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Autonomy, Incentives, and School Performance: Evidence from the 2009 Autonomous Private High School Policy in Korea†

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Improving the quality of school education is one of the key policy concerns in Korea. This paper examines whether providing schools with adequate autonomy and incentives can meet the policy goals by looking at a recent policy reform in Korea. In 2009, the Korean government granted autonomy to certain private high schools on the condition that no financial subsidies would be provided to the schools. Because the autonomous private high schools cannot receive a subsidy, they have a strong incentive to meet parental demands because schools failing to meet these demands will lose students and will have to close. Applying the value-added model to longitudinal data at the student level, I find that students entering these autonomous schools show faster growth in their academic achievement than their peers in traditional non-autonomous schools. These results suggest that providing schools with autonomy and incentives can be a useful policy tool for improving school education.

Key Word: Autonomous Private High School, School Effect, Academic Achievement, School Autonomy, Secondary Education in Korea
JEL Code: I20, I28, C21

I. Introduction

Improving the quality of education is one of the key policy concerns in Korea. A traditional approach is to provide more resources to schools, though there is growing evidence that such an input-oriented policy is not an effective means of improving school education. One explanation of the failure of such input-oriented policies is that they often fail to provide schools with adequate autonomy and

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incentives to use their resources efficiently and thus improve their educational output. For example, Barrera-Osorio *et al.* (2009) found that providing schools with autonomy has become a popular educational reform in both developed and developing countries in recent years. Hanushek (2003) reviewed many education policies around the world and concluded that creating incentives is much more effective than simply increasing inputs to schools.

This paper investigates whether, and to what extent, providing schools with adequate autonomy and incentives improves the quality of education by looking at a recent policy reform in Korea, referred to here as the autonomous private high school policy of 2009. Traditionally, private high schools in Korea have been heavily subsidized and regulated by the government. For example, private schools cannot select their own students, cannot charge higher tuitions than public schools, and must follow the uniform national curriculum. In exchange for these regulations, any financial deficits in the operating costs of private schools are fully refunded by the government. Perhaps naturally, some commentators argue that private schools have little incentive to improve their teaching technology and make efforts to meet parental demands given the strong governmental control and support (Kim and Lee 2003).

To provide better incentives to private high schools, the Korean government introduced a new type of school, termed the "autonomous private high school", in 2009. As the name suggests, these schools are characterized by a certain level of autonomy in school management in the absence of financial subsidies from the government. Because they receive no governmental subsidies, they have a strong financial incentive to meet parental demands, as schools failing to do so will lose pupils and will likely to close. Thus, one may consider the introduction of these autonomous schools as an opportunity to investigate whether, and to what extent, providing schools with autonomy and incentives -can induce them to improve the quality of their education.

To understand how autonomy and related incentives affect school performance, I compare academic achievement levels in three subjects (Korean, math, and English) of students attending autonomous schools with that of their peers who attend traditional non-autonomous schools using the value-added model suggested by Todd and Wolpin (2007). The results show that students at autonomous schools tend to show more rapid growth in academic achievement than their peers at traditional non-autonomous schools in all three subjects. To assess whether the estimated gap in achievement growth is spuriously driven by non-school educational inputs such as private tutoring, I include the amount of private tutoring expenditures as an additional control variable. In addition, in order to check whether the estimated autonomous school premium is spuriously driven by the unobserved pre-determined academic quality of the students, I also perform a falsification test using pre-determined test scores as (falsified) dependent variables. The results suggest that the estimated gap in achievement growth is mainly attributable to the type of school the students attend and not by private tutoring expenditures or the unobserved pre-determined quality of the students for Korean and math subjects. However, the parallel results for the subject of English suggest that the estimated gap in achievement growth may not reflect a causal impact. These results provide valuable but tentative evidence indicating that providing

adequate incentives and autonomy to schools may be a useful policy tool for improving school education.

II. Institutional Background and Related Literature

A. *Autonomous Private High School Policy*

Since the 1970s, most high schools in Korea have been heavily regulated and subsidized by the government. The high school equalization policy of 1974 required that virtually all high schools in large cities, either private or public, follow a set of governmental rules which outlined nearly all aspects of their operation. For example, all high schools under the equalization policy must follow a uniform national curriculum, cannot select their own students, must charge tuition amounts set by the government, and must recruit teachers and principals certified by the government. Given that the equalization policy mandated private schools to charge an equal amount of tuition - to that of public schools, any deficits in the operating costs of private schools were fully refunded by the government. Consequently, private schools became nearly "public" in their operation, and this raised serious concerns about the lack of diversity and incentives in school education (e.g., Kim and Lee 2003).

