Evaluating Bank Recapitalization Programs in Japan: How Did Public Capital Injections Work?

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Abstract. Public capital injection into the banking system is a comprehensive policy program aimed at reducing the financial risks faced by capital-injected banks and thereby stimulating their lending and profitability. Using a difference-in-difference estimator in a fixed effect panel model, this paper evaluates empirically two large-scale capital injections in Japan; the first was the public capital injection into 21 Japanese banks in March 1998 based on the Financial Function Stabilization Act, and the second was into 15 Japanese banks in March 1999 based on the Prompt Recapitalization Act. We find that the two public injections significantly reduced the financial risks faced by the capital-injected banks, but did not stimulate their lending and profitability. Next, we investigate what factors impeded bank lending after the two public injections. We provide evidence that the deterioration of borrowers' creditworthiness because of Japan's severe recession and banks' accompanying increased perception of the riskiness of lending inhibited not only the injected banks but also the noninjected banks from lending more.

JEL classification: G01, G21, G28.

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1. Introduction Public capital injection into the banking system is a comprehensive policy program aimed at reducing the financial risks of capital-injected banks and thereby stimulating their lending and profitability. The financial crisis after the Lehman shock in 2008 and the global recession that followed forced industrialized countries, including England, France, Germany, Ireland, the US and Switzerland, to implement such bank recapitalization programs. Accordingly, a macroeconomic framework to conceptualize theoretically how the policy program works has been developing (see, e.g., Gertler and Kiyotaki (2010)), but no empirical consensus exists on whether it has produced the desired results. This paper utilizes Japan's two large-scale capital injections in 1998 and 1999, which are regarded as precedents for the European and US public capital injections, as a natural experiment in bank recapitalization policy, and attempts to offer new insights into the actual implementation of public capital injection into the banking system.

Theoretically, when asymmetric information exists, an increase in a bank's financial risk profile can cause its lending behavior to deteriorate. The phenomenon where a bank restrains its lending because of an increase in its financial risk is called a "capital crunch". Indeed, several papers found evidence supporting the existence of capital crunches both in the US and in Japan in the 1990s (see, e.g., Bernanke and Lown (1991) and Peek and Rosengren (1995) for studies of capital crunches in the US, and Woo (2003) and Watanabe (2007) for studies of Japan' s experience). Previous studies of Japanese bank recapitalization programs in 1998 and 1999 mainly focused on whether the two programs resolved the capital crunch of banks needing a capital injection.

The favorable view of the effect of Japan's public capital injections suggests that they reduced the default risk of the capital-injected banks, thereby improving their lending (see Allen et al. (2009) and Giannetti and Simonov (2009)). Figures 1 and 2 show the historical paths of the probability of default and bank loans to domestic enterprises of Japanese banks, divided into two groups: the treated group that has entered into the bank recapitalization programs, and the control group that has not. [1] Figure 1 shows that the probability of default of the treated group decreased drastically after the two public capital

[1] See Section 2 for the method of calculating the probability of default, and for the definition of bank loans.

injections in 1998 and 1999, while that of the control group rarely changed before and after the capital injections. On the other hand, Figure 2 demonstrates that the bank loans not only of the treated group but also of the control group decreased continuously after the capital injections. Casual observation reveals that the favorable view cannot successfully explain why the lending by the injected banks did not improve, even though their financial conditions improved substantially.

One promising explanation is that the policy framework of the two Japanese capital injections that forces each capital-injected bank to maintain and raise its capital ratio ends up impeding its lending, as pointed out by Osada (2010). The unfavorable view, however, ignores and fails to explain why the relatively stable financial condition of noninjected banks and their decreasing loans to domestic enterprises coexisted.

Irrespective of these opposing views about the effect of Japan's public capital injections, they share a common assumption, that the lending of Japanese banks after the capital injections was substantially determined by lender-side factors such as the banks' financial condition and profitability. However, once we note that many borrowers' creditworthiness deteriorated during the severe recession after Japan's two large-scale capital injections, we cannot simply ascribe the stagnant bank lending after the capital injections to lender-side factors. In other words, the increased default risk and the decreased profitability of borrowing firms shown in Figure 3 appear to be dominant factors causing stagnant bank lending after the public capital injections. [2]

Given that public capital injection is a comprehensive policy program designed to stabilize the banking system first and then to stimulate bank lending and profitability, this paper evaluates Japan's public capital injections in 1998 and 1999 in terms of the following three issues:

1. To what extent did the public capital injections in 1998 and 1999 contribute to reducing capital-injected banks' financial risks such as default risk and nonperforming loans?

[2] The empirical study of US capital injections by Berrospide and Edge (2010) demonstrated that the US slowdown in loan growth after the capital injections can be attributed more to the prevailing recession in the US and banks' increased perception of the riskiness of lending than to their capital position.

2. If the public capital injections contributed to a decrease in the financial risks of the injected banks, did they also increase their lending to domestic enterprises and profitability?

3. Was there room to improve bank lending to domestic enterprises using the two capital injections in the first place? If not, how can we explain the sluggish bank lending after the public capital injections shown in Figure 2?

This paper addresses the above issues empirically by estimating the treatment effects of the public capital injections with a difference-in-difference estimation method in a fixed effect panel model. Our approach to evaluating Japan's public capital injections is different from that of previous studies as follows.

First, previous studies (Allen et al. (2009) and Giannetti and Simonov (2009)) measured the responses of bank lending at the time of public capital injection. Taking into consideration the characteristics of public capital injection, it is more important to select outcome variables linked to their policy objectives and then to measure their change over time. We attempt to capture such a duration effect for Japan's public capital injections in terms of causal inference. [3]

Second, as pointed out by Conley and Taber (2011), when the number of members belonging to the treated group is much smaller compared with that belonging to the control group, the standard large-sample approximations are not appropriate for conducting statistical inference of a treatment-effect estimate obtained using a fixed effect panel model. Following the method of Conley and Taber (2011), we conduct rigorous statistical inference of the treatment-effect estimate based on the empirical distribution.

Third, using information on borrowing enterprises listed on the Tokyo Stock Exchange, we control for borrower-side factors including borrowers' default risks and profitability to investigate in detail the lending of Japanese banks after the public capital injections. Such borrower-side factors have not been considered explicitly by previous studies. We examine

[3] Spiegel and Yamori (2003) and Yamori and Kobayashi (2007) employed an event study approach to evaluate Japan's capital injections. Unlike our study, these attempted to analyze the very short-term effects of Japan's capital injections by estimating stock market responses to the announcement of the public capital injections into the banking system.

whether bank lending after the public capital injections was more determined by lender-side or borrower-side factors respectively.

Our paper is organized as follows. Section 2 discusses a method for the estimation of the treatment effect and our data set. Section 3 reports the estimation results. Section 4 analyzes bank lending after the capital injections by controlling for borrower-side factors. Section 5 provides a conclusion. Appendix I illustrates the treatment effect. Appendix II discusses the method of statistical inference developed by Conley and Taber (2011). Appendix III discusses how the amount of capital injection into each bank influenced its default risk by introducing heterogeneity in the treatment effect.

2. Estimation Method of the Treatment Effect of Public Capital Injection In November 1997, four financial institutions (Sanyo Securities, Hokkaido Takushoku Bank, Yamaichi Securities and Tokuyo City Bank) failed, and Japan experienced its greatest financial crisis in the postwar period. Since then, the Japanese government has decided to use public funds in order to deal with the financial crisis, although until then the Japanese government had feared public opinion against the use of public funds (see Nakaso (2001) and Hoshi and Kashyap (2010)).

For the actual implementation of public capital injection, the Financial Function Stabilization Act (hereafter FFSA) came into effect in February 1998. The first capital injection based on the FFSA was approved in March 1998 for 21 banks, and a total of 1,815.6 billion yen (1,080 billion yen for subordinated debt, 414.6 billion yen for subordinated loans and 321 billion yen for preferred stock) was paid on March 30 (see Table 1).

Six months later, in October 1998, the FFSA was abolished, and instead the Prompt Recapitalization Act (hereafter PRA) came into effect. Accordingly, the limit of the amount of public capital allowed to be injected into banks was increased from 13,000 billion to 25,000 billion yen. The second capital injection based on the PRA was approved for 15 banks in March 1999, and a total of 7,459.25 billion yen (1,300 billion yen for subordinated debt and loans and 6,159.3 billion yen for preferred stock) was paid on March 30.

The laws of the FFSA and PRA stipulate the policy objectives of public capital injection, and hence the banking supervisory agency supervises a capital-injected bank to ensure that its actions are consistent with the policy objectives. In this section, we first outline an econometric method for estimating the treatment effects of the two public capital injections in 1998 and 1999. We then define financial variables corresponding to the policy objectives stipulated by the FFSA and the PRA, thereby specifying our econometric models more precisely.

2.1. **Econometric Method** Let t^* denote the time at which public capital is injected into problematic banks. Then, we denote $D_{it} = 1$ if bank i belongs to the treated group at time $t = t^* + k$ $(k \geq 0)$ in which banks have entered into a recapitalization program at time t^*, and $D_{it} = 0$ if bank i belongs to the control group at time t in which banks have not entered into the program at time t^*. Let us suppose that this indicator variable takes the value $D_{it^*-1} = 0$ for all banks i at time $t^* - 1$.

Given the treatment indicator D_{it}, we introduce the following two models to estimate the treatment effect on the capital-injected banks:

$$\begin{aligned} \textbf{Model I}: \quad y_{it} &= \alpha y_{it-1} + \beta' \mathbf{X_{it}} + \gamma_t t + \delta D_{it} + v_i + \varepsilon_{it}, \\ \textbf{Model II}: \quad y_{it} &= \alpha y_{it-1} + \beta' \mathbf{X_{it}} + \gamma_t t + \delta_t (t \cdot D_{it}) + v_i + \varepsilon_{it}, \end{aligned}$$

where y_{it} is an outcome variable for bank i, and $\mathbf{X_{it}}$ are time-varying covariates other than y_{it-1}. t is a time dummy variable that regards time $t^* - 1$ as the reference point of time, and its coefficient parameter γ_t captures the time effect that is common to all banks but varies across time. v_i is the fixed effect term for bank i, and ε_{it} is the stochastic error term. [4]

The fixed effect term v_i plays a critical role as an unobservable confounding factor that determines managerial decisions of bank i, including the decision about whether the bank

[4] For studies on labor economics and health economics, it is more common to use the method based on the estimated propensity score to estimate the treatment effect (for example, in the context of labor economics, see Heckman, Ichimura and Todd (1997)). The method based on the propensity score requires a cross-sectional data set in which the cross-section dimension is very large and the time dimension is one. On the other hand, for our study on Japanese banks, we have a panel data set in which the time dimension is two or more and the cross-section dimension is not very large because the number of domestic banks is not large. Taking into account the structure of our data set, we do not use the propensity score method, but use the method based on individual fixed effects to estimate the treatment effects of Japan's public capital injections. For studies on labor economics that measure the treatment effect using panel data as in this paper, see Ashenfelter and Card (1985) and Jacobson, Lalonde and Sullivan (1993).

enters into the recapitalization program. In the above models, we assume that the fixed effect term v_i captures banks' time-invariant managerial ability. As in the conventional fixed effect models, v_i can be correlated not only with the treatment indicator D_{it}, but also with the covariates y_{it-1} and $\mathbf{X_{it}}$ and each other. [5]

Now, let us suppose that the outcome variable of bank i takes a value of y_{1it} at time $t = t^* + k$ $(k \geq 0)$ if it has received the capital injection at time t^* $(D_{it} = 1)$ and y_{1it} at time t if it has not $(D_{it} = 0)$. Then, we can define the treatment effect on the treated group, denoted by TE, as follows:

$$\begin{aligned} TE &= E(y_{1it} - y_{0it} | D_{it} = 1) \\ &= E(y_{1it} | D_{it} = 1) - E(y_{0it} | D_{it} = 1). \end{aligned}$$

To measure TE, we have to estimate $E(y_{0it}|D_{it} = 1)$: the expected value of the counterfactual outcome that would be realized if a capital-injected bank has not been recapitalized. We cannot, however, estimate the expected value directly from the observational data because the counterfactual outcome is not observable. [6] Then, we introduce the following assumption into Models I and II:

$$E(y_{0it}|D_{it},\ y_{it-1},\ \mathbf{X_{it}},\ t,\ v_i) = E(y_{0it}|y_{it-1},\ \mathbf{X_{it}},\ t,\ v_i). \tag{1}$$

Equation (1) implies that the recapitalization program is randomly assigned across banks at time $t = t^* + k$ $(k \geq 0)$ as long as y_{it-1}, X_{it}, t and v_i are conditional. By employing this assumption, the treatment effect at time t can be expressed as an estimate of the parameter

[5] The main issue in identifying the treatment effect involves how to deal with the unobservable confounding factor. In the framework of panel data analysis, one promising way is to introduce an individual fixed effect term as the unobservable confounding factor. For details, see Angrist and Pischke (2009, Chapter 5) and Wooldridge (2010, Chapter 10). Wooldridge (2009) demonstrated that for estimation of the treatment effect, the fixed effect model can be applied even to an unbalanced panel data set where sample selection can be correlated with an individual fixed effect term and possibly also with treatment indicators and time-varying covariates.

