Towson University
Department of Economics
Working Paper Series

Working Paper No. 2010-11

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April, 2010

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April 27, 2010

Abstract

We investigate whether and how the credit crunch during the financial crisis in Japan affected household welfare. We estimate the consumption Euler equation with endogenous credit constraints using household panel data for 1993–1999, generating several findings. First, a small but non-negligible portion of the households faced credit constraints during the crisis, rejecting the standard consumption Euler equation. Second, the credit crunch affected household welfare negatively, albeit not seriously. The estimated welfare loss ranges between two to ten percent increases in marginal utility, depending on income level. Finally, our results corroborate that the credit crunch in Japan was supply-driven.

Keywords: Credit crunch; Consumption Euler equation; Household Welfare

JEL Classification Numbers: D91; E21

1. Introduction

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The authors would like to thank the Institute for Research on Household Economics for providing their panel data and Prof. Takeshi Amemiya for encouraging them to initiate this project. They would also like to thank the editor and an anonymous referee of the journal, Naohito Abe, Tony Cavoli, Dean Corbae, Ippei Fujiwara, Mototsugu Fukushige, Yoshio Higuchi, Charles Yuji Horioka, Takeo Hoshi, Hidehiko Ichimura, Yasushi Iwamoto, Hiroyuki Kasahara, Ryo Kambayashi, Miki Kohara, Takashi Kurosaki, John Litwack, Colin McKenzie, Chiaki Moriguchi, Makoto Saito, Masaya Sakuragawa, Michio Suzuki, Tomoya Suzuki, and Wako Watanabe for their helpful comments and Shunichiro Bessho for his excellent research assistance.

In Japan, the collapse of mega-banks in 1997 increased regulatory pressure, market scrutiny, and the distress of the financial system, thereby causing a crisis in the domestic financial sector (Woo, 2003). The financial crisis in Japan is often referred to as a typical example of a "credit crunch," which is conventionally defined as a sharp decline in bank loans caused by supply factors such as risk-based capital standards imposed on banks.[2] According to the Bank of Japan's Diffusion Indices (DIs) of "banks' willingness to lend," there was a sharp deterioration of this indicator in the first quarter of 1998 (Figure 1). There is a large emerging body of studies that investigates whether and how the credit crunch in Japan constrained firm investments (Caballero, Hoshi, and Kashyap, 2008; Hayashi and Prescott, 2002; Hori, Saito, and Ando, 2006; Hosono, 2006; Ito and Sasaki, 2002; Motonishi and Yoshikawa, 1999; Woo, 2003).[3]

However, there are no clear explanations for the way in which the financial crisis and credit crunch affected the welfare of Japanese households.[4] Casual evidence shows that the negative impact of the credit crunch in Japan was serious at the household level; for example, the growth rate of the credit supply of private banks to individuals shrunk significantly in the 1990s and even became negative in 1992 and 1993 (Ogawa, 2003). In addition, the number of applications for individual bankruptcies increased from 43,545 in 1993 to 122,741 in 1999. Thus, it is natural to hypothesize that this sharp increase in individual bankruptcies was caused by problems in the financial sector. There are many media reports stating that the credit crunch in Japan damaged small firms disproportionately because unlike large listed firms, the only source of their external funding for investments was still bank loans. Accordingly, it has been said that many owners of small firms or businesses went bankrupt after facing a steep decline

[2] However, such a crunch can also be caused by a slump in demand.

[3] For example, Peek and Rosengren (2000) indicate that the credit crunch in Japan became a serious issue even around 1993, by finding that Japanese banks reduced their commercial real estate lending in the US where the financial crisis in Japan was an external event. Ito and Sasaki (2002) corroborate this view. In contrast, Woo (2003) finds empirical support for only the credit crunch experienced in 1997.

[4] Kang and Sawada (2008) examine the manner in which the credit crunch in Korea affected household welfare in 1997 and 1998.

in bank loans during the credit crunch. Although this credit crunch can be assumed to have negatively affected the welfare of small-firm owners and their employees, it is not categorically clear whether the credit crunch really affected small firms disproportionately. For example, Hayashi and Prescott (2002) concluded that small firms could rely on their own cash and deposits to weather the sharp decline in bank loans. Moreover, the negative impact on employment was actually smaller than the original prediction; further, it was not comparable to the impact of other major economic crises such as the Great Depression (Genda, 2003; Hoshi, 2006). The responses of firms, households, and the government appeared to have played important roles in coping with the macroeconomic crisis. Hence, unless a thorough empirical study is undertaken, the overall social impact of the credit crunch is not necessarily clear.

This study aims to bridge the gap in the existing studies by addressing the extent to which households were affected by the crisis. In our empirical analysis, household-level panel data collected from 1993 to 1999 are used. We estimate a consumption Euler equation that is augmented by endogenously imposed credit constraints using a switching regression or Type 5 Tobit model. The methodology permits us to derive a density function of the probabilities of binding credit constraints for each year, which makes it possible to quantify the seriousness of the credit crunch at the household level.

In order to identify whether the credit crunch was driven by supply or demand, we also estimate Kimball's prudence parameter (1990) before and during the credit crunch using the approach suggested by Lee and Sawada (2007); this approach is an extension of the Euler equation approach proposed by Dynan (1993). A change in the prudence parameter characterizes a change in households' saving behavior and thus indicates the demand for credit.

The remaining paper is organized as follows. Section 2 provides a theoretical framework based on which the econometric framework is derived in Section 3. Section 4 presents a discussion of the data and empirical results, and Section 5 concludes the paper.

2. The Model Framework

Under the financial crisis, households faced the problem of reconciling the realized income shortfall with a desirable level of stable consumption. Households have devised several methods, such as self-insurance and mutual insurance, to protect their levels of consumption against the ex-post risks of negative income shocks (Fafchamps, 2003). In this paper, we focus on a particular measure that households use to cope with hardship—credit. Households can utilize credit to smooth consumption by reallocating future resources for current consumption. The lack of consumption insurance can be compensated for by having access to a credit market (Sawada and Shimizutani, 2008). However, households may be constrained from borrowing for a variety of reasons such as asymmetric information between lenders and borrowers. The existence of credit constraints has important negative impacts on the risk-coping ability of poor households. For example, faced with a shortfall in the real income, credit-constrained households may be forced to reduce consumption expenditure since credit cannot be used as insurance. A credit crunch could magnify this negative impact of credit constraints. This channel may explain the mechanisms of the so-called consumption slump and the resultant deflationary spirals in Japan (Krugman, 1998). Interestingly, Olney (1999) also showed that given the high cost of default, reducing consumption was the only viable strategy of American households against the recession in 1930. Accordingly, consumer spending collapsed in 1930, turning a minor recession into the Great Depression. While credit constraints were self-imposed by households in the case of the Great Depression, there appears to be a common element of reinforced credit constraints in recessions in the US and Japan.

2.1 The Model of an Augmented Euler Equation

In order to formalize the role of credit in smoothing consumption, following Zeldes (1989) and Deaton (1991), we construct a model that provides optimal consumer behavior under uncertain income and potential credit constraints. Suppose a household *i*'s decision-maker has a concave instantaneous utility $U(\bullet)$ of the household consumption C_t. The household then has to choose a value of C_{it} with a subjective discount rate δ such that the discounted lifetime utility is maximized, which is subject to intertemporal budget constraints with the interest rate r. Generally, when the household income is stochastic, analytical solutions to this problem cannot be derived. However, in order to obtain an optimum solution, we can derive a set of first-order necessary conditions by forming a value function and a Bellman equation. Let λ represent the Lagrange multiplier associated with credit constraint $A + y - C + z \geq 0$, where A is the household asset at the beginning of the period, y represents the stochastic household income, and the maximum supply of credit possible for this household is represented by z.[5] On combining the envelope condition derived from the first-order conditions, we obtain a consumption Euler equation that is augmented by the credit constraint:

$$
\begin{aligned}
U'(C_{it}) &= E_t\left[U'(C_{it+1})\left(\frac{1+r_{it}}{1+\delta}\right)\right] + \lambda_{it}, \\
A_{it} + y_{it} - C_{it} + z_{it} &\geq 0 \;\; if\; \lambda_{it} = 0, \\
A_{it} + y_{it} - C_{it} + z_{it} &= 0 \;\; if\; \lambda_{it} > 0.
\end{aligned} \tag{1}
$$

2.2 Measuring Welfare Losses from Binding Credit Constraints

This augmented Euler equation (1) was first derived by Zeldes (1989). The term λ is an implicit price of consumption, which is equal to the increase in the expected lifetime utility that would result if the current constraint were relaxed by one unit. Since the household is constrained from borrowing more, but not from saving more, λ assumes a positive sign. Accordingly, we can interpret the Lagrange multiplier λ as an indicator of the welfare losses arising from binding credit constraints (Kang and Sawada, 2008). Note that the Lagrange multiplier λ_{it} is a negative function of the current income y_{it} (Zeldes, 1989). Given other variables, an increase in the current income of a credit-constrained household leads to a decline

[5] When z is sufficiently large, the household can lend and borrow freely at a rate of interest r. A case of complete borrowing constraint, in which a household cannot borrow at all, can be represented by $z = 0$.

in the marginal utility of current consumption, thereby causing the Lagrange multiplier to decline. This theoretical property provides us with a basis for testing the validity of the theoretical framework.

