denominated debt, which leads to a somewhat lower investment rate in the initial period.

We now turn to the aggregate implications of our structural estimates. Figure 5 considers 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.[22] The resulting path for investment is plotted in the solid line in figure 5. We then 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 debt ratio of fifty percent (dotted line). The foreign-denominated debt levels 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 a 45% reduction in investment. This is lower than the observed 100% reduction in investment during the crisis. Thus our macroeconomic simulation understates the overall contraction in investment spending.[23] According to our simulations, foreign-denominated debt explains one half of the model's aggregate investment dynamics – in the absence of foreign-denominated debt, investment contracts by 20% rather than 40%. Even in the absence of foreign-denominated debt, financial frictions are 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 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. As a result, with financial frictions, the model predicts a 40% drop in investment, whereas, absent financial frictions, the model predicts an investment response which is close to zero.

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

[22] We have also computed a value weighted response in a similar manner. These estimates imply similar conclusions regarding the role of foreign denominated debt and the role of the balance sheet operating through the exchange rate.

[23] We are currently investigating alternative macroeconomic assumptions and their role in generating a sizeable contraction in investment. These include relaxing UIP and increasing the expected persistence to the aggregate demand shock (which currently is set at 0.3 on an annual basis). Relaxing the UIP condition generates an investment contraction on the order of 70%. Our conclusions regarding the importance of foreign-denominated debt remain unchanged however.

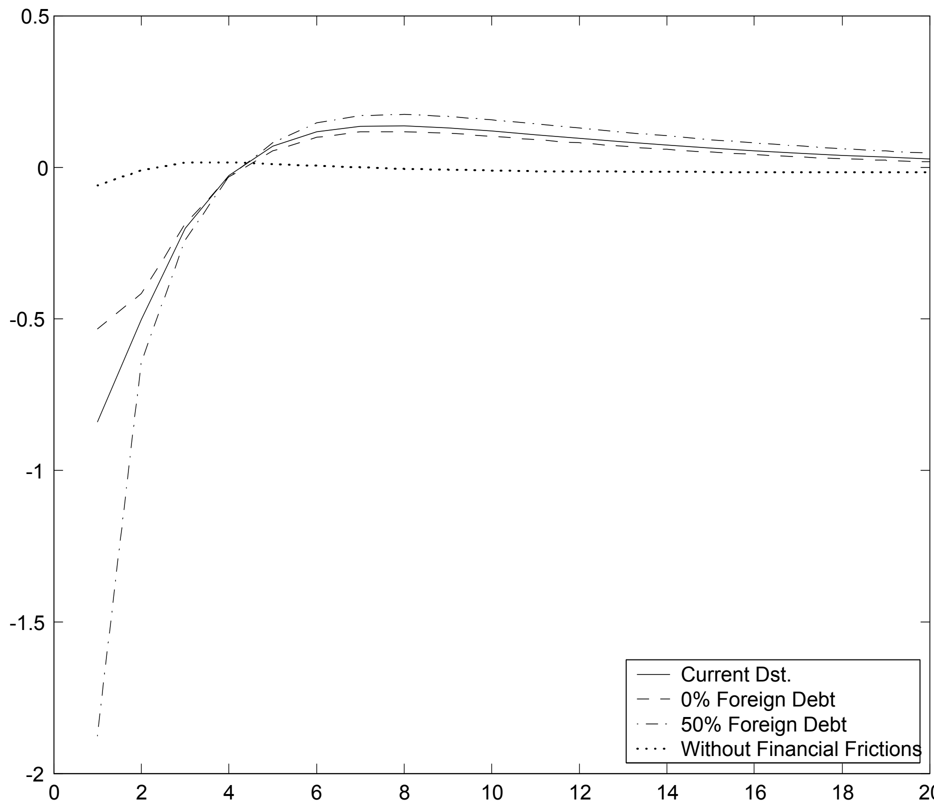

Figure 6: Aggregate Investment Under the Near Random Walk Hypothesis

a 50% foreign-denominated debt ratio. Here, we find a sizeable effect of foreign-denominated debt on investment spending. The simulated contraction in investment is now on the order of 100%. 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.