[6] If the public recapitalization program is randomly assigned across all banks, $E(y_{0it}|D_{it} = 1) = E(y_{0it}|D_{it} = 0)$ holds for time $t = t^* + k$ $(k \geq 0)$. However, that assumption is not appropriate because the program is not randomly assigned.

coefficient δ_t in Model II as follows:

$$\begin{aligned}\delta_t &= E(y_{1it} - y_{0it}|\ y_{it-1},\ \mathbf{X_{it}},\ t,\ v_i) \\ &= E(y_{1it} - y_{0it}|\ D_{it} = 1,\ y_{it-1},\ \mathbf{X_{it}},\ t,\ v_i) \\ &= \{E(y_{1it}|\ D_{it} = 1,\ y_{it-1},\ \mathbf{X_{it}},\ t,\ v_i) - E(y_{1it^*-1}|\ D_{it^*-1} = 0,\ y_{it^*-2},\ \mathbf{X_{it-1}},\ t^* - 1,\ v_i)\} \\ &\quad - \{E(y_{0it}|\ D_{it} = 1,\ y_{it-1},\ \mathbf{X_{it}},\ t,\ v_i) - E(y_{1it^*-1}|\ D_{it^*-1} = 0,\ y_{it^*-2},\ \mathbf{X_{it-1}},\ t^* - 1,\ v_i)\}\,,\end{aligned} \quad (2)$$

where the second equality follows from equation (1). From the third equality in equation (2), we can interpret the estimate of δ_t as a difference-in-difference estimate in which time $t^* - 1$ is the reference point of time. More precisely, the duration effect of the public capital injection, or δ_t, can be defined as the difference between the actual variation of the outcome variable (the first brace term) and the counterfactual variation of it (the second brace term). The difference between the actual and counterfactual variations measures the treatment effect of the capital injection on the outcome variable in terms of causal inference. Appendix I illustrates the difference-in-difference estimate, and discusses it in more detail.

The treatment effect in Model I can be expressed as an estimate of the parameter coefficient on D_{it} as follows:

$$\delta = E(\delta_t) = E(y_{1it} - y_{0it}|D_{it} = 1, y_{it-1},\ \mathbf{X_{it}},\ t,\ v_i).$$

In the following, we measure the treatment effect on the capital-injected banks by estimating the parameter coefficients δ and δ_t in Models I and II. For estimation of the parameter coefficients, we conduct a two-stage within-group estimation using y_{it-2} as an instrumental variable for y_{it-1} because Models I and II have the dynamic structure with a lagged dependent variable y_{it-1}. [7]

For consistency of an estimator of a coefficient parameter in a fixed effect panel model,

[7] In a dynamic panel model with a lagged dependent variable y_{it-1}, either y_{it-2} or $y_{it-2} - y_{it-3}$ is promising as an instrumental variable (for a more detailed discussion, see Arellano and Bond (1991) and Blundell and Bond (1998)). Our results do not qualitatively change when we use $y_{it-2} - y_{it-3}$ as an instrumental variable.

the strict exogeneity condition, which requires that the stochastic error term should be uncorrelated with covariates over time, is necessary. Given the structure of our dynamic panel models, however, the strict exogeneity condition could not be met. Accordingly, we also estimate Models I and II that do not include the lagged dependent variable y_{it-1} as an explanatory variable using the conventional within-group estimation method, thereby checking the robustness of our estimation results. [8]

As stated above, we estimate the causal effect of Japan's public capital injection using a fixed effect panel model. Conley and Taber (2011), however, pointed out that the standard large-sample approximations are not appropriate for conducting statistical inference of a treatment effect estimate obtained using a fixed effect panel model when the number of members of the treated group is much smaller than that of the control group. Thus, following Conley and Taber (2011), we conduct statistical inference for estimates of the treatment effect parameters δ and δ_t. The method of statistical inference developed by Conley and Taber (2011) is based on the empirical distribution calculated using residuals ε_{jt}, generated from the control group equation of a noninjected bank j. Appendix II discusses the procedure to calculate the empirical distribution in more detail.

Our sample period for estimating the treatment effect of the first recapitalization program ranges from September 1997 to September 1998 because $t^* =$ March 1998, while that for estimating the treatment effect of the second recapitalization program ranges from September 1998 to March 2002 because $t^* =$ March 1999. The reason why the sample period for analyzing the second recapitalization program is up to 2002 is that the third recapitalization program based on the Deposit Insurance Law was implemented for Resona Bank in 2003. To extract the pure effect of the second recapitalization program before the third program, we set the end of the second subsample period at 2002.

2.2. **Data Set and Benchmark Model** In this subsection, we define the outcome variable y_{it} and covariates $\mathbf{X_{it}}$ corresponding to the policy objectives of Japan's public capital injections in 1998 and 1999, thereby resulting in an improved specification of Models

[8] For estimating the treatment effect with a fixed effect model, Angrist and Pischke (2009, Chapter 5) proposed using two types of fixed effect models with and without a lagged dependent variable and then checking the robustness of treatment effect estimates.

I and II. As stated above, the public capital injections in 1998 and 1999 were based on the two laws of the FFSA and the PRA, respectively. To discipline capital-injected banks, these laws stipulate the following policy objectives: 1) reduction of the default risks of capital-injected banks; 2) write-offs of nonperforming loans; 3) improvements in profitability; 4) improvements in bank lending to domestic enterprises, including small and medium-sized ones; and 5) expenditure cuts through adjustment of employment costs, the number of board members and the number of branch offices.

The FFSA and the PRA discipline capital-injected banks in line with these policy objectives, but the ultimate purpose of policy objective 5 is to improve the financial condition of capital-injected banks and revitalize their profitability. Accordingly, this paper focuses on policy objectives 1 to 4, and consequently offers the following seven financial variables as the outcome variable y_{it}:

Related to policy objective 1): a variable measuring the default risk of bank i

1. Probability of default (PD_{it}),
2. Tier I ratio (TIER_{it}),
3. Capital ratio (RATIO_{it}),

Related to policy objective 2): a variable measuring the nonperforming loans of bank i

4. Nonperforming loan (NPL_{it}),

Related to policy objective 3): a variable measuring the profitability of bank i

5. Return on assets (ROA_{it}),

Related to policy objective 4): a variable measuring loans to enterprises offered by bank i

6. Loans for domestic enterprises (LOAN_{it}),
7. Loans for small and medium-sized enterprises (SMELOAN_{it}).

The three variables related to a bank's default risk, the probability of default (PD_{it}), the Tier I ratio (TIER_{it}) and the capital ratio (RATIO_{it}), are alternatives to each other.

We thus include one of the three variables in Models I and II prepared from four financial variables: NPL_{it}, ROA_{it}, LOAN_{it} and SMELOAN_{it}. In addition, we include either the Tier I ratio or the capital ratio in place of the probability of default in Models I and II, and thereby check the robustness of the estimation results.

The probability of default (PD_{it}) is theoretically based on Merton's (1974) structural option-pricing model. In what follows, let V_A represent the banks' asset value (market value), σ_A the asset volatility and r the risk-free rate. Furthermore, we denote by D the book value of the debt that has maturity equal to T. According to Merton (1974), the market value of equity V_E can be thought of as a call option on V_A with time to expiration of T, and hence plays the role of the strike price of the call option. The market value of equity V_E and the volatility of equity valuation σ_E are then given by the Black and Scholes (1973) formula for call options:

$$V_E = V_A N(d_1) - De^{-rT} N(d_2), \tag{3}$$

$$\sigma_E = \left(\frac{V_A}{V_E}\right) N(d_1)\sigma_A, \tag{4}$$

where

$$d_1 = \frac{\ln\left(V_A/D\right) + \left(r + \frac{1}{2}\sigma_A^2\right)T}{\sigma_A\sqrt{T}}, \qquad d_2 = d_1 - \sigma_A\sqrt{T},$$

and N denotes the cumulative density function of the standard normal distribution.

In the framework of Merton's (1974) structural model, once the numerical value of d_2 is obtained, the risk-neutral probability of bank default is calculated as $N(d_2)$. We use the risk-neutral probability as a measure of banks' credit risk, converting it to percentage terms.

To compute the risk-neutral probability of default, it is necessary to estimate two unknowns, the banks' asset value V_A and asset volatility σ_A, using data for each time period of the five observables, the market value of equity V_E, the volatility of equity valuation σ_E, the book value of debt liabilities D, the time to maturity of the debt T and the risk-free rate r, from the two nonlinear simultaneous equations (3) and (4). To solve this system, we employ the reduced gradient method, and use the market value of equity V_E calculated from both the banks' daily stock-price data and the number of shares outstanding provided by Nomura Research Institute (hereafter NRI). [9] To estimate the volatility of equity valuation σ_E, we calculate the standard deviation of the market value of equity V_E for the past 20 business days of each trading day. In addition, we express the estimated volatility of the equity valuation at annual rates as in the following equation:

$$\sigma_{Et} = \sqrt{\frac{1}{20-1} \times \sum_{i=t}^{t-19} (ret_i - \overline{ret_t})^2} \times \sqrt{240},$$

where $ret_t = \ln(V_{Et}) - \ln(V_{Et-1})$ denotes the daily rate of change in equity valuation, and $\overline{ret_t}$ is the average rate of change in equity valuation of the previous 20 days.

The book value of debt liabilities D is obtained from semiannual published accounts (nonconsolidated base) compiled by NRI, and is linearly interpolated to yield daily observations. The time to maturity of the debt T is set at one year, which is the conventional assumption to construct a measure of default risk theoretically based on Merton's (1974) model including the distance to default marketed by the Moody/KMV Corporation (see Crosbie and Bohn (2003) for more details). [10] For the risk-free rate r, the one-year swap rate observed for each trading day is used. More precisely, we construct the swap rates based on the average rate of offers and bids quoted by Yagi Euro, one of the major dealers in the interest-rate swap market in Japan. Finally, we compute the monthly average of the probability of default to ensure consistency with our semiannual data set. Nonperforming loans (NPL_{it}) are defined as the ratio of the reported amount of nonperforming loans to total loans. We use the book values of the nonperforming loans and total loans, each of which is from semiannual published accounts compiled by NRI and includes the logarithm value of the nonperforming loan ratio in Models I and II. [11]

[9] The number of shares outstanding used for our empirical analysis is adjusted according to a TOPIX-type computation from the secondary capital transfer data.

[10] Vassalou and Xing (2004), Gropp et al. (2006), Gilchrist et al. (2009) and Harada and Ito (2011) set the time to maturity of debt liabilities to one year for calculation of their indicators of default risk based theoretically on Merton (1974).

[11] The book value of nonperforming loans is defined as the sum of loans to borrowers in legal bankruptcy and past due loans for which there had been no payments of interest or principal for six months or more. We use both bank and trust accounts to calculate the nonperforming loans. We also define the nonperforming loan ratio as the ratio of such loans to the book value of total assets. The difference in our definition of

Return on assets (ROA_{it}) is constructed by dividing banks' net profits by the book value of their total assets, and is expressed in percentage terms. The banks' net profits are preliminarily adjusted on a monthly basis by dividing them by the number of accounting periods. Net profits and total assets are financial statements compiled by NRI.