3. Econometric Framework

Our econometric framework aims to test the implications of the augmented Euler equation (1). Here, we employ two different empirical strategies. First, following Zeldes (1989), we suppose that households form their rational expectations and that utility is described by the constant relative risk aversion (CRRA) utility function, that is, $U(C_{it}) = C_{it}^{1-\gamma}(1-\gamma)^{-1}\exp(\theta_{it})$, where θ represents the household size and tastes. Next, equation (1) becomes

$$(2)\quad \hat{C}_{it+1} = \frac{1}{\gamma}\left\{ \log\left[(1+r_{it})\beta\right] + (\theta_{it+1}-\theta_{it}) + \log\left[1+\frac{\lambda_{it}}{E_t\left(C_{it+1}^{-\gamma}\exp(\theta_{it+1})\left(\frac{1+r_{it}}{1+\delta}\right)\right)}\right] - \log(1+e_{it+1}) \right\},$$

where $\hat{C}$ is the consumption growth rate, e denotes the household's expectation error, and $E(e_{it+1}|I_t) = 0$, with I_t being the information set available at time t.

The second approach follows the approach adopted by Lee and Sawada (2007), which is an extension of Dynan (1993), by including endogenous credit constraints. As shown by Lee and Sawada (2007), we can take a second-order Taylor approximation of equation (1) to obtain

$$(3)\qquad \hat{C}_{it+1} = \frac{1}{\gamma}\left(\frac{r_{it}-\delta}{1+r_{it}}\right) + \frac{\rho}{2}\left(\hat{C}_{it+1}\right)^2 - \left(\frac{1+\delta}{1+r_{it}}\right)\frac{\lambda_{it}}{C_{it}U''(C_{it})} + e_{it+1},$$

where $\rho \equiv -C(U'''/U'')$ is the coefficient of relative prudence, as elaborated by Kimball (1990). Note that we follow Merrigan and Normadin (1996) and replace the expected consumption growth with the observed consumption growth as well as the expectation error.

In the case of either approach, the approximated estimable equation becomes

(4) $$\hat{C}_{it+1} = X_{it}\beta + f(\lambda_{it}) + v_{it+1},$$

where X includes the proxy variables for θ, r, and β; $f(\bullet)$ is an increasing function that takes zero if λ becomes 0; and v_{it+1} indicates a well-behaved stochastic error term. In order to control for the changes in preferences and household characteristics, items such as household size, age of the respondent, and age squared were included in X (Zeldes, 1989). Further, in the second approach, X includes the squared consumption growth rate. In consumption growth equation (3), it is natural to regard the squared consumption growth and error terms as being correlated. Thus, we follow Dynan (1993) and treat the squared consumption growth term as an endogenous variable.

Now, let C^* represent the optimal consumption in the absence of a current credit constraint. $C^* = C$ if the credit constraint is not binding, and $C^* > C$ if the credit constraint is binding. Next, we define the gap between optimal consumption under the perfect credit accessibility and cash in hand without credit constraints, that is, $H = C^* - (A + y + z)$. Further, following Hayashi (1985) and Jappelli (1990), we assume that the conditional expectation of optimal consumption C^* can be approximated by a quadratic function. Hence, the reduced form of the optimal consumption C^* can be expressed as a linear function of observables, such as current income, wealth, and age, as well as the quadratic terms of these variables. The maximum amount of borrowing is also assumed to be a linear function of the same variables. Defining a matrix, W, as a set of these variables, we obtain a reduced-form equation.

(5) $$H_{it} = W_{it}\beta_W + \varepsilon_{it},$$

where ε is an error term that captures unobserved elements and measurement error.

From equations (4) and (5), we can derive the following econometric model of the augmented Euler equation with the following endogenous credit constraints:

(6) $$\hat{C}_{it+1} = X_{it}\beta + f(\lambda_{it}) + v_{it+1},$$

$$\lambda_{it}' > 0 \;\; if \;\; H_{it} \geq 0,$$
$$\lambda_{it}' = 0 \;\; if \;\; H_{it} < 0,$$

$$H_{it} = W_{it}\beta_W + \varepsilon_{it}.$$

3.1 Exogenous versus Endogenous Credit Constraints

The conventional empirical approach to estimate equation (6) (Zeldes, 1989) ignores the endogeneity of the Lagrange multiplier and exogenously splits the sample into those households that are likely to be credit-constrained, that is, $\lambda_t > 0$, and those that are not likely to be credit-constrained, that is, $\lambda_t = 0$, by using observable household characteristics. Zeldes (1989) splits the sample based on the wealth-to-income ratio.

The exogenous split approach, however, has two problems (Garcia, Lusardi, and Ng, 1997, p. 158; Hu and Schiantarelli, 1998, p. 466–467). First, it is unlikely that a single variable, such as the wealth-to-income ratio, would serve as a sufficient statistic of a consumer's ability to borrow. Usually, lenders screen credit applicants based on multiple factors. Second, if the variables used as the criteria for splitting a sample were correlated with the unobserved factors in consumption growth, this correlation would generate a sample selection problem. Accordingly, sample selection bias should be controlled for properly.

3.2 Type 5 Tobit Model with Observed Regimes

In order to overcome these two issues, an alternative approach elaborated by Jappelli (1990) is adopted; this approach constructs a qualitative response model of an endogenous credit constraint by defining the indicator variable of a credit constraint, which would take the value of one if the credit constraints are binding, and zero otherwise. Jappelli, Pischeke, and Souleles (1998) combined this model of endogenous credit constraint with a consumption Euler equation. Accordingly, in order to estimate a system of equations (6), we can combine the endogenous credit constraint approach of Jappelli (1990) with the augmented Euler equation.

Let the function of the Lagrange multiplier $f(\lambda)$ be a piecewise linear function of a set of variables Z with a coefficient vector ψ that is specific to the credit constraint status. With subscripts N and C representing the credit non-constrained and constrained groups, respectively, the estimable augmented Euler equation (6) can be rewritten as follows:

$$\hat{C}_{it+1} = X_{it}\beta_X + (1-d_{it})Z_{it}\psi_N + d_{it}Z_{it}\psi_C + v_{it+1}, \tag{7}$$

$$d_{it} = 1[\, W_{it}\beta_W + \varepsilon_{it} \geq 0], \tag{8}$$

where 1[•] is an indicator function. The testable restriction is that the elements of the coefficient vector for the non-constrained group in equation (7), that is, ψ_N, are all zero. We assume that errors follow a joint normal distribution:

$$\begin{pmatrix} v_{Nit+1} \\ \varepsilon_{it} \end{pmatrix} \sim N\left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma^2 & \sigma_{v\varepsilon} \\ \sigma_{v\varepsilon}^2 & 1 \end{pmatrix}\right]. \tag{9}$$