In an attempt to diversify school education and spur competition among schools, the Korean government introduced the autonomous private high school policy in 2009. In this policy, autonomous private high schools are granted a greater degree of autonomy in their operation compared to traditional non-autonomous high schools. Essentially, these schools have substantial autonomy in many aspects of school management, including student/teacher/principal recruitment, tuition amounts, curriculum, textbooks, and academic terms (i.e., their choice of semester, trimester, or quarter systems). However, in order to enjoy this autonomy, the schools must bear the following responsibilities. First, autonomous schools cannot receive any financial subsidies from the government. Thus, they have a strong incentive to meet parental demand because schools failing to enroll enough students will not be able to finance their operations and will thus have to close. Second, they cannot charge more than three times the tuition of traditional non-autonomous schools. Third, they must reserve at least 20 percent of their places for students from low-income families. Fourth, they are allowed to select their own students, but not through entrance exams or interviews about academic knowledge. Instead, autonomous schools may select their students according to middle school grades, (non-academic) interviews, recommendation letters, and a lottery system. Particularly, the autonomous schools in Seoul in 2013, which constitute the main subject of this study, select their students in a lottery among applicants whose middle school grades are above the 50th percentile within their middle schools. Finally, the licenses of the autonomous schools must be re-evaluated by the local government (i.e., the education superintendent) every five years. If the schools do not pass this evaluation, they are converted to traditional non-autonomous schools.

Between 2009 and 2011, the Korean government designated 49 autonomous

private high schools, mostly from existing traditional non-autonomous schools across the nation. More than half (25 out of 49) of the autonomous schools operate in Seoul, and the present study focuses on these. Since the introduction of autonomous schools, however, there has been heated debate as to whether the autonomous private high school policy should be maintained. Advocates of the policy argue that it can improve the overall quality of school education by spurring competition among schools. In contrast, opponents emphasize that autonomous schools can deteriorate the educational equity by providing better educational services only to those who can afford higher tuition levels of the autonomous schools. Particularly, as candidates pledging to abolish autonomous schools were elected in many cities and provinces in the 2014 local education superintendent elections,[1] the conflict between advocates and opponents became even more serious. These conflicts garnered media headlines and were followed by a series of lawsuits involving parents, autonomous schools, local education superintendents, and the central government (i.e., the Ministry of Education).

B. *Empirical Studies of Autonomous Private High School*

In spite of the heated debate, surprisingly little is known about how autonomous schools affect the educational performance of students. Kim and Namkung (2014) evaluated the impact of autonomous private high schools on the academic achievement of their students. They compared the educational performance of students in autonomous schools with that of their peers in traditional non-autonomous schools after controlling for students' family backgrounds. They conclude that there is a large gap in academic achievement (with standard deviations of approximately 0.6~0.7 and 0.7~1.0 for reading and math, respectively) between the two types of schools. However, given that autonomous schools select students based on their middle school grades, recommendations, interviews and related factors, there may be unobserved differences in students' pre-determined academic quality levels that are not captured by their family backgrounds. In this respect, the estimated autonomous school premium reported by Kim and Namkung (2014) is likely to be overestimated.

Lee and Shin (2014) attempt to evaluate the spillover effect of autonomous schools. Specifically, they estimate how the designation of autonomous schools affects the academic achievement of the incumbent students in these schools (i.e., students entering the autonomous schools before the designation) and the incumbent students in the closest non-autonomous schools (i.e., students entering neighboring non-autonomous schools before the designation). They found that the designation of an autonomous school does not affect the academic achievement of incumbent students within the schools, whereas itnegatively affected the academic achievement of students in neighboring non-autonomous schools.

This study contributes to the literature by examining how autonomous schools affect the academic achievement of students. Although understanding whether and

[1]Out of 17 local education superintendent elections, candidates pledging to shut down autonomous schools won 13 elections, including that in Seoul. The exceptions were Daegu, Ulsan, Daejeon, and Gyungbuk.

to what extent autonomous schools can improve the academic achievement of students would be the first step towards an evaluation of the desirability of the policy, there is surprisingly little evidence on this issue. Kim and Namkung (2014) reported a large impact, but their estimates are likely to be upwardly biased. Lee and Shin (2014) analyzed the spillover effect of autonomous schools but not the direct effect (i.e., how they affect students who enrolled in them after the designation), which constitutes the main objective of this study.

III. Data

My empirical analysis relies on the Seoul Education Longitudinal Study of 2010 (SELS 2010). SELS 2010 has tracked three cohorts (students in the fourth, seventh, and tenth grades as of 2010) of pupils in Seoul annually since 2010. Table 1 summarizes the survey timings of SELS 2010.

Among the three cohort samples, I use the seventh-grade sample in this study. As shown in Table 1, students in the seventh-grade sample were surveyed from the seventh grade (i.e., their first year of middle school) to the eleventh grade (i.e., their second year of high school). This allows an estimation of the effect of attending an autonomous private high school on academic achievement after controlling for pre-determined achievement as measured during the middle school years.