Loans for domestic enterprises (LOAN_{it}) and loans for small and medium-sized enterprises (SMELOAN_{it}) are defined as the ratios of loans for domestic enterprises and loans for small and medium-sized enterprises (hereafter SMEs) to the book value of total assets, respectively. [12] To construct the two loan ratios, we use the data set compiled by NRI for loans to domestic enterprises and SME loans. We then include the logarithm values of the ratios in Models I and II. Some previous studies that investigate bank-lending behavior in Japan used the growth rates of bank loans (see, e.g., Montgomery (2005), Giannetti and Simonov (2009) and Osada (2010)). We also use the period-by-period growth rates of bank loans, but estimation results obtained using the growth rates do not qualitatively differ from those using the loan ratios. We report only the estimation results obtained using the loan ratios in Section 3.

We rotate the use of the above five financial variables (PD_{it}, NPL_{it}, ROA_{it}, LOAN_{it} and SMELOAN_{it}) as the output variable and the time-varying covariates. More precisely, when we cast one of the five variables as the outcome variable y_{it} in Models I and II, we include one-period lags of the remaining four variables and the one-period lag of the relative size (SIZE_{it}) of bank i as the covariates $\mathbf{X_{it}}$. [13] We define the relative size (SIZE_{it}) of bank i at time t as follows:

$$SIZE_i = \ln\left(\frac{V_{Ai}}{\sum_{j=1}^{n} V_{Aj}}\right),$$

where V_A is the banks' asset value defined in the above construction of the probability of default, and n is the number of banks listed on the Tokyo Stock Exchange at time t.

As discussed above, we include the probability of default (PD_{it}) as a measure of banks'

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the nonperforming loan ratio does not make any qualitative difference to our estimation results.

[12] The book values of loans for domestic enterprises and loans for SMEs are defined by including loans of trust accounts as well as of bank accounts.

[13] For example, when casting the probability of default as the output variable y_{it}, we include the one-period lags of NPL_{it}, ROA_{it}, Loan_{it}, SMELOAN_{it} and SIZE_{it} as the time-varying covariates $\mathbf{X_{it}}$. SIZE_{it} serves as a covariate in Models I and II, but not as the output variable.

default risk. We also include the Tier I ratio (TIER_{it}) and the capital ratio (RATIO_{it}) as alternative measures of default risk instead of the probability of default to check the robustness of our estimation results. The Tier I ratio is defined as the percent ratio of Tier I capital to banks' risk assets. For the capital ratio, we use the book value of banks' self-capital ratios. The book values of Tier I, the risk assets and the self-capital ratios are from semiannual published accounts compiled by NRI.

This paper evaluates the treatment effects on 21 banks that received the first capital injection in March 1998, and 15 banks that received the second capital injection in March 1999 (see Table 1). The sample size for the first capital injection is 303 with 103 Japanese banks (21 banks in the treated group and 82 banks in the control group) listed on the Tokyo Stock Exchange for the sample period from September 1997 to September 1998, while that for examining the second capital injection is 751 for 99 Japanese banks (15 banks in the treated group and 84 banks in the control group) for the sample period from September 1998 to March 2002. During the second subsample period after the second capital injection in March 1999, 17 regional banks other than the 15 injected banks sporadically received public capital injections under the PRA. To extract the pure effects of the first and the second public injections, we initially exclude data for the 17 regional banks. [14]

In the construction of our data set, it is worth noting that by March 2002, after the second capital injection in March 1999, four mergers took place among injected banks in the treated group of the second subsample. [15] The four mergers partially changed the composition of the treated group. To control for such a merger effect on the composition of the treated group, we use the four continuing banks formed by the four mergers before

[14] Capital injections based on the PRA were implemented intermittently until March 2002. The capital injection in March 1999 was implemented for major banks, while each capital injection after April 1999 was implemented for 17 regional banks based on the subprogram "Basic vision for strengthening the capital bases of regional banks" announced by the Japanese government in June 1999. We exclude the data of the 17 regional banks to extract the pure effects of the capital injection in March 1999, even though the capital injection in March 1999 and each capital injection after April 1999 were based on the PRA. Furthermore, in June 2003, a capital injection was implemented for Resona Bank based on the Deposit Insurance Law. For our data set, we do not define Resona Bank as a bank that received the first and second capital injections in 1998 and 1999.

[15] Four mergers took place. Chuo Trust Bank and Mitsui Trust Bank in 2000; Daiichi Kangyo Bank, Fuji Bank, Industrial Bank of Japan and Yasuda Trust Bank in 2000; Sakura Bank and Sumitomo Bank in 2001; Sanwa Bank, Tokai Bank and Toyo Trust Bank in 2001.

March 2002 as survival banks of the premerged banks in the treated group. [16]

Figures 1, 2 and 4 show the historical path of each financial variable from 1997 to 2002. The solid line indicates the path of the treated group that received the two capital injections, and the dashed line indicates the path of the control group that did not receive them. We point out the following observations about the figures. First, Figure 1 indicates that the default risk of the treated group became much higher than that of the control group just before the first injection in 1998, but decreased drastically just after this injection. Second, as shown by Figure 2, the bank loans of both the treated and control groups decreased consistently irrespective of the public capital injections in 1998 and 1999. Third, Figure 4 observes that SME loans were always higher in the control group. Nonperforming loans were larger in the treated group before the second injection in 1999, whereas they were smaller after the second injection. The path of the return on assets indicates that bank profitability fell sharply in 1999. The historical path of the relative size shows that the firm size of the treated group was always considerably larger than that of the control group.

In the following section, we report the estimation results of Models I and II obtained using our data set.

3. Estimation Results The estimation results of Model I appear in Tables 2 to 5. In particular, Tables 2 and 3 show the parameter estimates obtained by including the probability of default (PD_{it}) as a measure of bank default risk, while Tables 4 and 5 include the Tier I ratio (TIER_{it}) and the capital ratio (RATIO_{it}), respectively.

All estimates of the treatment effect δ are initially converted to percentages. Furthermore, for each of the estimates, Tables 2 to 5 report its level of significance with asterisks and its 95% confidence interval in parentheses, each based on the empirical distribution constructed following Conley and Taber's (2011) method. For estimates of covariates, on the other hand, Tables 2 to 5 report their 95% confidence intervals in parentheses, each based on the large-sample approximations. Appendix II discusses Conley and Taber's

[16] Regarding the mergers among Japanese banks that took place in the late 1990s and 2000s, Harada and Ito (2011) empirically demonstrated that merged banks inherit the financial conditions of the premerged banks. Similar to our study, they used an indicator of bank fragility theoretically based on Merton's (1974) model. According to their findings, our approach to dealing with the four mergers after the second capital injection would not be substantial for our estimation results reported in the next section.

(2011) method for constructing the empirical distribution.

As shown in Tables 2 and 3, our estimation results of Model I are qualitatively unaffected by whether the lagged dependent variable y_{it-1} is included or not.

In the estimation results of Model I in which a measure of bank default risk is used as the output variable, estimates of the treatment effect δ on the probability of default (PD_{it}), the Tier I ratio (TIER_{it}) and the capital ratio (RATIO_{it}), each reported in Tables 2 to 5, indicate that the first and second capital injections reduced the default risks of the capital-injected banks significantly. [17] Parameter estimates of the covariates of NPL_{it} and ROA_{it} imply that larger nonperforming loans and lower profitability were significantly responsible for the higher default risks of Japanese banks. Parameter estimates of LOAN_{it} and $\mathrm{SMELOAN}_{it}$, even though some are not significant, indicate that an increase in the amount of credit granted to domestic enterprises is likely to cause an increase in the default risks of Japanese banks. The estimated coefficients for the relative size (SIZE_{it}) imply that the default risks of larger banks decreased more than those of smaller banks.

For Model I in which nonperforming loans are determined, the treatment-effect estimates of NPL_{it} indicate that the first and second capital injections both reduced the amount of nonperforming loans held by the capital-injected banks, while the second did so more significantly than the first. The parameter estimates of the covariates of the default risk indicators (PD_{it}, TIER_{it}, RATIO_{it}) and ROA_{it} imply that Japanese banks with higher default risks and lower profitability had a larger number of nonperforming loans. The parameter estimates of LOAN_{it} and $\mathrm{SMELOAN}_{it}$ indicate that the number of nonperforming loans was positively associated with that of bank loans to domestic enterprises. [18]

[17] There are two issues concerning the estimates of the treatment effect for the first capital injection, which is considered to have been implemented based on an ineffective policy scheme in which injected banks' capital requirements were not fully tested (see, e.g., Allen et al. (2009)). First, our estimation results for the first capital injection are quite consistent with the movement of the Japan premium demonstrated by Hoshi and Kashyap (2010). Indeed, the Japan premium reached a peak of almost 110 basis points in December 1997. However, it started to fall in January 1998, when the government outlined a policy scheme for injecting public funds into problematic banks, and ended up falling below 20 basis points in March 1998. In this way, the Japan premium dropped to a much lower value in March 1998. Second, some studies for US public capital injections based on the Troubled Assets Relief Program (TARP) (see, e.g., Greenspan (2010) and Veronesi and Zingales (2010)) provided evidence that it significantly reduced bank default risk even though it was implemented without a bank stress test to determine the injected banks' capital requirements.

[18] Hoshi (2001) analyzed the determinants of the nonperforming loans of Japanese banks in the 1980s

When the measure of bank profitability is the output variable in Model I, estimates of the treatment effect on ROA_{it} imply that the first and second capital injections did not improve the profitability of the capital-injected banks. Parameter estimates of the covariate of NPL_{it} imply that the profitability of Japanese banks was negatively associated with their number of nonperforming loans. In addition, the parameter estimates of $\mathrm{SMELOAN}_{it}$ suggest that an increase (decrease) in SME loans impaired (improved) the profitability of Japanese banks.

Regarding the estimation results of the loan supply functions of Japanese banks, there are two issues in relation to previous studies of the two public capital injections.

First, our treatment-effect estimates of LOAN_{it} and $\mathrm{SMELOAN}_{it}$ indicate that the first and second capital injections did not have substantial effects on the lending behavior of the capital-injected banks. These estimation results do not support those of Allen et al. (2009) and Giannetti and Simonov (2009), which argued that the second capital injection in 1999 improved the lending behavior of the capital-injected banks, but supported those of Osada (2010), which argued that the first and second capital injections in 1998 and 1999 did not improve it. [19]

Second, our parameter estimates of the covariates in Model I indicate that the four indicators of bank fragility (the probability of default, the Tier I ratio, the capital ratio and nonperforming loans) did not significantly determine bank lending (LOAN_{it} and $\mathrm{SMELOAN}_{it}$) for our sample period from 1998 to 2002. On the other hand, Osada (2010) found that his bank fragility indicators (the Tier I ratio and the capital ratio) significantly determined bank lending for his sample period from 1993 to 2006. [20] As discussed later

and Ogawa (2003, Chapter 2) in the 1990s. They found that increases in the number of SME loans, and loans to real estate businesses, the construction industry and the finance and insurance industry, were responsible for increases in the number of nonperforming loans.

[19] Unlike this paper, Allen et al. (2009) and Giannetti and Simonov (2009) used a discrete variable at each time of the two capital injections in one equation; the former defined the discrete variable as the amount of the capital injection, and the latter simply as a dummy variable for indicating that a bank did or did not receive the capital injection. Such a specification for identifying policy effects does not involve a particular period as the reference period. Hence, it would lead to a wrong judgment on the policy effects, because the value of an outcome variable after the preceding policies serves as the reference value to evaluate subsequent policies.

[20] Ito and Sasaki (2002) estimated a loan supply function in Japan from 1990 to 1993, Woo (2003) from 1989 to 1997 and Ogawa (2003, Chapter 2) from 1992 to 1999. They all found that an increase in nonperforming loans would cause a decrease in bank loans. Hosono (2006) observed that a decrease in

in Section 4, such a difference in parameter estimates of the financial fragility indicators carries different implications regarding the reason why the lending of the capital-injected banks did not improve.