If the sign of H is observable, the model can be estimated by the Type 5 Tobit model with observed regime (Amemiya 1985, p. 399–408). The Type 5 Tobit model explicitly considers the endogenous sample selection bias arising from the Ordinary Least Squares (OLS) estimation of equation (7). We can consistently estimate the parameters in the Euler and credit-constrained equations by maximizing the following log-likelihood function:

$$\begin{aligned} &l_t(\beta_X, \psi_N, \psi_C, \beta_W, \sigma\ , \sigma_{v\varepsilon}) \\ (10)\ &= \sum_{i=1}^{n}\ (1-d_{it})\ln\{f(v_{it+1} \mid H_{it} < 0)\mathrm{Prob}(H_{it} < 0)\} + d_{it}\ln\{f(v_{it+1} \mid H_{it} \geq 0)\,\mathrm{Pr\,ob}(H_{it} \geq 0)\} \end{aligned}$$

$$= \sum_{i=1}^{n} (1-d_{it}) \ln \left\{ \left[\frac{1}{\sigma} \varphi\left(\frac{v_{it+1}}{\sigma} \right) \right] \Phi\left(\frac{-W_{it}\beta_W - \frac{\sigma_{v\varepsilon}}{\sigma^2} v_{it+1}}{\sqrt{1-\frac{\sigma_{v\varepsilon}^2}{\sigma^2}}} \right) \right\}$$

$$+ d_{it} \ln \left\{ \left[\frac{1}{\sigma} \varphi\left(\frac{v_{it+1}}{\sigma} \right) \right] \left[1-\Phi\left(\frac{-W_{it}\beta_W - \frac{\sigma_{v\varepsilon}}{\sigma^2} v_{it+1}}{\sqrt{1-\frac{\sigma_{v\varepsilon}^2}{\sigma^2}}} \right) \right] \right\},$$

where $\phi(\bullet)$ and $\Phi(\bullet)$ represent the density and cumulative distribution functions of standard normal distribution, respectively. We take the OLS estimation results as the starting values and set the auxiliary parameters (σ, $\sigma_{v\varepsilon}$) to be (1, 0) initially. The estimated parameters are then employed as the updated initial values to re-estimate all the parameters.[6]

3.3 Type 5 Tobit Model with Unobserved Regimes I

However, a precise measurement of the credit constraint is not straightforward. A direct approach involves utilizing the information on a household's willingness and ability to obtain credit (Jappelli, 1990; Jappelli, Pischeke, and Souleles, 1998). Generally, household-level data on credit availability is not available in standard household surveys (Scott, 2000) because the credit constraint status cannot be identified by only considering the amount of attained credit. Even in cases where the indicator variable for the credit constraint is not observed, we can apply the estimation method of a switching model with unknown regimes. Following a recent study by Garcia, Lusardi, and Ng (1997), we estimate the Euler equation augmented by endogenous credit constraints as a switching regression model. Although we cannot observe *H* directly, we can estimate the probability of being credit-constrained jointly with other parameters in Euler equations by maximizing the following log-likelihood function:

[6] Alternatively, we can estimate the Type 5 Tobit model of equations (7) and (8) by using Heckman and Lee's two-step procedure. Nevertheless, Nawata (1994) showed that the maximum likelihood estimator is superior to the two-step estimator.

(11)

$$L_t(\beta_X, \psi_N, \psi_C, \beta_W, \sigma \ , \sigma_{v\varepsilon})$$

$$= \sum_{i=1}^{n} \log\left\{\left[\frac{1}{\sigma}\varphi\left(\frac{v_{it+1}}{\sigma}\right)\right]\Phi\left(\frac{-W_{it}\beta_W - \frac{\sigma_{v\varepsilon}}{\sigma^2}v_{it+1}}{\sqrt{1-\frac{\sigma_{v\varepsilon}^2}{\sigma^2}}}\right) + \left[\frac{1}{\sigma}\varphi\left(\frac{v_{it+1}}{\sigma}\right)\right]\left[1-\Phi\left(\frac{-W_{it}\beta_W - \frac{\sigma_{v\varepsilon}}{\sigma^2}v_{it+1}}{\sqrt{1-\frac{\sigma_{v\varepsilon}^2}{\sigma^2}}}\right)\right]\right\}.$$

While the nonlinearity of the system made convergence difficult, we achieved interior solutions from the OLS starting values and the random attempts that were made to ascertain better feasible initial values. We again set auxiliary parameters (σ, $\sigma_{v\varepsilon}$) as (1, 0) initially. In order to re-estimate all the parameters, we employ the estimated parameters as the initial values.

According to the testable restriction derived from the theoretical result of augmented Euler equation (1), for the non-constrained group, the elements of the coefficient vector, that is, ψ_N, are all zero, while for the constrained group, the elements of the coefficient vector, that is, ψ_C, are non-zero.

3.4 Type 5 Tobit Model with Unobserved Regimes II

We also estimate the augmented Euler equation with unobserved regimes by letting the parameter vector β_X differ depending on the regime. In this case, we have the following econometric model, which was employed by Garcia et al. (1997) and Kang and Sawada (2008):

(12) $$\hat{C}_{it+1} = X_{it}\beta_N + Z_{it}\psi_N + v_{Nit+1} \;\; if \;\; H_{it} < 0,$$

(13) $$\hat{C}_{it+1} = X_{it}\beta_C + Z_{it}\psi_C + v_{Cit+1} \;\; if \;\; H_{it} \geq 0,$$

(14) $$H_{it} = W_{it}\beta_W + \varepsilon_{it}.$$

As earlier, the testable restriction of our framework is that the elements of the coefficient vector in equation (12), that is, ψ_N, are all zero for the non-constrained group. We assume that errors follow a joint normal distribution with zero means and the following covariance matrix:

(15)
$$\begin{pmatrix} v_{Nit+1} \\ v_{Cit+1} \\ \varepsilon_{it} \end{pmatrix} \sim N\left[\begin{pmatrix} 0 \\ 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_N^2 & \sigma_{NC} & \sigma_{N\varepsilon} \\ \sigma_{CN} & \sigma_C^2 & \sigma_{C\varepsilon} \\ \sigma_{\varepsilon N} & \sigma_{\varepsilon C} & \sigma_\varepsilon^2 \end{pmatrix}\right].$$

For identification, we assume that $\sigma_\varepsilon^2 = 1$. In addition, note that σ_{NC} and σ_{CN} are not identifiable because only one of the two regimes can be observed. All other parameters can be identified.

Although we cannot observe H directly, we can estimate the probability of being credit-constrained jointly with other parameters in Euler equations by maximizing the following log-likelihood function:

(16)

$$L_t(\beta_N, \beta_C, \psi_N, \psi_C, \beta_W, \sigma_N, \sigma_C, \sigma_{N\varepsilon}, \sigma_{C\varepsilon})$$
$$= \sum_{i=1}^{n} \log\left\{\left[\frac{1}{\sigma_N}\varphi\left(\frac{v_{Nit+1}}{\sigma_N}\right)\right]\Phi\left(\frac{-W_{it}\beta_W - \frac{\sigma_{N\varepsilon}}{\sigma_N^2}v_{Nit+1}}{\sqrt{1-\frac{\sigma_{N\varepsilon}^2}{\sigma_N^2}}}\right) + \left[\frac{1}{\sigma_C}\varphi\left(\frac{v_{Cit+1}}{\sigma_C}\right)\right]\left[1-\Phi\left(\frac{-W_{it}\beta_W - \frac{\sigma_{C\varepsilon}}{\sigma_C^2}v_{Cit+1}}{\sqrt{1-\frac{\sigma_{C\varepsilon}^2}{\sigma_C^2}}}\right)\right]\right\}.$$

4. Data and Estimation Results

This paper uses the Japanese Panel Survey of Consumers data set from 1993 to 1999, which was collected by the Institute for Research on Household Economics. The survey was conducted in all Japanese prefectures. The data comprised the responses to multipurpose surveys with household and individual modules. The initial survey in 1993 was completed by 1,500 women between the ages of 24 and 35.[7] After excluding the observations with missing information, we construct a balanced panel data set that comprises the survey responses of 807 women. The definitions and summary statistics of the variables are provided in Table 1.

[7] The respondents are only women by construction of the survey.

4.1 Descriptive Statistics

Table 1 summarizes the descriptive statistics of the variables used in this paper. Household income is calculated as the total family income before tax minus asset income such as returns from securities. The income is then deflated by the consumer price index. The surveys asked about detailed income components in the calendar year immediately before the interview. By following Zeldes (1989, p. 326), we utilize this initial income variable as the only element of Z when we implement econometric analysis. This also mitigates the endogeneity bias arising from the correlation between the initial income and the error term. According to Table 1, real household income constantly increased from 4.8 million yen in 1993 to 6 million yen in 1998.[8]

With respect to total expenditure, questions were asked about the family expenditure in the month of September each year. We multiplied this expenditure figure by 12 and deflated it by the consumer price index.[9] Real annual household expenditure also increased from 1.74 million yen in 1993 to 2.17 million yen in 1999.

