Internet Appendix: Do declines in bank health affect borrowers’ voluntary disclosures? Evidence from international propagation of banking shocks

This document provides the details of the robustness checks mentioned in Section 4.3 of the paper, including tests that (i) use alternative regression specifications, (ii) study borrowers matched on firm size, (iii) minimize borrowers’ exposures to foreign markets, (iv) address concerns due to incomplete coverage in the CIG database, and (v) study the non-crisis periods to confirm that the results are limited to periods of declining bank health. The details follow.

i) Alternative specifications: The results are qualitatively similar if I change the regression to (1) a logit model without firm-fixed effects, or (2) a Poisson specification and using forecast frequency during the quarter as the dependent variable. Also, the results persist if I include additional control variables (e.g., borrowers’ credit rating or Bharath et al. (2008)’s default score used in Section 4.4) to control for changes in borrowers’ credit risk. In addition, controlling for changes in the status of the equity market (as proxied by the number of SEOs in a quarter) provide similar results.

ii) Studying borrowers matched on firm size: The treatment group of borrowers includes firms larger than those in the control group. The difference in firm size could affect the results if larger firms tend to change their disclosures in the crisis period for (unknown) reasons other than banking shocks. To address this issue, I study a restricted sample of borrowers matched on firm size. For each treatment borrower, I first find all control borrowers whose market value in 1997:Q2 is within 5 percent of the treatment borrower’s market value. Then I pick the closest control borrower in terms of ROA as the matching firm. The matching is performed without replacement, and I am able to find 384 matched pairs using this procedure. As expected, firm size is similar between the treatment and the matching borrowers (p-value = 0.904). Despite the smaller sample, the results remain qualitatively similar (coefficient on Crisis × Exposed = 0.306, z = 3.074), suggesting differences in firm size between the two borrower groups do not drive the results.

iii) Additional restrictions on firms’ foreign operations: As discussed in Section 3.1, my sample had little direct exposure to the crisis areas, so it is unlikely that different borrower exposure to the crises would confound my results. Nonetheless, I repeat the test using a restricted sample of 865 firms with negligible exposures to the foreign markets, i.e., firms whose aggregate sales across Asia, Europe, Pacific Basin, and South America were lower than 5 percent of the firm’s total sales in the fiscal year prior to the crises. Despite the smaller sample, the results remain similar (coefficient on Crisis × ExpomBK = 0.300, z = 3.127).

iv) Addressing concerns about incomplete forecast coverage in the CIG database: Chuk et al. (2009) suggest that forecast coverage in the CIG database is less complete before 1997. These results imply that the sample of forecasts used in my tests may be incomplete. If for unknown reasons forecast coverage increases more for borrowers of exposed banks in 1997 than for other borrowers, then this will create the appearance of the reported differential forecast changes in the

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1 Please note that the main test already has included a number of time-varying variables to control for changes in borrowers’ credit quality, including the firm’s market size, past stock returns, return volatility, and ROA. Prior studies find that these factors are the most important determinants of a firm’s default risk (e.g., Shumway 2001).
crisis period (regardless of banking shocks). This alternate explanation, however, does not explain why I observe a gradually weakening forecast tendency for borrowers of exposed banks when the crises subsided, nor does it readily explain the results from the cross-sectional tests of differences in treatment effects (Section 4.4). Importantly, Chuk et al find that forecasts not tracked by the CIG database are more likely to have low precision. To the extent that the potential greater increase in forecast coverage for borrowers of exposed banks captures these previously missing forecasts, it should drive down the average forecast precision and forecast informativeness for these borrowers more than for other borrowers, which is opposite to what is reported in Section 4.5. Nonetheless, to minimize the concern due to incomplete forecast coverage, I follow Chuk et al’s suggestion and repeat the test using a restricted sample of borrowers followed by at least three analysts. Despite the smaller sample, the results remain similar (coefficient on Crisis × ExpoMBK = 0.379, z = 4.087).

ev) Null results from examining non-crisis periods: In Table 3, I use borrowers’ forecasts in the post-crisis period to reduce concerns that the results for the crisis period are spuriously driven by factors that cause a continuing forecast increase over time for borrowers of exposed banks (e.g., structural trends that mainly affect these borrowers). Alternatively, I study the non-crisis periods to rule out such concerns. If the concerns were to hold, I expect to find similar differential forecast increases over time even in periods with no banking shocks. The test examines three different two-year windows (around June 30 of 1996, 2003, and 2004) during which no banking shocks are expected. See the following figure for the timeline.

The regression specification is shown below:

\[ Pr(MF\_DUM_{it} = 1) = \alpha_i + \beta \ Pseudo\_Crisis_t + \gamma \ Pseudo\_Crisis_t \times ExpoMBK_i + \delta \ Controls_{it-1} + \epsilon_{it} \]

Essentially, the test examines whether borrowers of exposed banks exhibit differential increases in forecast tendency in the second half of the non-crisis period (denoted by Pseudo\_Crisis) relative to the first half. I conduct the test separately for each of the three alternate periods. In line with the lack of treatment effect in these periods, results show that in none of the three cases is the coefficient on Pseudo\_Crisis × ExpoMBK significant ((a) Pseudo\_Crisis\_1996 × ExpoMBK = -0.070, z = -0.586; (b) Pseudo\_Crisis\_2003 × ExpoMBK = -0.282, z = -1.385; and (c) Pseudo\_Crisis\_2004 × ExpoMBK = -0.017, z = -0.068).
REFERENCES