We now report the estimation results of Model II. Figures 5 and 6 plot estimates of the time-varying treatment effect δ_t obtained using Model II. The respective paths of the treatment effect for the first and second capital injections intersect at March 1999, with September 1997 serving as the reference period. Figure 5 shows the treatment-effect estimates of the probability of default, the Tier I ratio and the capital ratio, each of which is obtained by including the respective indicators in Model II using five variables: NPL_{it}, ROA_{it}, LOAN_{it}, SMELOAN_{it} and SIZE_{it}. Figure 6 shows treatment-effect estimates of NPL_{it}, ROA_{it}, LOAN_{it} and SMELOAN_{it} obtained using Model II in which the probability of default is included as a measure of bank default risk. The treatment-effect estimates of the four variables are expressed in percentages. In addition, for all the estimates of the treatment effect δ_t, Figures 5 and 6 show their 90% confidence intervals in dashed lines as well as their point estimates in solid lines. The confidence intervals are constructed following the method of Conley and Taber (2011), which is discussed in more detail in Appendix II.

Overall, the estimates of the time-varying treatment effect δ_t obtained with Model II are consistent with those of the treatment effect δ obtained with Model I. More precisely, Figure 5 indicates that the two public capital injections reduced the default risks of the capital-injected banks. Figure 6 also shows that the second capital injection in 1999 worked particularly well in reducing the number of nonperforming loans of the capital-injected banks. On the other hand, as with the estimates of the treatment effect for ROA_{it}, LOAN_{it} and SMELOAN_{it} obtained with Model I, Figure 6 also provides unfavorable evidence about the effect of Japan's public capital injections; the first and second capital injections in 1998 and 1999 had no significant effect on capital-injected banks' profitability and loans to domestic enterprises, including SME loans.

the self-capital ratio caused a decrease in bank loans in the 1990s. The difference between our estimation results and theirs is the sample periods used for estimation; the previous studies, including Osada (2010), used a data set covering much of the 1990s, during which the capital crunch resulting from the deterioration of banks' assets was severe.

In sum, our estimation results indicate that the first and second capital injections in 1998 and 1999 were quite effective in reducing the financial risks of the capital-injected banks through recapitalization and write-off of nonperforming loans, while they did not play a critical role in improving the profitability and the lending behavior of the capital-injected banks. In particular, our findings for the loan supply functions of Japanese banks partially support those of the previous study by Osada (2010) of the two public capital injections in terms of the fact that the loan supply of the capital-injected banks was sluggish, but partially do not in terms of the fact that the lending of Japanese banks after the capital injections was not substantially determined by lender-side factors.

In the next section, we examine in depth the determinants of bank lending after the capital injections by controlling for borrower's default risk and profitability.

4. Bank Lending after the Public Capital Injection In the previous section, we observed that the first and second capital injections in 1998 and 1999 probably reduced the financial risks of the capital-injected banks through recapitalization and write-offs of nonperforming loans. Furthermore, as with Osada (2010), we also observed that the two capital injections did not significantly improve the lending behavior of the capital-injected banks.

Osada (2010) observed that financial risk factors of Japanese banks such as the Tier I ratio and the capital ratio had significant effects on bank lending. From this observation, he inferred that the policy framework of public capital injection that prompts a capital-injected bank to maintain and raise its capital ratio ended up impeding its lending, and he then concluded that capital adequacy requirements were fundamentally responsible for its sluggish loan supply to domestic enterprises. [21]

Did capital adequacy requirements really impede the lending of the capital-injected banks? As described above, we observed that financial risk factors such as the probability of default and the number of nonperforming loans do not explain the lending of Japanese banks after the public capital injections. In this section, we examine in depth the determinants

[21] Montgomery (2005) and Osada (2010) demonstrated theoretically that an increase in a bank's financial risk would reduce its lending when the capital ratio is small or the supervision on the capital ratio is strict, while the bank's lending does not depend on its financial risk when the bank is not strictly subject to capital ratio regulation and supervision of the capital ratio.

of the banks' lending after the public capital injections.

More specifically, we investigate whether bank lending after the public capital injections was more determined by loanable funds supply or demand factors using the following loan supply function:

$$\textbf{Model III}: \quad LOAN^{j}_{it} = \alpha LOAN^{j}_{it-1} + \beta'_i \mathbf{X_{it}} + \beta'_j \mathbf{X^{j}_{it}} + \gamma_t t + v_i + \varepsilon_{it},$$

where LOAN^{j}_{it} indicates the total amount of loans outstanding that domestic listed company j borrows from bank i at time t. We define the total amount of loans outstanding by adding short-term debt with a maturity of one year or less to long-term debt with a maturity of more than one year, and then taking the logarithm of the total amount of loans outstanding. We obtain the annual amount of total loans outstanding for each listed company j from the Corporate Borrowings from Financial Institutions Database compiled by Nikkei Digital Media Inc.

$\mathbf{X_{it}}$ and $\mathbf{X^{j}_{it}}$ are time-varying covariates to capture the financial risks and the profitability of bank i and listed enterprise j that borrows from bank i, respectively. t is a time dummy variable to control for the common factors for the Japanese economy at time t, and v_i is a bank dummy variable to capture a time-invariant bank-specific factor.

For the lender-side covariates $\mathbf{X_{it}}$, we use the one-period lags of PD_{it} and NPL_{it}, defined in Section 2, as a proxy for bank i's financial risks. As an alternative to PD_{it}, we also use the one-period lag of the "adjusted capital ratio" ($\mathrm{ADRATIO}_{it}$), defined by subtracting the target capital ratio (8% for international banks and 4% for domestic banks) from a reported capital ratio to control for the effect of capital adequacy requirements. In addition, as a proxy for the profitability of bank i, we include the one-period lag of ROA_{it} in Model III.

For the borrower-side covariates $\mathbf{X^{j}_{it}}$, we use the one-period lag of the probability of default of borrower j (PD^{j}_{it}), which embodies the default risk of domestic listed company j that borrows from bank i. Additionally, we use the one-period lag of the return on assets (ROA^{j}_{it}) to examine whether the profitability of borrower j determines the lending behavior of bank i. The procedure for constructing the two covariates of borrower j is the same as the procedure for constructing those of lender i, which is discussed in Section 2. For construction of the covariates, we use annual published accounts data compiled by

NRI.

The sample period used for estimation is from March 1998 to March 2002, which is in accordance with that used for estimating the treatment effect of the capital injections in Section 3. Our data set covers 103 banks and 1,026 borrowing firms listed on the Tokyo Stock Exchange. Thus, our data set does not fully cover SMEs. The number of samples is 35,749 (out of which 11,226 are involved with the capital-injected banks and 24,523 are involved with the noncapital-injected banks). Figure 3 shows the probability of default and the return on assets on the borrower side. The figure indicates that the default risk of borrowing firms increased drastically after 1998, while their profitability decreased.

Table 6 reports the estimation results of Model III. For the estimation of Model III, we employ the ordinary least squares estimation method. The upper panel reports estimation results obtained using the adjusted capital ratio as a financial risk factor of bank i, while the lower panel uses the probability of default. We make the following remarks concerning the estimation results of Model III.

First, bank loans to domestic listed companies are not determined by lender-side financial risk (ADRATIO_{it}, PD_{it} and NPL_{it}) and profitability (ROA_{it}), but by borrower-side financial risk (PD^{j}_{it}) and profitability (ROA^{j}_{it}).

Second, the estimation results for the injected and noninjected banks are not qualitatively different from each other. We can therefore infer that at least for domestic listed companies, their deterioration in creditworthiness would cause the sluggish loan supply of Japanese banks shown in Figure 2.

Third, the empirical study of TARP-related capital injections by Berrospide and Edge (2010) attributed the US slowdown in loan growth after the capital injections to the US recession and banks' accompanying increased perception of riskiness of lending, but not to their capital position. Our estimation results provide support for Berrospide and Edge's (2010) analysis of the US capital injections.

In sum, our estimation results of Model III do not indicate that the policy framework of capital injection that is subject to capital adequacy requirements impeded the lending of the capital-injected banks. On the contrary, our results indicate that lenders' increased perception of the riskiness of lending due to borrowers' increases in default risk and deterioration of profitability were primarily responsible for impeding the lending not only of the injected banks, but also of the noninjected banks during the sample period after the two public capital injections.

5. Conclusion This paper draws three main substantive conclusions.

First, the first and second capital injections reduced the default risks of the capital-injected banks and their nonperforming loans. We therefore conclude that the two public capital injections significantly reduced the financial risks of the capital-injected banks.

Second, the two injections did not improve the profitability of the capital-injected banks and their lending behavior substantially.

Third, the main reason why the lending of the capital-injected banks did not increase is most likely because the profitability of their borrowers deteriorated and the borrowers' default risks increased in the severe recession after the two injections. In addition, such borrowers' increased default risk and deterioration of profitability would impede not only the injected but also the noninjected banks from lending more. In other words, the deterioration of borrowers' creditworthiness due to Japan's severe recession and banks' accompanying increased perception of the riskiness of lending would impede overall bank lending to domestic enterprises after the two public injections.

The two capital injections in Japan probably had a favorable effect in terms of decreasing the financial risks of the capital-injected banks. Such a favorable effect is likely to have stabilized greatly the Japanese banking system. On the other hand, the public capital injections would not have successfully stimulated the lending and the profitability of the injected banks.

Public capital injection is a comprehensive policy program to reduce the financial risk of a capital-injected bank and thereby stimulate its lending and profitability. However, given that a large-scale capital injection, such as Japan's first and second capital injections in 1998 and 1999, is usually implemented in a severe recession, the implementation of such a large-scale injection may not lead to improvement in the lending and the profitability of the injected banks; in other words, it may work only "to stabilize the banking system", that is, "to save banks".

One significant aspect not addressed in the paper involves controlling for borrower factors in the loan supply function; our data set used to investigate the determinants of the lending of Japanese banks does not fully cover SMEs as borrowers because the data set only includes domestic listed companies. Our analysis of the loan supply function should be complemented along this line.

Appendix I: Illustrating the Treatment Effect This appendix examines the third equality in equation (2), which specifies the treatment effect δ_t in Model II. In the following, let us suppose that capital injection toward bank i is implemented at time t^* and cast the probability of default as its outcome variable y_{it}. Figure 7 illustrates the treatment effect δ_t with time $t^* - 1$ serving as the reference time.

Figure 7 shows that there are two possible paths of the probability of default of bank i. One is the path shown as the bold dashed line, along which the probability of default at time $t = t^* + k$ $(k \geq 0)$, y^A_{1i,t^*+k}, becomes lower than that at the reference time $t^* - 1$, y_{it^*-1}. The other is the path shown as the thinner dashed line, along which the probability of default at time t, y^B_{1i,t^*+k}, becomes higher than that at the reference time. Furthermore, we suppose that if bank i has not received the capital injection, the probability of default would take the path shown as the solid line; the probability of default at time t, $y_{0i,t}$, would become higher than y^B_{1i,t^*+k}.

Figure 7 indicates that if the probability of default follows the bold dashed path, the treatment effect of the capital injection δ_t is described as follows:

$$\delta^A_t = y^A_{1i,t^*+k} - y_{0i,t} = (y^A_{1i,t^*+k} - y_{it^*-1}) - (y_{0i,t} - y_{it^*-1}).$$

This equation corresponds to the third equality in equation (2), which specifies an estimate of the treatment effect δ_t as a difference-in-difference estimate. That is to say, the difference between the actual (the first parenthetic term) and counterfactual variation (the second parenthetic term) measures the treatment effect of the capital injection.

On the other hand, if the probability of default follows the thinner dashed line, the treatment effect is described as follows:

$$\delta^B_t = y^B_{1i,t^*+k} - y_{0i,t} = (y^B_{1i,t^*+k} - y_{it^*-1}) - (y_{0i,t} - y_{it^*-1}).$$

This equation implies that if an outcome variable of an agent that received a policy program follows the thinner dashed path, casual observation of the actual variation of the outcome variable (first parenthetic term) leads to a wrong judgment on the effect of the policy program. Therefore, the accurate evaluation of a policy program including a capital injection program requires estimating its treatment effect by causal inference.