The household asset variable is obtained as the total amount of deposits, bonds, and securities owned by the family. It increased from 2.84 million yen in 1993 to 4.02 million yen in 1998. The amount of outstanding debt is equal to the sum of housing and other loans. It increased sharply from 3.11 million yen in 1993 to 6.99 million yen in 1998. This may be due to life cycle effects—the respondents were at an early stage of their lives, and thus, they were likely to accumulate debts to finance housing and other investments. In fact, the housing ownership rate is very high. Initially, the rate was 60% in 1993, which increased to 67% in

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[8] However, it stayed almost unchanged in 1999.

[9] We use a different data source, the Family Income and Expenditure Survey (FIES), to calculate the ratio of September consumption to annual consumption. In 1993–1999, the September expenditure levels are less than the average and the ratio fluctuates over the years. This means that the consumption variable that we use may not be similar to the annual consumption. In order to mitigate this potential data problem, we conduct a robustness test with data adjustments by dividing our individual consumption variables by the ratios of September consumption per household to annual per household consumption from the FIES. The qualitative results are similar to the results reported in the paper.

1998. Note, however, that this increase in debt does not necessarily imply that credit constraints are less binding, since the demand for credit might have increased significantly at the same time.

In order to characterize the respondents further, we included marital status and education variables. While approximately 27% of the surveyed women were single in 1993, 17% remained single in 1998. With respect to the education level in the sample, 43% of the respondents were senior high school graduates; 19%, vocational school graduates; 20%, junior college graduates; and 12%, college or university graduates.

Finally, city size was categorized into three groups, that is, major cities with more than 0.5 million people, other middle- or small-sized cities, and other areas such as towns and villages. It is notable that approximately 25% and 60% of the respondents were living in major and middle-sized cities, respectively.

4.2 Variables Used in the Analysis

Throughout the estimation of consumption Euler equations, we consider the annual growth rate of total expenditures as a dependent variable.[10] Independent variables include the respondent's age and age squared, co-residence dummy, the first difference of the household size, and initial household income with a slope dummy variable for credit-constrained and non-constrained households. The co-residence dummy variable is exclusively for single women; this takes the value of one if a single woman lives with others. In the final specification based on equation (3), we follow Dynan (1993) and Lee and Sawada (2007) and include the squared consumption growth rate, treating this variable as endogenous.

With respect to the credit constraint equation, the explanatory variables are as follows: household income, household income squared, asset, asset squared, debt, debt squared, age, age squared, marital status, household size, and the education and indicator variables for city

[10] This variable is computed by taking the first difference of log real household expenditure.

size. We also include the co-residence dummy for single women. Following the probit analysis of the credit constraint equation of Jappelli (1990), the dependent variable should be a credit constraint dummy that takes the value of one if the credit constraint is binding, and zero otherwise. Notably, the 1993 and 1998 surveys include a special section that precisely measures credit constraint by directly asking about a household's willingness and ability to obtain credit. In our sample, the percentages of the households facing credit constraint are 7.8% in 1993 and 8.4% in 1998 (Table 1).

4.3 Estimation Results I: The Case of Observed Regimes

By using the 1993 and 1998 surveys in which the credit constraints can be identified directly from the data set, Table 2 shows the estimation results of augmented Euler equation (7) under observable but endogenous credit constraints (8), achieved by maximizing the likelihood function in equation (10). In the Euler equation section, we find that the income coefficient for credit-constrained households is negative and statistically significant at 10%, while the coefficient for non-constrained households is not statistically different from zero. The result confirms our hypothesis that the source of violation of the standard LC-PIH is the binding credit constraint.

These results are consistent with the theory of consumption smoothing under endogenous credit constraints. With regard to the credit constraint equation, as observed in Table 2, the probability of binding credit constraints is a positive function of debt; however, the amount of assets decreases the probability. These results imply that loan provisions are positively affected by the net amount of assets that can be used as collateral. Further, we also found that the respondents with higher education tended to have a lower probability of being credit-constrained; this confirmed the finding of Horioka and Kohara (1999, 2006).

4.4 Estimation Results II: The Case of Unobserved Regimes

We present the second set of estimation results using the full sample from 1993 to 1999. This part comprises two subparts. First, we estimate an econometric model of augmented Euler equation (7) under unobservable credit-constraint indicators by maximizing the likelihood function in equation (11). Second, we estimate the augmented Euler equation by relaxing the assumption of the same coefficients across regimes. In this approach, we estimate equations (12), (13), and (14) jointly by maximizing the likelihood function in equation (16) under the distributional assumption in equation (15).[11]

In Table 3, we represent the estimation results of the augmented Euler equation under unobservable but endogenous credit constraints, achieved by maximizing the likelihood function in equation (11). In other words, we estimate the model of equations (12), (13), and (14) under constraints $\beta_N = \beta_C$ and $v_N = v_C = v$. In the Euler equation section, we find that the income coefficients for credit-constrained and non-constrained households are negative and positive, respectively. The direction and statistical significance of the former coefficient is consistent with the theory of consumption smoothing under endogenous credit constraints. The estimation results of the credit constraint equation are summarized in the lower section of Table 3. While none of the individual coefficients are statistically significant, they are jointly significant at the 10% level.

Table 4 summarizes the estimation results of equations (12), (13), and (14) jointly by maximizing the likelihood function in equation (16). The qualitative findings in Table 4 are similar to those in Table 3 even after relaxing the same coefficients on Euler equations, while the *t*-value of the income coefficient for the constrained group is a mere 1.34. Thus, the results only marginally support the theory. Interestingly, in the credit constraint equation, the coefficient on the year dummy variable for 1998 takes a large positive value and is statistically

[11] We also follow Zeldes' exogenous grouping approach (1989) and split the households based on the wealth-to-income ratio. A household is regarded as being credit-constrained if the calculated total non-housing wealth is less than two months' worth of the average income. The coefficients on the initial income for the constrained and non-constrained groups are negative as well as marginally significant. The level of statistical significance is slightly larger for the credit-constrained households.

significant. This implies that the seriousness of the credit crunch at the household level had increased in 1998 regardless of household characteristics. Moreover, the estimated covariance of the error terms in (13) and (14) is positive and statistically significant, suggesting that the endogeneity bias generated by credit constraints is not necessarily negligible.

4.5 Probability of Binding Credit Constraints and Estimated Welfare Losses

In order to further compare credit accessibility across the years, we computed the probability of binding credit constraints $\Phi\left(W_{it}\hat{\beta}_W\right)$ by using the estimated parameter vectors shown in Tables 2 and 4. Note that the joint Wald test results, where all the coefficients in the credit constraint equation are jointly different from zero, support the validity of these estimated coefficients.[12]

Figures 2a and 3a represent the kernel density functions of the predicted credit-constrained probabilities using the switching regression results for 1993–1994 & 1998–1999 and 1993–1999, respectively. We employed a Gaussian kernel to estimate the density functions. The band width of the density function is selected such that the mean integrated square error is minimized. In both these figures, the distribution of credit-constrained households in 1993 and 1998 is skewed toward the left, indicating that only a small portion of households face credit constraints. In addition, as Figure 3a reveals, the density function moves slightly toward the right over time, indicating that the probability of binding credit constraints has increased. The two-sample Kolmogorov-Smirnov tests of the equality of distributions reject the equality between the pairs of the three probability distributions of binding credit constraints in 1993, 1995, and 1998 at the 5% level of statistical significance. These findings may imply that the credit crunch generated negative effects on the households' consumption smoothing behavior especially at the end of the 1990s.

[12] The Wald test statistics are 78.39 with a p-value of 0.000 and 124.37 with a p-value of 0.000 for the results in Tables 4 and 6, respectively.