Figure 6 considers the same experiment but now allows the for near random walk behavior in the exchange rate. By increasing the persistence in the exchange rate process, firms no longer expect a sharp drop in interest rates following the devaluation. Thus investment contracts by much more in this scenario. Using the current distribution of foreign-denominated debt, our model predicts a 90% drop in investment, which is line with the actual experience.

Increasing the foreign-denominated debt ratios to 50% implies a doubling of the investment response relative to the baseline case.

6 Conclusion:

This paper studies the effect of foreign-denominated debt on investment spending during the Korean financial crisis. The presence of foreign-denominated debt exerted a strong influence on investment at the micro-level. This is found to be true in both reduced-form regressions and structural parameter estimates obtained from a model of firm-level investment that allows for both real and financial frictions. Our structural parameter estimates allow us to conduct counterfactual exercises. These exercises imply that the presence of foreign-denominated debt likely reduced aggregate investment by 25% to 50% during crisis. This finding holds true despite the fact that foreign-denominated debt levels are relatively low for Korean manufacturing firms. Our counterfactual simulations suggest that foreign-debt ratios on the order of 50% would have doubled the observed contraction in investment following the devaluation. Foreign-denominated debt levels on the order of 50% are high relative to Asian economies but in line with observed foreign-denominated debt levels for Latin American economies during the 1980's and 1990's. These estimates suggest that policy makers should take into account the effect of a devaluation on the debt burden when assessing the potential gains to various policy responses during exchange rate crises.

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Appendix A: Derivation of the Profit Function.

The profit function is defined as

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

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

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

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

$$
\begin{aligned}
e_t w_{m,t} m_{i,t}(j) &= (1-\alpha)\sigma\chi_i \left[\theta_i\left(j\right) Z_{i,t}\right]^{1-\chi_i} e_t^{\mathbf{1}(i=f)} y_{i,t}(j)^{\chi_i} \\
w_{n,t} n_{i,t}(j) &= (1-\alpha)(1-\sigma)\chi_i \left[\theta_i\left(j\right) Z_{i,t}\right]^{1-\chi_i} e_t^{\mathbf{1}(i=f)} y_{i,t}(j)^{\chi_i}
\end{aligned}
$$

where $\mathbf{1}(i = f)$ is an indicator function which takes one if $i = f$, and zero otherwise. Subsituting these conditional demand functions in the profit results in the following profit function

$$
\pi_t(j) = \sum_{i=d,f} \Gamma_i s_{i,t}(j)
$$

where the mark-up ratios and the sales for each market are given by

$$\begin{aligned}
\Gamma_{i,t} &= 1 - \chi_i (1 - \alpha) \\
s_{i,t}(j) &= e_t^{\mathbf{1}(i=f)} y_{i,t}(j)^{\chi_i} \left[\theta_i(j) Z_{i,t}\right]^{1-\chi_i}
\end{aligned}$$

Note that the mark-up ratios are constants for both markets. If the firm has the same market power in both markets, the mark-up ratios in both markets are equalized in the steady state since $e_{ss} = 1$. To get the closed form of profit function, we substitute the conditional demand functions in the sales functions to get

$$s_{i,t}(j) = \theta_i(j)^{\varsigma_i} \Xi_{i,t} \exp\left[a_t(j)\right]^{\vartheta_i} e_t^{\xi_i} k_t(j)^{\gamma_i} \quad \text{for } i = d, f$$

where the elasticities of sales functions with respect to state variables are the same as described in the text. $\Xi_{i,t}$ is a complicated function of exogenous aggregate variables. It is a decreasing function of variable factor prices and a increasing function of aggregate income shocks.

Appendix B: Data Construction.

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

$$k_{j,t+1} = (1-\delta)k_{j,t} + \frac{I_{j,t}}{P_{k,t}} \tag{39}$$

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.