Appendix II: Statistical Inference of the Estimated Treatment Effect Conley and Taber (2011) demonstrated that standard large-sample theory is not appropriate for statistical inference of the treatment effect estimated using the within-group estimation method for a fixed effect panel model when the number of members of the treated group, N_1, is much smaller than that of the control group, N_0. Accordingly, Conley and Taber (2011) suggested an alternative method for statistical inference that employs the information of members of the control group. More precisely, their method is based on the empirical distribution constructed using residuals generated from a control group equation in a fixed effect panel model. Following Conley and Taber's (2011) method, we then conduct statistical inference based on the empirical distribution constructed by the following procedures:

1. Estimate Model I using the within-group estimation method.

2. Generate residuals ε_{jt} from an estimated equation for bank j ($j = 1, \ldots, N_0$) that belongs to the control group, and then calculate the centered residuals $\tilde{\varepsilon}_{jt} = \varepsilon_{jt} - \bar{\varepsilon}_j$. N_0 denotes the number of banks belonging to the control group.

3. Construct the empirical distribution of the estimated treatment effect using the centered residuals as follows:

$$\frac{\sum_{i=1}^{N_1} \sum_{t=t^*-1}^{T} (D_{it} - \bar{D}_i)\tilde{\varepsilon}_{jt}}{\sum_{i=1}^{N_1} \sum_{t=t^*-1}^{T} (D_{it} - \bar{D}_i)^2} \quad (j = 1, \ldots, N_0),$$

   where $\bar{D}_i = (T - t^*)^{-1} \sum_{t=t^*-1}^{T} D_{it}$, and T indicates the endpoint of the sample period. The 95% confidence intervals of treatment effect δ reported in Tables 2 to 5 are obtained as "a point estimate of δ plus the 2.5 percentage quantile of the empirical distribution" and "the point estimate plus the 97.5 percentage quantile".

4. To test the null hypothesis $\delta = 0$, estimate Model I, imposing the parameter restriction $\delta = 0$, and then construct the empirical distribution of the null hypothesis following the above procedures 2 and 3. If a point estimate of δ falls into the rejection region of this empirical distribution at the required level of significance, reject the null. Tables 2 to 5 report the 1, 5 and 10 percent levels of significance with the corresponding number of asterisks.

5. The 90% confidence intervals of the treatment effect δ_t at each time $t = t^* + k$ $(k \geq 0)$, shown in Figures 5 and 6, are constructed by modifying the above procedures in the following way. First, in procedures 1 and 2, we estimate Model II and generate centered residuals $\tilde{\varepsilon}_{jt} = \varepsilon_{jt} - \bar{\varepsilon}_j$ for bank j $(j = 1, \ldots, N_0)$ that belongs to the control group. Next, in procedure 3, using the centered residuals, we calculate the empirical distribution at each time $t = t^* + k$ $(k \geq 0)$ as follows:

$$\frac{\sum_{i=1}^{N_1}(D_{it} - \bar{D}_i)\tilde{\varepsilon}_{jt}}{\sum_{i=1}^{N_1}(D_{it} - \bar{D}_i)^2} \quad (j = 1, \ldots, N_0).$$

   Finally, the 90% confidence intervals of the treatment effect δ_t at time $t = t^* + k$ $(k \geq 0)$ are obtained as "a point estimate of δ_t at time $t = t^* + k$ $(k \geq 0)$ plus the five percentage quantile of the empirical distribution" and "the point estimate plus the 95 percentage quantile".

Appendix III: The Heterogeneous Effect on the Probability of Default The main text of this paper observed that the two capital injections in 1998 and 1999 decreased the default risks of capital-injected banks. This appendix introduces the heterogeneous treatment effect corresponding to the amount of capital injection into each bank, and thereby examines how the amount of capital injection affected the default risks of capital-injected banks.

More precisely, we identify such a heterogeneous effect using the following model:

$$\textbf{Model IV}: \quad PD_{it} = \alpha PD_{it-1} + \beta' \mathbf{X_{it}} + \gamma_t t + \delta_q(D_{it}^q) + v_i + \varepsilon_{it},$$

where dummy variable D_{it}^q is set to each bank i depending on the amount of capital injection, and hence its parameter coefficient δ_q captures heterogeneity in the policy effect corresponding to the amount of the injection. PD_{it} and $\mathbf{X_{it}}$ indicate the probability of default and covariates other than PD_{it-1}, respectively. t denotes time dummy variables and v_i denotes the fixed effect term of each bank. For covariate $\mathbf{X_{it}}$, this appendix uses the one-period lags of NPL_{it}, ROA_{it}, LOAN_{it}, SMELOAN_{it} and SIZE_{it}. For estimation of Model IV, we use a two-stage within-group estimation method using PD_{it-2} as an instrumental variable for PD_{it-1}. In addition, to confirm the robustness of the estimation results, we also estimate Model IV, which does not include PD_{it-1} as an explanatory variable.

Table 7 shows the estimation results of the heterogeneous effect corresponding to the amount of the injection. [22] At first, we report the heterogeneous effect of the first capital injection in 1998. As shown in Table 1, 100 billion yen was injected into 11 out of 21 banks, and hence the first capital injection is often characterized as the "yokonarabi (herd behavior) policy". Nevertheless, we observe that the injection of 100 billion yen significantly reduced PD_{it}. Given the significant effect of the injection of 100 billion yen, it is inferred that the overall effect of the first capital injection, discussed in Section 3, primarily reflects this 100 billion yen injection. On the other hand, Table 7 indicates that the first capital injection did not reduce the default risks of the Long-term Credit Bank of Japan and the Nippon Credit Bank, the former having received the largest capital injection of 176.6 billion, while the latter received 60 billion yen. The estimation results for the two banks are consistent with the fact that the two banks fell into bankruptcy after the first capital injection. In sum, our estimates of the heterogeneous effect imply that for the first capital injection, the difference in the amount of injection into each bank did not make any qualitative difference to the amount by which its default risk was reduced.

[22] The 95% confidence intervals related to the treatment effect δ_q reported in Table 7 are constructed by modifying procedures 1 to 3 in Appendix II. First, we obtain centered residuals $\tilde{\varepsilon}_{jt}$ $(j = 1, \ldots, N_0)$ of the control group by estimating Model IV. Next, we construct the empirical distribution using the following equation:

$$\frac{\sum_{i=1}^{N_1} \sum_{t=t^*-1}^{T} (D_{it}^q - \bar{D}_i^q)\tilde{\varepsilon}_{jt}}{\sum_{i=1}^{N_1} \sum_{t=t^*-1}^{T} (D_{it}^q - \bar{D}_i^q)^2} \quad (j = 1, \ldots, N_0),$$

where $\bar{D}_i^q = (T - t^*)^{-1} \sum_{t=t^*-1}^{T} D_{it}^q$. The confidence intervals are obtained as "a point estimate of δ_q plus the 2.5 percentage quantile of the empirical distribution", and "the point estimate plus the 97.5 percentage quantile". To test the null hypothesis of δ_q, we estimate Model IV and impose the restriction $\delta_q = 0$, and then construct the empirical distribution of the null hypothesis.

Next, we report the heterogeneous effect of the second capital injection in 1999. As shown in Table 7, the second injection significantly reduced PD_{it} in more cases than the first injection. Such a favorable result for the second injection may be because of the fact that the injected banks' capital was initially adequate. Unlike the first injection in 1998, the second injection was conducted after a bank stress test to determine the injected bank's capital requirements (see, e.g., Allen et al. (2009) and Hoshi and Kashyap (2010)). On the other hand, although Daiwa Bank and the Asahi Bank, which later merged into the Resona Bank, received capital injections of 408 billion and 500 billion yen, respectively (see Table 1), the second capital injection did not significantly reduce their default risks. Finally, for the three banks that received the largest capital injections, the Sakura Bank (800 billion yen), the Daiichi Kangyo Bank (900 billion yen) and the Fuji Bank (1,000 billion yen), Table 7 indicates that the second capital injection reduced their default risks significantly. In sum, our empirical results for the second capital injection imply that as long as a bank stress test to determine the injected bank's capital requirements was fully conducted, the difference in the size of the capital injection of each bank possibly caused a quantitative difference in the amount by which its default risk was reduced.

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Table 1: The size of public capital injections

Bank name	The first capital injection based on the Financial Function Stabilization Act			The second capital injection based on the Prompt Recapitalization Act		
	Preferred shares	Subordinated bonds and loans	Total	Preferred shares	Subordinated bonds and loans	Total
Daiichi Kangyo	99	-	99	700	200	900
Fuji	-	100	100	800	200	1000
Industrial Bank of Japan	-	100	100	350	250	600
Yasuda Trust	-	150	150	-	-	-
Sakura	-	100	100	800	-	800
Sumitomo	-	100	100	501	-	501
Tokyo Mitsubishi	-	100	100	-	-	-
Mitsubishi Trust	-	50	50	200	100	300
Sanwa	-	100	100	600	100	700
Tokai	-	100	100	600	-	600
Toyo Trust	-	50	50	200	-	200
Asahi	-	100	100	400	100	500
Daiwa	-	100	100	408	-	408
Sumitomo Trust	-	100	100	100	100	200
Mitsui Trust	-	100	100	250.3	150	400.3
Chuo Trust	32	28	60	150	-	150
Yokohama	-	20	20	100	100	200
Hokuriku	-	20	20	-	-	-
Ashikaga	-	30	30	-	-	-
Long-Term Credit Bank of Japan	130	46.6	176.6	-	-	-
Nippon Credit Bank	60	-	60	-	-	-
Total	321	1494.6	1815.6	6159.3	1300	7459.3

1. Data are expressed in billions of yen.

Table 2: Estimation results of Model I based on the probability of default: Two-stage within-group estimation method

The first capital injection (September 1997 - September 1998)

	Outcome variable: y_t^i				
	PD	NPL	ROA	Loan	SMELOAN
Treatment effect: δ	-2.692** (-5.138, -0.246)	-6.305 (-15.92, 3.309)	-0.046 (-0.120, 0.028)	0.484 (-1.525, 2.493)	0.043 (-1.730, 1.816)
y_{t-1}^i	0.132 (-0.178, 0.442)	0.420 (-0.266, 1.106)	0.870*** (0.652, 1.088)	0.701*** (0.315, 1.087)	0.251 (-1.307, 1.809)
PD_{t-1}^i	-	0.002 (-0.002, 0.006)	-0.001 (-0.005, 0.003)	-0.0001 (-0.0007, 0.0004)	-0.0003 (-0.0012, 0.0006)
NPL_{t-1}^i	0.016* (-0.002, 0.034)	-	-0.089** (-0.160, -0.018)	-0.016 (-0.071, 0.039)	-0.021 (-0.058, 0.016)
ROA_{t-1}^i	-0.323 (-4.713, 4.067)	-0.045 (-0.280, 0.190)	-	0.030 (-0.011, 0.071)	0.044* (-0.001, 0.089)
$LOAN_{t-1}^i$	20.66 (-14.992, 56.312)	0.428 (-1.440, 2.296)	-0.003 (-0.767, 0.761)	-	0.058 (-0.961, 1.077)
$SMELOAN_{t-1}^i$	3.972 (-35.836, 43.780)	0.965 (-1.093, 3.023)	-0.413* (-0.878, 0.052)	1.032 (-0.850, 2.914)	-
$SIZE_{t-1}^i$	-8.221 (-37.327, 20.885)	1.197 (-0.298, 2.692)	-0.213 (-0.576, 0.150)	0.330 (-0.111, 0.771)	0.186 (-0.498, 0.870)

The second capital injection (September 1998 - March 2002)

	Outcome variable: y_t^i				
	PD	NPL	ROA	Loan	SMELOAN
Treatment effect: δ	-0.937* (-1.939, -0.007)	-18.34*** (-31.84, -4.842)	0.073 (-0.143, 0.289)	-2.084 (-5.041, 0.873)	1.068 (-9.011, 11.15)
y_{t-1}^i	0.754 (-0.197, 1.705)	0.375*** (0.248, 0.502)	0.915*** (0.905, 0.925)	0.982*** (0.802, 1.162)	0.174* (-0.006, 0.354)
PD_{t-1}^i	-	0.009** (0.001, 0.018)	-0.0030* (-0.0061, 0.0001)	-0.001 (-0.011, 0.009)	-0.013 (-0.046, 0.020)
NPL_{t-1}^i	0.246* (-0.007, 0.499)	-	-0.005 (-0.019, 0.009)	-0.003 (-0.001, 0.006)	-0.027 (-0.078, 0.024)
ROA_{t-1}^i	-14.12*** (-19.849, -8.391)	-0.784*** (-1.319, -0.249)	-	0.072 (-0.081, 0.225)	0.070 (-0.326, 0.466)
$LOAN_{t-1}^i$	4.660 (-4.666, 13.986)	0.884** (0.125, 1.643)	-0.103 (-0.242, 0.036)	-	-1.117 (-3.191, 0.957)
$SMELOAN_{t-1}^i$	3.448 (-0.882, 7.778)	0.012 (-0.215, 0.239)	-0.069** (-0.118, -0.020)	0.045*** (0.016, 0.074)	-
$SIZE_{t-1}^i$	1.793 (-1.794, 5.380)	0.057 (-0.190, 0.304)	-0.036* (-0.077, 0.005)	0.064*** (0.037, 0.091)	0.153 (-0.082, 0.388)

1. We conduct a two-stage within-group estimation using y_{t-2}^i as an instrumental variable for y_{t-1}^i.
2. For the estimates of the treatment effect δ, the 95% confidence intervals calculated using Conley and Taber's (2011) method are in parentheses. See Appendix II for Conley and Taber's (2011) method. For the estimates of the covariates, the 95% confidence intervals calculated based on the large-sample approximation are in parentheses.
3. *, ** and *** indicate the 10%, 5% and 1% levels of significance, respectively.