However, in Figure 2a, the median (average) probabilities of binding credit constraint are 4.82% (7.82%) in 1993 and 5.77% (8.44%) in 1998. The corresponding probabilities in Table 3a are merely 8.88% (13.80%), 10.69% (14.88%), and 10.79% (13.83%), respectively, in 1993, 1995, and 1998. Hence, the negative welfare effects of credit crunch on household welfare may be negligible. Figure 2b and 3b further compare the density functions by using the cumulative density functions of the predicted probabilities of binding credit constraints. By looking at these figures, we can perform an eye-ball test of the stochastic dominance; we ascertain that the second-order stochastic dominance holds. Initially, in 1993, for the relatively less credit-constrained group, the probability of binding credit constraints appears to have increased significantly. However, it is not necessarily obvious that the predicted probabilities in 1993 are dominated by the probabilities in 1995, which, in turn, are dominated by the probabilities in 1998. In fact, based on the Kolmogorov-Smirnov tests for Figures 3a and 3b, we cannot reject the same distribution if we take the 10% and 1% levels of significance, respectively; this supports our result that credit crunch generated a rather minor effect on household welfare.

With respect to the magnitude of the welfare losses arising from binding credit constraints, we compute the expected value of the ratio of the Lagrange multiplier associated with credit constraints, λ, to the expected marginal utility of consumption in the next period, that is, $E\{\lambda_{it}/E_t[U'(C_{it}+1)(1+r_{it})\ \beta\ ]\}$. This ratio shows the degree of welfare losses in the increase in implicit consumption prices. We apply the method of welfare computation developed by Kang and Sawada (2008) to compute the expected value of this ratio by the formula $\Phi\left(W_{it}\hat{\beta}_W\right)\bullet\left[\exp\gamma(\psi_0+\hat{\psi}_C \ln y_{it})-1\right]$ based on the estimated parameters of equations (13) and (14) reported in Table 4. We assume that ψ_0 is equals to the absolute value of the minimum value of $\psi_C \ln y$ so that the estimated λ becomes non-negative. Table 5 shows the calculated welfare losses due to binding credit constraints for the year 1998 by income quartile normalized by the expected marginal utility for the year 1999. In our computation, we postulate that the coefficient of relative risk aversion takes the value of two, that is, $\gamma = 2$. As can be seen in Table 5, our results indicate that the expected welfare loss from binding credit constraints is higher for poorer households, suggesting the gravity of the credit crunch on poverty. However, the estimated price increase is merely 10% even for the poorest quartile. Hence, we may conclude that the credit crunch affected household welfare negatively after 1997, albeit not seriously.

4.6 Estimation Results III: *Household Prudence*

While these overall empirical results are in accordance with our theoretical framework, the level of statistical significance in Table 6 is not necessarily supportive of our theoretical framework. In order to improve the accuracy of our estimation, we employ an alternative estimation strategy based on equation (3). By doing so, we can also estimate the prudence parameter that summarizes the degree of precautionary saving.

We employ Type 5 Tobit model with unobserved regimes as before by including a new independent variable—squared consumption growth rate. However, an additional issue that we consider here is an endogeneity bias arising from a correlation between the error term of the Euler equation and the squared consumption growth rate on the right-hand side of equation (3) (Dynan, 1993). Specifically, we postulate a linear regression equation for the squared consumption growth rate. Assuming that the error terms of the squared consumption growth rate and the consumption growth rate equation follow a bivariate normal distribution, we can employ the two-stage method suggested by Smith and Blundell (1986) to control for the endogeneity bias (Lee and Sawada, 2007, 2010).

In the first step, we regress the squared consumption growth rate on a set of instrumental variables, including marriage status, co-residence status, education variables, number of household members, household ownership information, asset variables, and debt variables. We reject the hypothesis that the coefficients of the instruments are jointly equal to zero in the first stage at the 1% level. These figures are comparable to those in previous studies, including Dynan (1993). In the second stage, we follow Smith and Blundell (1986) and include the residual from the first-stage regression to control for endogeneity bias. We then estimate augmented Euler equation (7) under unobservable credit constraint indicators given in equation (8) by maximizing the likelihood function in equation (11).

Table 8 shows the estimation results. Specification (1) represents the replication specification shown by Dynan (1993). It confirms the small prudence puzzle because the coefficient of squared consumption growth rate is not statistically different from zero. The prudence estimates continue to be indistinguishable from zero (Specifications (2) and (3)) even when income is included. We also use the system of equations (7) and (8) by allowing different coefficients for the Euler equation parameters except the constant term.[13] If we split the sample endogenously, we find that income coefficients are negative and statistically significant for the credit-constrained households but insignificant for the non-constrained households (Specification (4)). In addition, the constrained households have stronger precautionary saving motives. They behave more prudently than the non-constrained ones. Because the liquidity constrained households have accumulated less net wealth than the unconstrained households, the observed difference in prudence estimates could be supporting evidence of the Decreasing Relative Risk Aversion (DRRA) or the Decreasing Absolute Risk Aversion (DARA), which is theoretically studied by Kimball (1990). The degree of prudence for the constrained households ($0.15 \times 2 = 0.3$) is larger than that in Dynan (ranging from 0.14 to 0.166). However, our estimates may still be smaller than the expected size of the prudence that ranges from 2 to 5 (Dynan, 1993; Lee and Sawada, 2007). The remaining gap may be

[13] We also tried to estimate the augmented Euler equation by jointly relaxing the assumption of the same coefficients across regimes, that is, across equations (12), (13), and (14), by maximizing the likelihood function in equation (16) under the distributional assumption in equation (15). However, we failed to achieve convergence in this model.

related to the approximation bias, as argued by Ludvigson and Paxson (2001), and the concavity of the consumption function, as discussed by Carroll (2001) and Carroll and Kimball (2007).

We include both unmarried and married respondents in our sample, but there is a serious problem with the data on unmarried individuals living with their parents or others. Data are available only on the household consumption financed by the respondent and not on the household consumption financed by other household members. In order to check a potential bias arising from this defect, we eliminate co-resident respondents. The results are represented as specification (6) in Table 6. As we can see, qualitative results are robust against the inclusion of co-resident respondents.

4.7 Supply- or Demand-Driven Crunch?

In the estimation results of the credit constraint equation of Table 6, we found that there exists a uniformly larger probability of binding credit constraints for the year 1998 than that for other years. This suggests that the credit crunch was supply-driven. In order to further identify whether the credit crunch was driven by supply or demand, we also estimate the prudence parameters for different phases. Specifically, we allow the prudence parameter to be different before and during the period from 1996 to 1999.[14] The result is reported in Specification (5) in Table 8. For the credit-constrained group, the prudence parameter is slightly smaller for the latter period. Our estimates may be regarded as being consistent with those of Hori and Shimizutani (2006), who found that the prudence parameter had dropped in the period from 1997 to 1998.

If the prudence parameter increases, precautionary saving should increase and the optimal consumption level should decline. In such a case, precautionary saving is likely to act

[14] We attempted to differentiate the period from 1997 to 1999, but could not achieve convergence of the likelihood function.

as a self-imposed credit constraint, as discussed in Deaton (1992). Thus, such a scenario is expected to present a demand-driven credit crunch. In contrast, our result shows that precautionary saving declines and optimal consumption level increases. This result suggests that the credit crunch in 1997 is inconsistent with the demand-driven credit crunch hypothesis. Hence, it is more likely to be supply-driven than demand-driven.

5. Conclusions

In this study, we investigated the manner in which households in Japan were affected by the credit crunch. Several important empirical findings emerged. First, we found that a small but a non-zero portion of people face credit constraints. Accordingly, our results reject the standard consumption Euler equation, that is, the necessary condition of the life cycle permanent income hypothesis. The maximum likelihood estimation results support our framework of the Euler equation with endogenous credit constraints. Finally, the analyses of the full data for the period from 1993 to 1999 indicate that the credit crunch became particularly more prevalent at the household level after 1997. However, the overall negative impact of credit crunch may not be severe even for the poorest income group.

When market or non-market opportunities for risk sharing are limited, credit serves as an insurance substitute. In the event of negative transitory shocks to their income, households can obtain credit instead of receiving an insurance payment; this helps the households to smooth out such shocks. To a small group of marginally non-constrained households, the financial crunch in Japan appears to have disabled the role of credit as an important self-insurance device. However, our results suggest that the overall impact of the credit crunch on households was not significant. Our study suggests that unlike the Great Depression in the US in 1930 wherein consumer spending collapsed, thereby turning a minor recession into a serious depression (Olney, 1999), the credit crunch in Japan was unlikely to have magnified the negative macroeconomic shocks in 1997.