Table 3: Estimation results of Model I based on the probability of default : Within-group estimation method

The first capital injection (September 1997 - September 1998)

	Outcome variable : y_t^i PD	NPL	ROA	LOAN	SMELOAN
Treatment effect : δ	-2.615* (-5.506, 0.076)	-8.279* (-18.19, 1.638)	-0.989 (-3.305, 1.372)	0.401 (-0.966, 1.768)	0.029 (-2.901, 2.959)
y_{t-1}^i	-	-	-	-	-
PD_{t-1}^i	-	0.001 (-0.003, 0.005)	-0.004 (-0.016, 0.008)	-0.00009 (-0.0005, 0.0003)	-0.0003 (-0.0007, 0.0001)
NPL_{t-1}^i	0.328 (-3.753, 4.409)	-	-0.470** (-0.886, -0.054)	-0.001 (-0.027, 0.026)	-0.021 (-0.060, 0.018)
ROA_{t-1}^i	-0.062 (-1.132, 1.008)	0.015 (-0.063, 0.093)	-	0.020*** (0.012, 0.028)	0.044*** (0.028, 0.060)
$LOAN_{t-1}^i$	19.480 (-23.503, 62.463)	0.799 (-0.892, 2.490)	-6.108 (-17.676, 5.460)	-	0.481** (0.079, 0.883)
$SMELOAN_{t-1}^i$	3.457 (-32.685, 39.599)	1.383 (-0.406, 3.172)	-3.962 (-14.683, 6.759)	0.473*** (0.250, 0.696)	-
$SIZE_{t-1}^i$	-6.832 (-26.258, 12.594)	1.115 (-0.263, 2.493)	-4.092 (-9.752, 1.568)	0.206** (0.030, 0.382)	0.207 (-0.232, 0.646)

The second capital injection (September 1998 - March 2002)

	Outcome variable : y_t^i PD	NPL	ROA	LOAN	SMELOAN
Treatment effect : δ	-1.137** (-2.209, -0.065)	-25.93** (-50.67, -1.195)	0.039 (-0.036, 0.114)	-1.326 (-4.263, 1.611)	-2.005 (-7.851, 3.841)
y_{t-1}^i	-	-	-	-	-
PD_{t-1}^i	-	0.010** (0.002, 0.018)	-0.010 (-0.026, 0.006)	-0.002 (-0.005, 0.001)	-0.001 (-0.005, 0.002)
NPL_{t-1}^i	0.166** (0.021, 0.311)	-	-0.319*** (-0.468, -0.170)	-0.014 (-0.049, 0.021)	-0.023 (-0.050, 0.004)
ROA_{t-1}^i	-12.70*** (-19.732, -5.668)	-1.401*** (-2.083, -0.719)	-	0.072 (-0.026, 0.170)	0.050 (-0.097, 0.197)
$LOAN_{t-1}^i$	1.209 (-3.185, 5.603)	1.160** (0.009, 2.311)	-0.167 (-1.723, 1.389)	-	0.775*** (0.555, 0.995)
$SMELOAN_{t-1}^i$	0.596 (-0.402, 1.594)	0.034 (-0.142, 0.210)	-0.148 (-1.467, 1.171)	0.051 (-0.023, 0.125)	-
$SIZE_{t-1}^i$	0.151 (-1.519, 1.821)	-0.159 (-0.445, 0.127)	-0.202 (-0.496, 0.092)	-0.001 (-0.040, 0.038)	-0.003 (-0.081, 0.075)

1. We conduct within-group estimation.
2. For the estimates of the treatment effect δ, the 95% confidence intervals calculated using Conley and Taber's (2011) method are in parentheses. See Appendix II for Conley and Taber's (2011) method. For the estimates of the covariates, the 95% confidence intervals calculated based on large-sample approximation are in parentheses.
3. *, ** and *** indicate the 10%, 5% and 1% levels of significance, respectively.

Table 4: Estimation results of Model I based on the Tier I ratio：Two-stage within-group estimation method

The first capital injection（September 1997 - September 1998）

	Outcome variable：y_t^i TIER	NPL	ROA	LOAN	SMELOAN
Treatment effect：δ	1.200* (0.023, 2,446)	-4.263* (-9.018, 0.492)	-0.050 (-0.106, 0.006)	-0.749 (-2.545, 1.047)	-0.513 (-2.427, 1.401)
y_{t-1}^i	0.402*** (0.200, 0.604)	0.382 (-0.271, 1.035)	0.866*** (0.844, 0.888)	0.660*** (0.286, 1.034)	0.269 (-1.499, 2.037)
TIER_{t-1}^i	-	-0.070* (-0.141, 0.001)	0.030* (-0.005, 0.065)	0.010 (-0.019, 0.039)	-0.017 (-0.039, 0.005)
NPL_{t-1}^i	-1.734* (-3.504, 0.036)	-	-0.089** (-0.160, -0.018)	-0.019 (-0.072, 0.034)	-0.016 (-0.053, 0.021)
ROA_{t-1}^i	2.980** (1.694, 4.266)	-0.471* (-0.996, 0.054)	-	0.092*** (0.039, 0.145)	0.152** (0.023, 0.281)
LOAN_{t-1}^i	-15.08* (-31.064, 0.904)	0.639 (-1.166, 2.444)	-0.058 (-0.836, 0.720)	-	0.573 (-0.505, 1.651)
SMELOAN_{t-1}^i	-3.046 (-29.780, 23.688)	1.703 (-0.375, 3.781)	-0.467* (-0.941, 0.007)	0.917 (-1.018, 2.852)	-
SIZE_{t-1}^i	20.98** (0.126, 41.834)	1.765** (0.150, 3.380)	-0.392 (-0.772, -0.012)	0.246 (-0.252, 0.744)	0.022 (-0.817, 0.861)

The second capital injection（September 1998 - March 2002）

	Outcome variable：y_t^i TIER	NPL	ROA	LOAN	SMELOAN
Treatment effect：δ	1.533*** (0.904, 2.162)	-14.22** (-27.98, -3.692)	0.028 (-0.309, 0.286)	-1.644* (-3.123, 0.089)	2.848 (-3.506, 9.151)
y_{t-1}^i	0.104 (-0.149, 0.357)	0.372*** (0.243, 0.501)	0.915*** (0.905, 0.925)	0.998*** (0.800, 1.196)	0.665*** (0.557, 0.773)
TIER_{t-1}^i	-	-0.017 (-0.048, 0.014)	0.033 (-0.016, 0.082)	0.001 (-0.003, 0.005)	-0.004 (-0.028, 0.020)
NPL_{t-1}^i	-0.226* (-0.465, 0.013)	-	-0.010 (-0.024, 0.004)	-0.003 (-0.011, 0.005)	-0.027 (-0.076, 0.022)
ROA_{t-1}^i	4.689*** (2.404, 6.974)	-0.670** (-1.231, -0.109)	-	0.063 (-0.017, 0.143)	0.115 (-0.310, 0.540)
LOAN_{t-1}^i	-5.760*** (-8.577, -2.943)	1.002** (0.226, 1.778)	-0.161 (-0.304, -0.018)	-	1.004 (-0.891, 2.899)
SMELOAN_{t-1}^i	-0.312 (-1.071, 0.447)	0.020 (-0.205, 0.245)	-0.035** (-0.086, 0.016)	0.050** (0.021, 0.079)	-
SIZE_{t-1}^i	-0.242 (-1.057, 0.573)	0.024 (-0.225, 0.273)	-0.004* (-0.047, 0.039)	0.060*** (0.031, 0.089)	0.153 (-0.082, 0.388)

1. We conduct two-stage within-group estimation using y_{t-2}^i as an instrumental variable for y_{t-1}^i.
2. For the estimates of the treatment effect δ, the 95% confidence intervals calculated using Conley and Taber's (2011) method are in parentheses. See Appendix II for Conley and Taber's (2011) method. For the estimates of the covariates, the 95% confidence intervals calculated based on the large-sample approximation are in parentheses.
3. *, ** and *** indicate the 10%, 5% and 1% levels of significance, respectively.

Table 5: Estimation results of Model I based on the capital ratio：Two-stage within-group estimation method

The first capital injection（September 1997 - September 1998）

	Outcome variable：y_t^i RATIO	NPL	ROA	LOAN	SMELOAN
Treatment effect：δ	1.598*** (0.749, 2.447)	-5.089* (-9.964, -0.359)	0.011 (-0.010, 0.004)	-0.470 (-2.423, 1.593)	-0.210 (-2.521, 1.559)
y_{t-1}^i	0.158*** (0.105, 0.211)	0.402 (-0.268, 1.072)	0.870*** (0.848, 0.892)	0.164 (-0.355, 0.683)	0.233 (-1.184, 1.650)
RATIO_{t-1}^i	-	-0.038 (-0.093, 0.017)	0.021* (-0.011, 0.043)	0.013 (-0.016, 0.042)	-0.008 (-0.020, 0.004)
NPL_{t-1}^i	-1.304** (-2.458, -0.150)	-	-0.089** (-0.160, -0.018)	-0.045 (-0.182, 0.092)	-0.014 (-0.049, 0.021)
ROA_{t-1}^i	2.284*** (1.751, 2.817)	-0.293 (-0.765, 0.179)	-	0.130 (-0.144, 0.404)	0.101*** (0.025, 0.177)
LOAN_{t-1}^i	-19.82*** (-31.933, -7.707)	0.143 (-1.821, 2.107)	-0.003 (-0.767, 0.761)	-	0.331 (-0.508, 1.170)
SMELOAN_{t-1}^i	-11.750 (-28.892, 5.392)	1.364 (-0.674, 3.402)	-0.413* (-0.878, 0.052)	1.568 (-2.109, 5.245)	-
SIZE_{t-1}^i	20.47** (4.692, 36.248)	1.852** (0.006, 3.698)	-0.213 (-0.576, 0.150)	0.253 (-0.186, 0.692)	0.110 (-0.611, 0.831)

The second capital injection（September 1998 - March 2002）

	Outcome variable：y_t^i RATIO	NPL	ROA	LOAN	SMELOAN
Treatment effect：δ	1.728*** (1.146, 2.490)	-14.05** (-27.86, -1.237)	-0.013 (-0.044, 0.028)	-1.588 (-3.078, 0.158)	2.041 (-6.448, 9.534)
y_{t-1}^i	0.044 (-0.240, 0.328)	0.377*** (0.250, 0.504)	0.915*** (0.905, 0.925)	0.100 (-0.102, 0.302)	0.166* (-0.003, 0.335)
RATIO_{t-1}^i	-	-0.013 (-0.038, 0.012)	0.033 (-0.496, 0.562)	0.001 (-0.003, 0.005)	-0.005 (-0.025, 0.015)
NPL_{t-1}^i	-0.291** (-0.561, -0.021)	-	-0.005 (-0.019, 0.009)	-0.003 (-0.011, 0.005)	-0.026 (-0.075, 0.023)
ROA_{t-1}^i	0.300** (0.051, 0.549)	-0.672** (-1.235, -0.109)	-	0.061 (-0.019, 0.141)	0.122 (-0.307, 0.551)
LOAN_{t-1}^i	-6.993*** (-10.527, -3.459)	1.014** (0.236, 1.792)	-0.103 (-0.242, 0.036)	-	0.997 (-0.898, 2.892)
SMELOAN_{t-1}^i	-0.025 (-0.903, 0.853)	0.019 (-0.206, 0.244)	-0.069** (-0.118, -0.020)	0.050*** (0.021, 0.079)	-
SIZE_{t-1}^i	-0.645 (-1.588, 0.298)	0.041 (-0.206, 0.288)	-0.036* (-0.077, 0.005)	0.062*** (0.033, 0.091)	0.121 (-0.091, 0.333)

1. We conduct two-stage within-group estimation using y_{t-2}^i as an instrumental variable for y_{t-1}^i.
2. For the estimates of the treatment effect δ, the 95% confidence intervals calculated using Conley and Taber's (2011) method are in parentheses. See Appendix II for Conley and Taber's (2011) method. For the estimates of the covariates, the 95% confidence intervals calculated based on the large-sample approximation are in parentheses.
3. *, ** and *** indicate the 10%, 5% and 1% levels of significance, respectively.