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Table 1. Descriptive Statistics

Variable name	Mean (Std. Dev.)	Mean (Std. Dev.)	Mean (Std. Dev.)	Mean (Std. Dev.)	Mean (Std. Dev.)	Mean (Std. Dev.)	Mean (Std. Dev.)
Year	1993	1994	1995	1996	1997	1998	1999
Number of household members	4.08	4.14	4.17	4.17	4.22	4.25	
	(1.45)	(1.47)	(1.48)	(1.54)	(1.53)	(1.55)	
Age of the respondent	29.03	30.03	31.03	32.02	33.02	34.03	
	(3.18)	(3.18)	(3.18)	(3.18)	(3.18)	(3.18)	
Amount of outstanding debt (1993 price, in 10,000 Yen)	310.53	435.35	532.35	581.11	672.01	698.96	
	(759.91)	(897.08)	(1020.73)	(1060.37)	(1119.68)	(1126.10)	
Annual household expenditure (1993 price, in 10,000 Yen)	173.69	181.68	189.97	193.76	195.50	213.42	216.61
	(89.24)	(92.79)	(96.26)	(91.16)	(91.59)	(110.43)	(114.18)
Expenditure growth rate between t and t + 1	0.05	0.04	0.03	0.03	0.08	0.00	
	(0.50)	(0.53)	(0.57)	(0.53)	(0.48)	(0.52)	
Annual household income (1993 price, in 10,000 Yen)	476.74	488.29	537.31	561.62	589.17	596.51	
	(247.06)	(254.74)	(257.44)	(255.73)	(284.65)	(320.50)	
Household assets (total amount of deposits plus total amount of securities) (1993 price, in 10,000 yen)	284.09 (394.26)	320.19 (425.73)	359.32 (456.28)	393.55 (593.97)	384.72 (541.51)	402.28 (579.75)	
Credit constraint dummy for 1993 based on Jappelli (1990) (= 1, if constrained; = 0, otherwise)	0.078					0.084	
Household ownership dummy (= 1, if own house; = 0, otherwise)	0.60	0.62	0.63	0.63	0.66	0.67	
Single dummy (= 1, if single; = 0, otherwise)	0.27	0.23	0.21	0.19	0.18	0.17	
Co-residence dummy for single women (= 1, if living with other(s); = 0, otherwise)	0.25	0.22	0.20	0.16	0.15	0.14	
Education level dummy (= 1, if senior high school graduate; = 0, otherwise)	0.43	0.43	0.43	0.43	0.43	0.43	
Education level dummy (= 1, if vocational school graduate; = 0, otherwise)	0.19	0.19	0.19	0.19	0.19	0.19	
Education level dummy (= 1, if junior college graduate; = 0, otherwise)	0.20	0.20	0.20	0.20	0.20	0.20	
Education level dummy (= 1, if university graduate; = 0, otherwise)	0.12	0.12	0.12	0.12	0.12	0.12	
Major city dummy (= 1, if major city; = 0, otherwise)	0.24	0.23	0.23	0.24	0.24	0.24	
Other cities dummy (= 1, if other cities; = 0, otherwise)	0.59	0.59	0.59	0.58	0.58	0.58	
Sample size	807	807	807	807	807	807	807

Table 2. Estimation Results of the Augmented Euler Equation Dependent Variable: Growth Rate of Expenditure for 1993/1994 and 1998/99

	Coef.	Std. Err.
Euler equation		
Single dummy (= 1, if single; = 0, otherwise)	0.100	(0.053)*
Co-residence dummy for single women (= 1, if living with other(s); = 0, otherwise)	–0.108	(0.074)
Age of the respondent	–0.098	(0.046)**
Age squared	–.001	(0.001)**
First difference of the household size	–0.025	(0.022)
Log income for credit-constrained households	–0.026	(0.015)*
Log income for non-constrained households	–0.006	(0.007)
Constant	–.660	(0.740)**
Credit constraint equation		
Single dummy (= 1, if single; = 0, otherwise)	0.080	(0.318)
Co-residence dummy for single women (= 1, if living with other(s); = 0, otherwise)	–0.322	(0.328)
Number of household members	–0.034	(0.038)
Annual household income (1993 price, in log)	0.000	(0.031)
Log income squared	–0.014	(0.010)
Household assets (total amount of deposits plus total amount of securities) (1993 price, in log)	0.065	(0.073)
Log asset squared	–0.032	(0.012)***
Amount of outstanding debt (1993 price, in log)	0.138	(0.063)**
Amount of outstanding debt squared (1993 price, log value squared)	–0.007	(0.009)
Household ownership dummy (= 1, if own house; = 0, otherwise)	–0.030	(0.128)
Age of the respondent	–0.026	(0.171)
Age squared	0.001	(0.003)
Education level dummy (= 1, if senior high school graduate; = 0, otherwise)	–0.403	(0.167)**
Education level dummy (= 1, if vocational school graduate; = 0, otherwise)	–0.456	(0.196)**
Education level dummy (= 1, if junior college graduate; = 0, otherwise)	–0.640	(0.203)***
Education level dummy (= 1, if university graduate; = 0, otherwise)	–0.666	(0.246)***
Year dummy for 1998 (= 1, if 1998; = 0, otherwise)	–0.020	(0.122)
Major city dummy (= 1, if major city; = 0, otherwise)	0.288	(0.170)*
Other cities dummy (= 1, if other cities; = 0, otherwise)	0.171	(0.151)
Constant	–0.072	(2.790)
Other parameters		
σ_v^2	0.508	(0.038)***
$\sigma_{v\varepsilon}$	0.542	(0.049)
Sample size	1614	

Note: Standard errors are indicated in parentheses. *** indicates significance at the 1% level; **, at the 5% level; and *, at the 10% level.

Table 3. Estimation Results of the Augmented Euler Equation Dependent Variable: Annual Growth Rate of Expenditure between 1993 and 1999

	Coef.	Std. Err.
Euler equation		
Single dummy (= 1, if single; = 0, otherwise)	0.049	(0.053)
Co-residence dummy for single women (= 1, if living with other(s); = 0, otherwise)	0.060	(0.079)
Age of the respondent	–0.052	(0.060)
Age squared	0.001	(0.001)
First difference of the household size	–0.015	(0.012)
Log income for credit-constrained households	–0.067	(0.012)***
Log income for non-constrained households	0.034	(0.012)***
Constant	0.954	(0.967)
Credit constraint equation		
Single dummy (= 1, if single; = 0, otherwise)	0.098	(0.131)
Co-residence dummy for single women (= 1, if living with other(s); = 0, otherwise)	0.185	(0.206)
Number of household members	–0.002	(0.015)
Annual household income (1993 price, in log)	0.278	(0.368)
Log income squared	–0.022	(0.034)
Household assets (total amount of deposits plus total amount of securities) (1993 price, in log)	–0.019	(0.046)
Log asset squared	0.002	(0.006)
Amount of outstanding debt (1993 price, in log)	–0.020	(0.026)
Amount of outstanding debt squared (1993 price, log value squared)	0.002	(0.004)
Household ownership dummy (= 1, if own house; = 0, otherwise)	0.018	(0.039)
Age of the respondent	–0.036	(0.166)
Age squared	0.001	(0.003)
Education level dummy (= 1, if senior high school graduate; = 0, otherwise)	–0.016	(0.096)
Education level dummy (= 1, if vocational school graduate; = 0, otherwise)	–0.056	(0.104)
Education level dummy (= 1, if junior college graduate; = 0, otherwise)	–0.029	(0.086)
Education level dummy (= 1, if university graduate; = 0, otherwise)	–0.086	(0.104)
Year dummy for 1994 (= 1, if 1994; = 0, otherwise)	–0.045	(0.058)
Year dummy for 1995 (= 1, if 1995; = 0, otherwise)	0.034	(0.054)
Year dummy for 1996 (= 1, if 1996; = 0, otherwise)	0.047	(0.069)
Year dummy for 1997 (= 1, if 1997; = 0, otherwise)	–0.099	(0.063)
Year dummy for 1998 (= 1, if 1998; = 0, otherwise)	–0.046	(0.073)
Major city dummy (= 1, if major city; = 0 otherwise)	–0.046	(0.053)
Other cities dummy (= 1, if other cities; = 0, otherwise)	–0.031	(0.051)
Constant	–0.255	(2.660)
Other parameters		
σ_v^2	0.706	(0.028)***
$\sigma_{v\varepsilon}$	0.701	(0.028)*:*
Sample size	4842	

Note: Standard errors are indicated in parentheses. *** indicates significance at the 1% level; **, at the 5% level; and *, at the 10% level.