Table 6: Estimation results of the loan supply function: Model III (March 1998 - March 2002)

The adjusted capital ratio

Dependent variable	$LOAN_t^{i,j}$	Capital-injected bank's loan			Noncapital-injected bank's loan		
Autoregressive variable	$LOAN_{t-1}^{i,j}$	0.508*** (0.494, 0.522)	0.504*** (0.490, 0.518)	0.504*** (0.490, 0.518)	0.460*** (0.450, 0.470)	0.458*** (0.448, 0.468)	0.458*** (0.448, 0.468)
Factor of bank i	$ADRATIO_{t-1}^{i}$	0.004 (-0.010, 0.018)	0.001 (-0.013, 0.015)	-0.004 (-0.016, 0.008)	0.002 (-0.004, 0.008)	0.003 (-0.007, 0.013)	0.002 (-0.008, 0.012)
	NPL_{t-1}^{i}	-	-	-0.064 (-0.146, 0.018)	-	-	-0.003 (-0.046, 0.040)
	ROA_{t-1}^{i}	-	0.008 (-0.006, 0.022)	0.009 (-0.005, 0.023)	-	0.001 (-0.001, 0.003)	0.001 (-0.001. 0.003)
Factor of borrower j	PD_{t-1}^{j}	-0.005* (-0.011, 0.001)	-0.012*** (-0.018, -0.006)	-0.012*** (-0.018, -0.006)	-0.0040* (-0.0079, -0.0001)	-0.010*** (-0.016, -0.004)	-0.010*** (-0.016, -0.004)
	ROA_{t-1}^{j}	-	0.003*** (0.002, 0.004)	0.003*** (0.002, 0.004)	-	0.002*** (0.001, 0.003)	0.002*** (0.001, 0.003)

The probability of default

Dependent variable	$LOAN_t^{i,j}$	Capital-injected bank's loan			Noncapital-injected bank's loan		
Autoregressive variable	$LOAN_{t-1}^{i,j}$	0.508*** (0.494, 0.522)	0.504*** (0.490, 0.518)	0.504*** (0.490, 0.518)	0.451*** (0.439, 0.463)	0.448*** (0.436, 0.460)	0.449*** (0.437, 0.461)
Factor of bank i	PD_{t-1}^{i}	-0.001 (-0.003, 0.001)	-0.001 (-0.002, 0.001)	-0.0004 (-0.0024, 0.0016)	-0.001 (-0.007, 0.005)	-0.002 (-0.008, 0.004)	-0.0015 (-0.0035, 0.0005)
	NPL_{t-1}^{i}	-	-	-0.060 (-0.140, 0.020)	-	-	-0.010 (-0.053, 0.033)
	ROA_{t-1}^{i}	-	0.008 (-0.006, 0.022)	0.009 (-0.005, 0.023)	-	0.001 (-0.001, 0.003)	0.001 (-0.001, 0.003)
Factor of borrower j	PD_{t-1}^{j}	-0.0060* (-0.0119, -0.0001)	-0.012*** (-0.018, -0.006)	-0.012*** (-0.018, -0.006)	-0.007*** (-0.011, -0.003)	-0.012*** (-0.018, -0.006)	-0.012*** (-0.018, -0.006)
	ROA_{t-1}^{j}	-	0.003*** (0.002, 0.004)	0.003*** (0.002, 0.004)	-	0.002*** (0.001, 0.003)	0.002*** (0.001, 0.003)

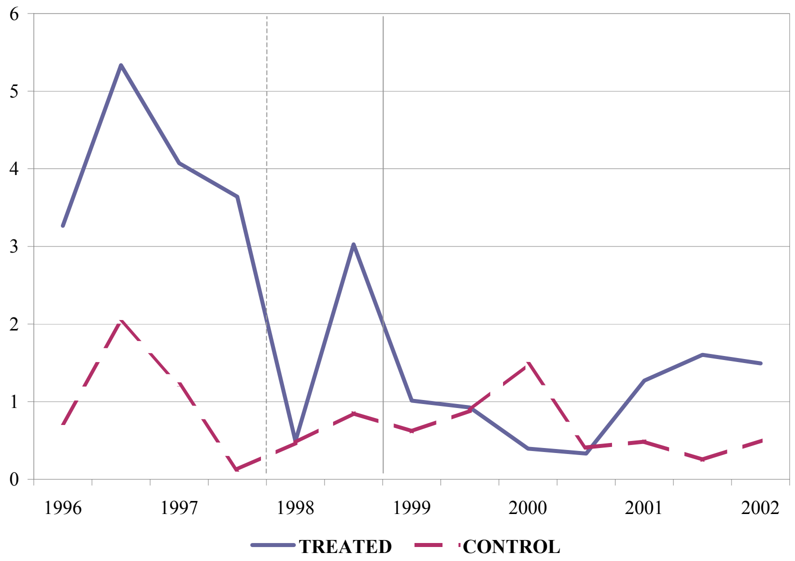
1. We employ the ordinary least squares estimation method for estimating Model III.
2. Estimates of the time dummy and bank dummy variables are not reported.
3. The adjusted capital ratio ($ADRATIO_{t-1}^{i}$ (s defined by subtracting the target capital ratio (eight percent for international banks and four percent for domestic banks) from a reported capital ratio.
4. The numbers in parentheses are the 95% confidence interval.
5. *, ** and *** indicate the 10%, 5% and 1% levels of significance, respectively.

Table 7: The heterogeneous treatment effect on the probability of default

The first capital injection			The second capital injection		
Size	Treatment effect : δ^q IV-FE estimation	FE estimation	Size	Treatment effect : δ^q IV-FE estimation	FE estimation
20 billion yen	-0.799 (-5.824, 3.226)	-0.452 (-1.472, 0.580)	150 billion yen	-0.309 (-6.753, 6.035)	-2.242*** (-3.324, -1.216)
30 billion yen	-0.757 (-7.654, 5.140)	-0.416 (-1.745, 0.813)	200 billion yen	-1.095 (-5.639, 2.339)	-1.466** (-2.705, -0.427)
50 billion yen	0.574 (-4.463, 7.155)	0.777* (-0.201, 1.755)	300 billion yen	-1.399** (-2.684, -0.118)	-1.350*** (-2.236, -0.474)
60 billion yen	-1.173 (-6.817, 2.472)	-0.717 (-2.521, 1.639)	400.3 billion yen	-3.382 (-10.51, 3.443)	-1.380*** (-2.248, -0.612)
99 billion yen	-0.405 (-7.274, 8.568)	-0.194 (-0.825, 0.289)	408 billion yen	2.006 (-4.287, 7.298)	4.820*** (4.352, 5.288)
100 billion yen	-4.509*** (-7.512, -2.211)	-4.604*** (-6.915, -2.836)	450 billion yen	-4.643 (-9.515, 0.229)	0.226 (-0.556,1.008)
150 billion yen	0.029 (-7.357, 7.415)	0.684 (-0.931, 1.899)	500 billion yen	1.487 (-4.710, 7.684)	3.797*** (3.327, 4.257)
176.6 billion yen	2.080*** (0.551, 3.250)	2.528*** (1.804, 3.251)	501 billion yen	-0.970* (-1.867, 0.026)	-0.564* (-1.207, 0.080)
-	-	-	600 billion yen	-1.753** (-3.356, -0.133)	-1.005** (-1.907, -0.104)
-	-	-	700 billion yen	-1.281 (-7.446, 4.884)	0.003 (-0.550, 0.555)
-	-	-	800 billion yen	-5.567* (-11.32, 0.369)	-2.165*** (-2.686, -1.654)
-	-	-	900 billion yen	-1.281 (-2.691, 0.310)	-2.163*** (-2.684, -1.643)
-	-	-	1000 billion yen	-6.930*** (-7.911, -5.901)	-2.670*** (-3.287, -2.013)

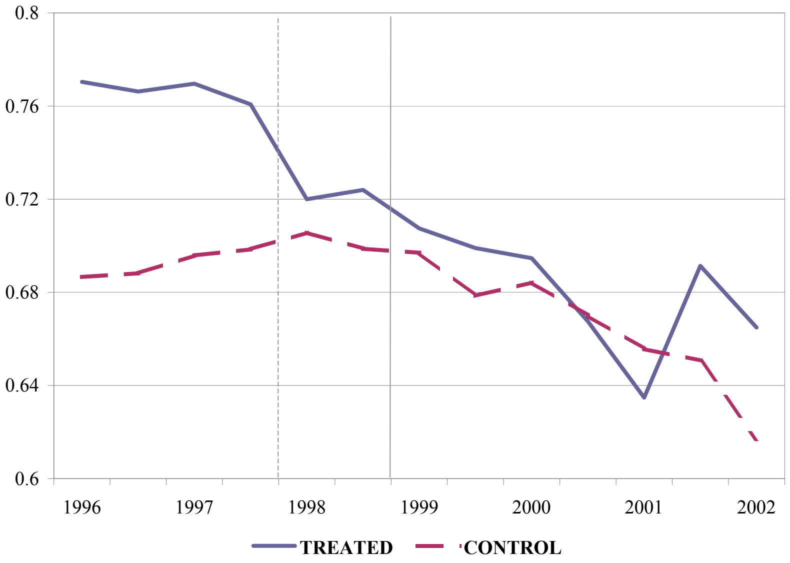
1. IV-FE indicates the two-stage within-group estimation using DP^i_{t-2} as an instrumental variable for DP^i_{t-1}. FE indicates the within-group estimation.
2. The numbers in parentheses are the 95% confidence interval calculated using Conley and Taber's (2011) method. See Appendix III for details.
3. *, ** and *** indicate the 10%, 5% and 1% levels of significance, respectively.

Figure 1: The Probability of Default of Japanese Banks



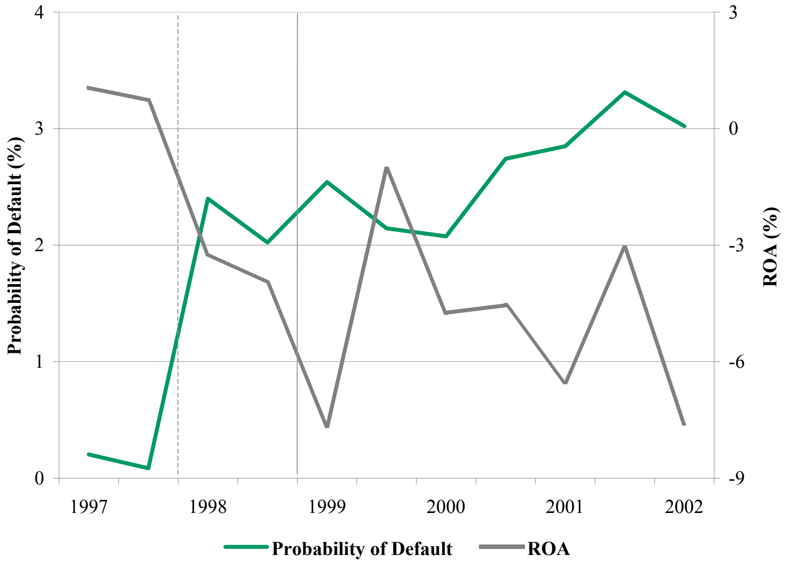
The vertical dotted line indicates the first injection period, and the vertical solid line indicates the second injection period. The solid line indicates the path of the injected banks (treated group), and the dashed line indicates that of the noninjected banks (control group). The probability default is calculated using Merton's (1974) structural model for option pricing. See Section 2 for more details.