Table 4. Estimation Results of the Augmented Euler Equation Dependent Variable: Annual Growth Rate of Expenditure between 1993 and 1999

	Non-Constrained.		Constrained	
Euler equation	Coef.	Std. Err.	Coef.	Std. Err.
Single dummy (= 1, if single; = 0, otherwise)	0.019	(0.031)	0.162	(0.546)
Co-residence dummy for single women (= 1, if living with other(s); = 0, otherwise)	–0.005	(0.039)	0.511	(0.826)
Age of the respondent	–0.021	(0.031)	–0.267	(0.235)
Age squared	0.0003	(0.0004)	0.004	(0.004)
First difference of the household size	–0.009	(0.015)	–0.014	(0.091)
Log income	–0.006	(0.009)	–0.073	(0.054)$^{+}$
Constant	0.433	(0.492)	3.406	(3.412)
Credit constraint equation				
Single dummy (= 1, if single; = 0, otherwise)	–0.840	(0.504)*		
Co-residence dummy for single women (= 1, if living with other(s); = 0, otherwise)	1.709	(0.496)***		
Number of household members	–0.033	(0.044)		
Annual household income (1993 price, in log)	–0.257	(0.133)**		
Log income squared	0.035	(0.018)**		
Household assets (total amount of deposits plus total amount of securities) (1993 price, in log)	0.010	(0.057)		
Log asset squared	–0.006	(0.008)		
Amount of outstanding debt (1993 price, in log)	0.043	(0.048)		
Amount of outstanding debt squared (1993 price, log value squared)	–0.009	(0.008)		
Household ownership dummy (= 1, if own house; = 0, otherwise)	0.023	(0.121)		
Age of the respondent	–0.225	(0.189)		
Age squared	0.003	(0.003)		
Education level dummy (= 1 if senior high school graduate; = 0, otherwise)	–0.070	(0.167)		
Education level dummy (= 1 if vocational school graduate; = 0, otherwise)	–0.103	(0.186)		
Education level dummy (= 1 if junior college graduate; = 0, otherwise)	–0.217	(0.203)		
Education level dummy (=1 if university graduate; = 0, otherwise)	–0.236	(0.231)		
Year dummy for 1994 (= 1, if 1994; = 0, otherwise)	0.220	(0.124)*		
Year dummy for 1995 (= 1, if 1995; = 0, otherwise)	0.259	(0.140)*		
Year dummy for 1996 (= 1, if 1996; = 0, otherwise)	0.176	(0.140)		
Year dummy for 1997 (= 1, if 1997; = 0, otherwise)	0.199	(0.174)		
Year dummy for 1998 (= 1, if 1998; = 0, otherwise)	0.404	(0.165)**		
Major city dummy (= 1, if major city; = 0, otherwise)	–0.351	(0.152)**		
Other cities dummy (= 1, if other cities; = 0, otherwise)	–0.245	(0.125)*		
Constant	3.473	(3.104)		
Other parameters				
σ_N^2	0.336	(0.020)***		
σ_C^2	1.376	(0.372)***		
$\sigma_{N\varepsilon}$	0.010	(0.026)		
$\sigma_{C\varepsilon}$	0.921	(0.530)*		
Sample size		4842		

Note: Standard errors are indicated in parentheses. *** indicates significance at the 1% level; **, at the 5% level; and *, at the 10% level. + indicates the p-value of 0.179.

Table 5. Welfare Losses Arising from Binding Credit Constraints for the Year 1998

(In %; By Income Quartile; Normalized by the Expected Marginal Utility for the Year 1999)

	Income quartile			
	Bottom 25%	50%	75%	Top 25%
$E\{\lambda_{it}/E_t[U'(C_{it}+1)(1+r_{it})\,\beta\,]\}$ The result based on type 5 Tobit model with observed regime in Table X	9.91%	8.68%	8.24%	5.94%
$E\{\lambda_{it}/E_t[U'(C_{it}+1)(1+r_{it})\,\beta\,]\}$ The result based on type 5 Tobit model with observed regime in Table X	10.27%	5.31%	3.43%	2.44%

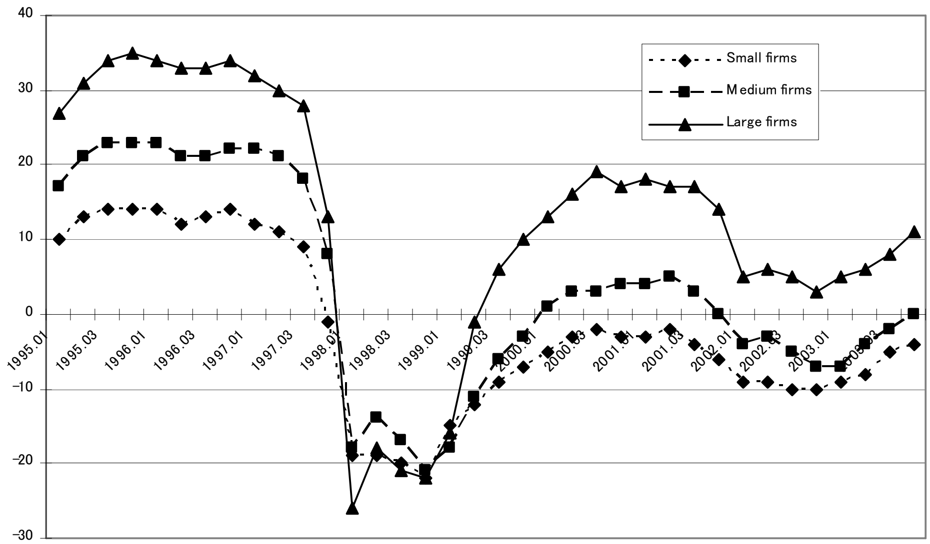
Table 6. Estimation Results of the Augmented Euler Equation with the Prudence Term Dependent Variable: Annual Growth Rate of Expenditure between 1993 and 1999

Specification	(1)	(2)	(3)	(4)		(5)		(6)
Method	IV	MLE	MLE	MLE		MLE		MLE
Period	1993–1999	1993/94 and 98/99	1993–1999	1993–1999		1993–1999		1993–1999 Without co-residents
Euler equation				Constrained	Unconstrained	Constrained	Unconstrained	
Single dummy (= 1, if single; = 0, otherwise)	0.045	0.133	0.053	–0.031	0.095	–0.034	0.106	0.164
	(0.053)	(0.061)**	(0.054)	(0.050)	(0.051)*	(0.052)	(0.057)*	(0.063)***
Co-residence dummy for single women (= 1, if living with other(s); = 0, otherwise)	0.005	–0.388	0.045	0.142	–0.148	0.148	–0.155	
	(0.057)	(0.223)*	(0.086)	(0.054)***	(0.057)***	(0.056)***	(0.061)**	
Age of the respondent	–0.041	–0.107	–0.052	–0.055	–0.060	–0.055	–0.061	–0.009
	(0.033)	(0.044)**	(0.063)	(0.030)*	(0.030)**	(0.032)*	(0.032)*	(0.055)
Age squared	0.001	0.002	0.001	0.001	0.001	0.0007	0.001	0.0003
	(0.001)	(0.001)**	(0.001)	(0.0004)	(0.0004)**	(0.0005)	(0.0005)**	(0.001)
First difference of the household size	–0.011	–0.018	–0.014	–0.022	0.023	–0.021	0.021	0.009
	(0.013)	(0.022)	(0.013)	(0.012)*	(0.022)	(0.013)*	(0.027)	(0.016)
Squared consumption growth rate	0.025	0.512	0.029	0.152	–0.097	0.151	–0.092	0.581
	(0.023)	(0.395)	(0.063)	(0.022)***	(0.017)***	(0.015)***	(0.017)***	(0.097)
Squared consumption growth rate*dummy for 1996–1999						–0.022	0.0002	–0.412
						(0.008)***	(0.007)	(0.194)
Residual from the squared consumption growth regression		–0.563	–0.019	0.004	–0.054	0.020	–0.058	–0.022
		(0.415)	(0.023)	(0.016)	(0.015)***	(0.019)	(0.015)***	(0.018)
Log income for credit-constrained households		–0.023	–0.067	–0.021		–0.020		–0.047
		(0.014)*	(0.013)***	(0.009)**		(0.009)**		(0.007)***
Log income for non-constrained households		–0.005	0.035		–0.006		–0.007	0.031
		(0.006)	(0.013)***		(0.013)		(0.013)	(0.007)***
Constant	0.702	1.683	0.913	1.012		1.013		–0.134
	(0.522)	(0.699)**	(0.939)	(0.475)**		(0.500)**		(0.919)
Credit constraint equation included?	NO	YES	YES	YES		YES		YES
Over-identification restriction tests based on Sargan pseudo-F tests	0.667							
Sample size	4842	1614	4842	4842		4842		3510

Note: Standard errors are indicated in parentheses. *** indicates significance at the 1% level; **, at the 5% level; and *, at the 10% level.

Figure 1

The Bank of Japan Diffusion Indices (DIs) of "banks' willingness to lend"



Data source: The Bank of Japan

Figure 2a

Kernel Density Function of Probability of Binding Credit Constraints in 1993 and 1998

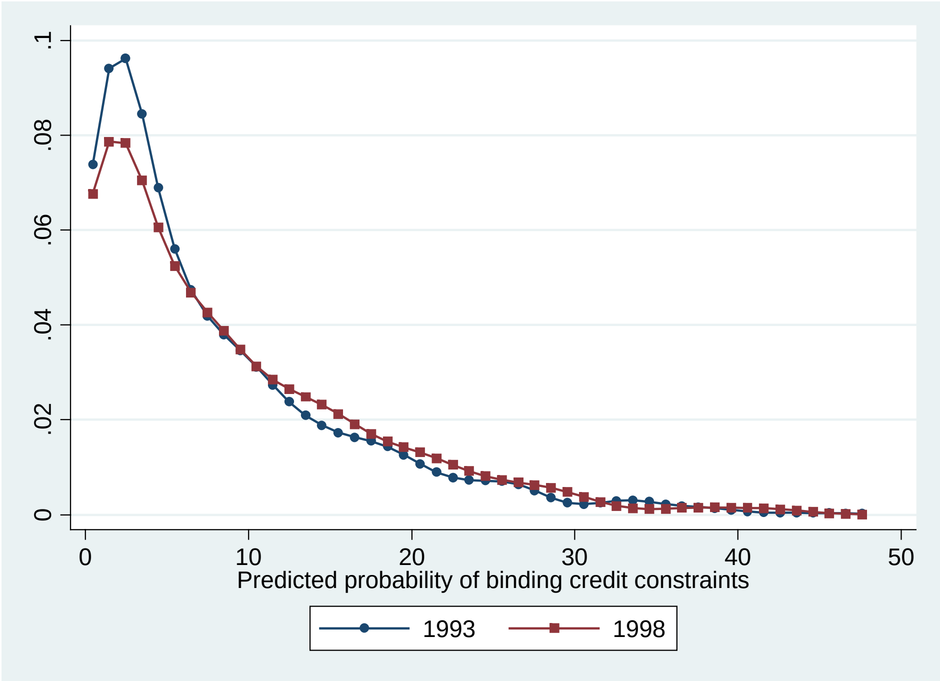


Figure 2b

Cumulative Density Function of Probability of Binding Credit Constraints in 1993 and 1998

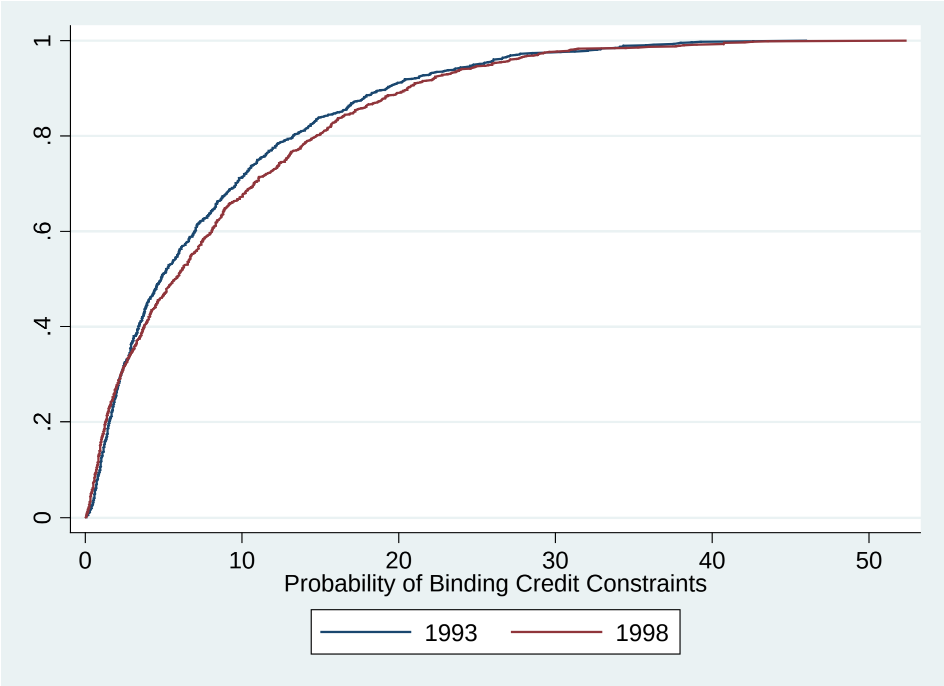


Figure 3a

Kernel Density Function of Probability of Binding Credit Constraints in 1993, 1995, and 1998

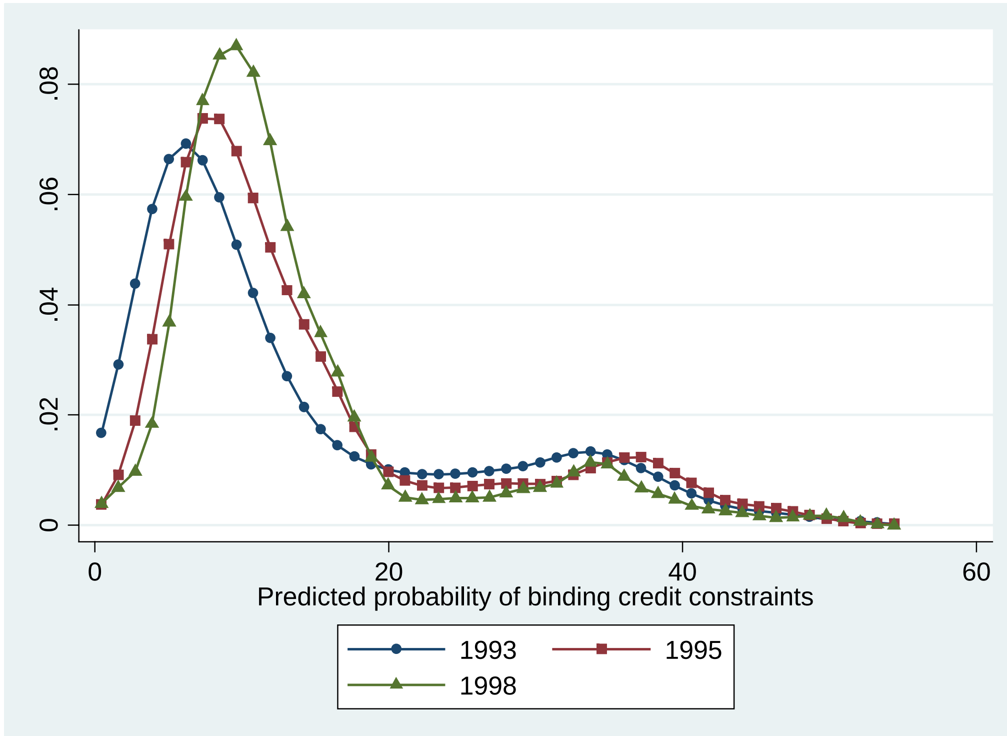


Figure 3b

Cumulative Density Function of Probability to be Credit-Constrained