Figure 2: Bank Loans to Domestic Enterprises



The vertical dotted line indicates the first injection period, and the vertical solid line indicates the second injection period. The solid line indicates the path of the injected banks (treated group), and the dashed line indicates that of the noninjected banks (control group). Bank loans is defined as the ratio of loans for domestic enterprises to total assets. See Section 2 for more details.

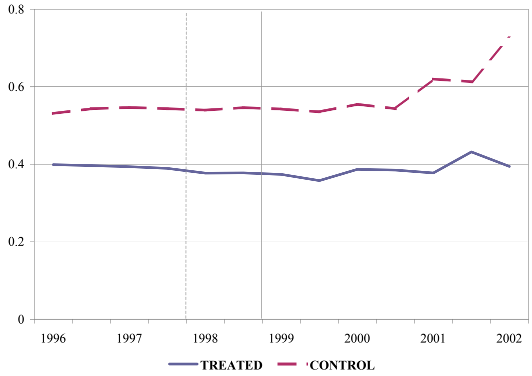
Figure 3: The Default Risk and Profitability of Borrowing Firms



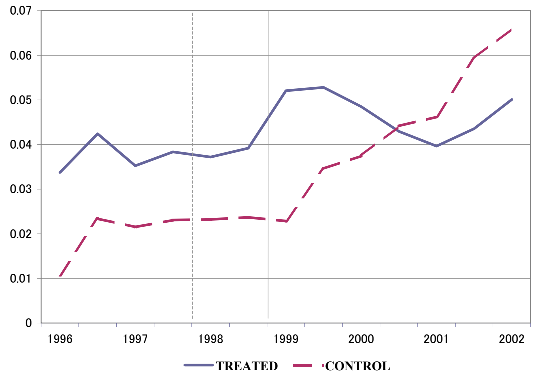
The vertical dotted line indicates the first injection period, and the vertical solid line indicates the second injection period. The probability of default of borrowing firms is calculated using Merton's (1974) structural model for option pricing. ROA (return on assets) is defined as $\frac{\text{net profits}}{\text{total assets}} \times 100$. See Section 3 for more details.

Figure 4: Historical Paths of Target Variables

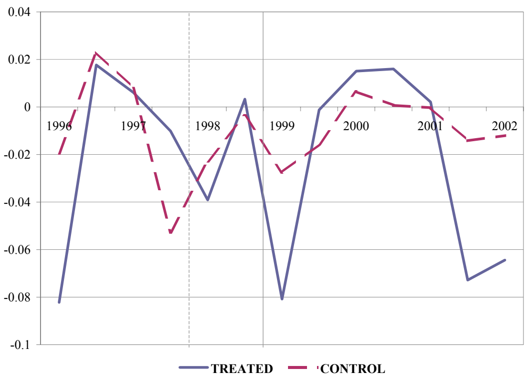
SME loans



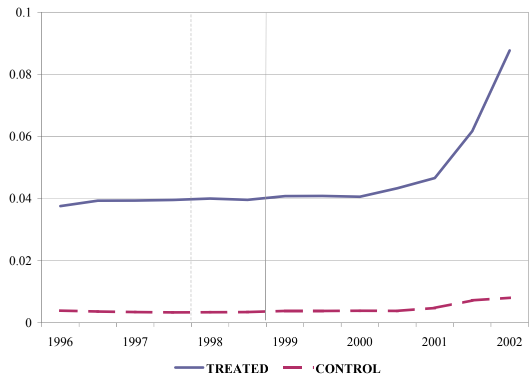
Nonperforming loans



Return on assets



Relative size



1. The vertical dotted line indicates the first injection period, and the vertical solid line indicates the second injection period.

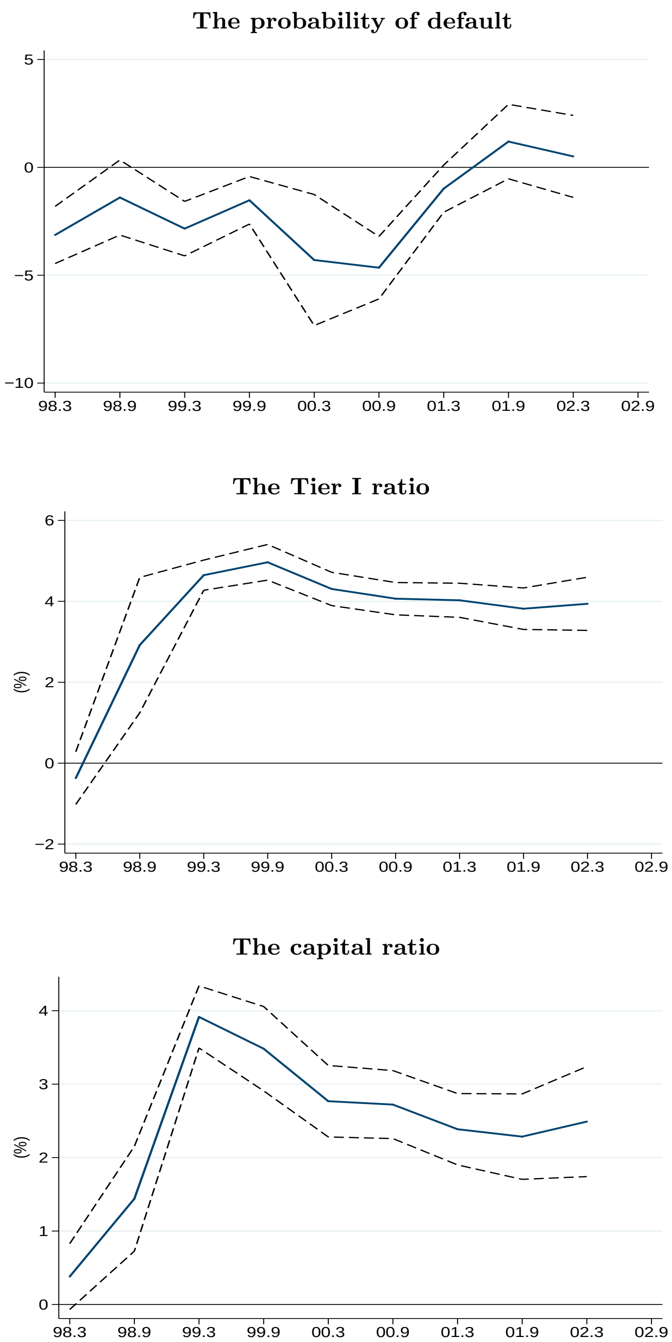
2. The solid line indicates the path of the injected banks (treated group), and the dashed line indicates that of the noninjected banks (control group).

3. SME loans and nonperforming loans are defined as the ratio of loans for small and medium enterprises to total assets and the ratio of nonperforming loans to total loans, respectively.

4. Return on assets is defined as $\frac{\text{net profits}}{\text{total assets}} \times 100$.

5. Relative size is defined as $V_{Ai}/\sum_{j=1}^{n} V_{Aj}$, where V_{Ai} is bank i's asset value and n is the number of banks listed on the Tokyo Stock Exchange at each time.

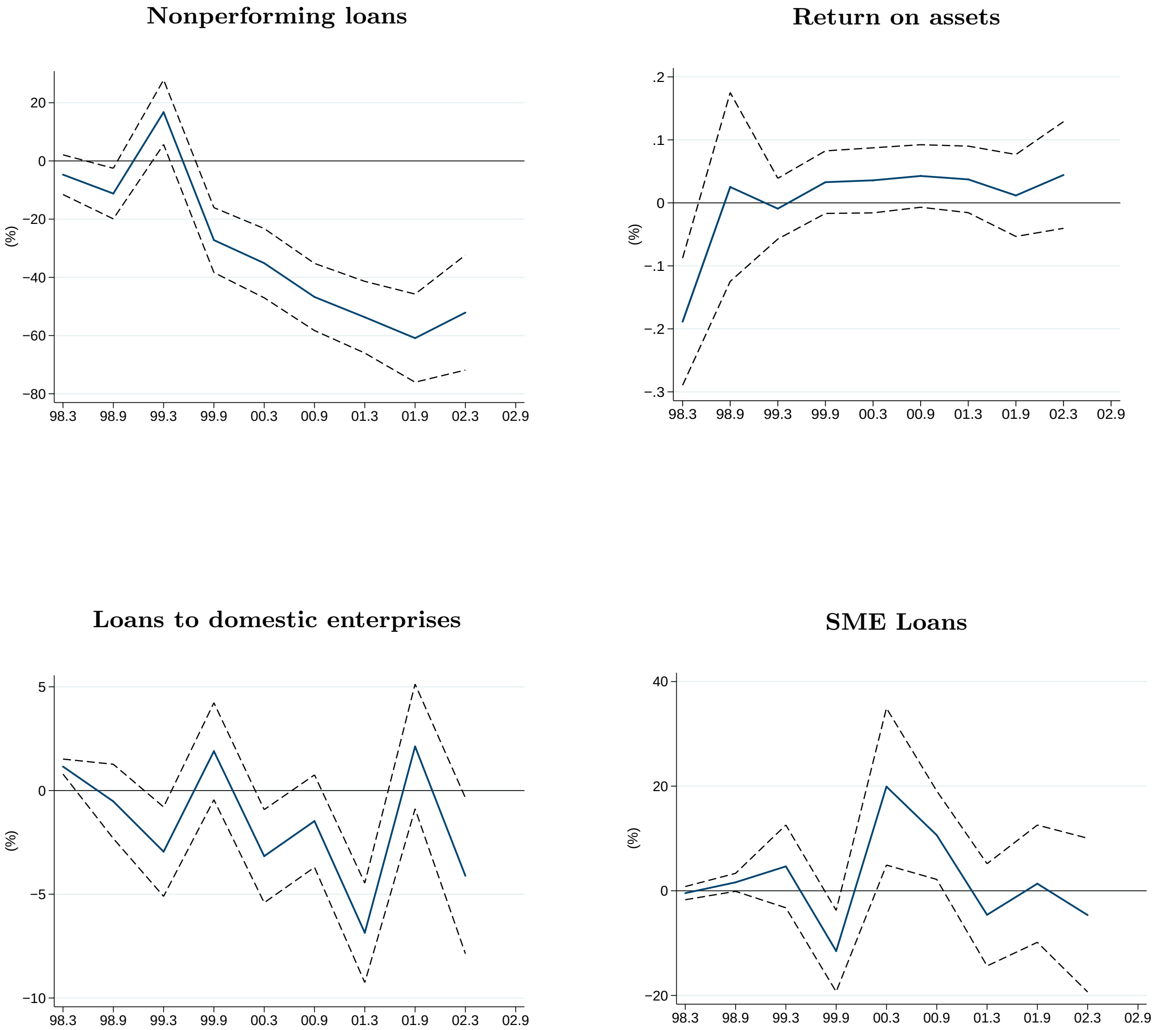
Figure 5: The Treatment Effect on Default Risk Indicators



1. The solid line indicates point estimates, and the dashed line indicates 90% confidence intervals.

2. The confidence intervals are calculated using the method of Conley and Taber (2011). See Appendix II for more details.

Figure 6: The Treatment Effects on Target Variables



1. The solid line indicates point estimates, and the dashed line indicates 90% confidence intervals.

2. The confidence intervals are calculated using the method of Conley and Taber (2011). See Appendix II for more details.

Figure 7: Illustration of the Treatment Effect δ_t