I restrict my estimation sample to students who entered either autonomous private high schools or traditional non-autonomous high schools. Students who entered vocational and special-purpose high schools are excluded from the estimation sample because these schools are considerably different from autonomous and non-autonomous schools in many aspects of school operation other than incentives.

For outcome variables, I consider test scores in Korean (reading), math, and English. The test scores were originally recorded on a scale of 100. I standardize these scores by subtracting the means and dividing by the standard deviations, with the results used as outcome variables.

Table 2 summarizes the descriptive statistics of my sample. The table clearly shows that students in autonomous schools tend to perform better than their peers in traditional non-autonomous schools. The achievement gap between the two groups ranges from approximately 0.7 to 1.0 in terms of the standard deviation in both the tenth and eleventh grades. However, it should be noted that students at autonomous high schools had outperformed their peers in traditional non-autonomous high schools *before* entering their high schools. Table 2 clearly shows that the students at autonomous high school had already scored higher than their

TABLE 1—SURVEY YEARS OF THE SEOUL EDUCATIONAL LONGITUDINAL STUDY 2010 (SELS 2010)

School level	Elementary School			Middle School			High School		
Cohort sample	Grade								
	4th	5th	6th	7th	8th	9th	10th	11th	12th
Fourth-grader sample	2010	2011	2012	2013	2014				
Seventh-grader sample				2010	2011	2012	2013	2014	
Tenth-grader sample							2010	2011	2012

TABLE 2—SUMMARY STATISTICS

Variable	Autonomous			Non-autonomous		
	N	Mean	S.D.	N	Mean	S.D.
Eleventh-grade test score (Z score)						
Korean	264	0.66	0.94	1990	-0.04	0.96
Math	264	0.80	1.19	1989	-0.09	0.91
English	263	0.82	1.02	1988	-0.09	0.93
Tenth-grade test score (Z score)						
Korean	262	0.71	0.86	1968	-0.05	0.97
Math	264	0.92	0.89	1964	-0.08	0.94
English	263	0.88	0.84	1969	-0.09	0.94
Ninth-grade test score (Z score)						
Korean	270	0.59	0.90	2002	0.08	0.96
Math	269	0.82	0.83	1995	0.08	0.98
English	270	0.72	0.79	2005	0.10	0.95
Eighth-grade test score (Z score)						
Korean	270	0.56	0.81	2009	0.05	0.97
Math	270	0.84	0.87	2009	0.08	0.96
English	270	0.71	0.81	2009	0.08	0.94
Seventh-grade test score (Z score)						
Korean	270	0.52	0.86	2001	0.06	0.96
Math	270	0.77	0.84	1997	0.08	0.94
English	269	0.73	0.80	1999	0.07	0.94
Female (yes=1)	270	0.22	0.41	2007	0.50	0.50
Disabled (yes=1)	268	0.02	0.14	2005	0.04	0.19
Number of siblings	270	2.08	0.41	2000	2.14	0.52
First-born (yes=1)	269	0.10	0.30	1988	0.12	0.33
Single parent (yes=1)	270	0.05	0.21	2008	0.09	0.29
Father's education (yes=1)						
Less than high school	270	0.01	0.09	1935	0.03	0.16
Some college	270	0.10	0.30	1935	0.13	0.34
College graduate	270	0.51	0.50	1935	0.44	0.50
Graduate school or more	270	0.21	0.41	1935	0.11	0.31
Mother's education (yes=1)						
Less than high school	268	0.02	0.14	1977	0.03	0.16
Some college	268	0.15	0.36	1977	0.16	0.37
College graduate	268	0.47	0.50	1977	0.34	0.47
Graduate school or more	268	0.10	0.30	1977	0.04	0.18
Father's age (yes=1)						
49 or younger	267	0.04	0.20	1931	0.06	0.23
60 or older	267	0.12	0.33	1931	0.12	0.33
Mother's age (yes=1)						
49 or younger	269	0.23	0.42	1986	0.23	0.42
60 or older	269	0.03	0.18	1986	0.04	0.20
Parents' employment (yes=1)						
Only mother employed	269	0.05	0.22	2001	0.07	0.26
Both employed	269	0.44	0.50	2001	0.52	0.50
Neither employed	269	0.00	0.00	2001	0.01	0.08
Parental income (10,000 KRW/month)	270	586.02	421.76	2009	497.53	522.05
Parental income missing (yes=1)	270	0.02	0.15	2009	0.04	0.20
Male-only school (yes=1)	270	0.71	0.46	2009	0.31	0.46
Female-only school (yes=1)	270	0.12	0.32	2009	0.34	0.48
Private school (yes=1)	270	1.00	0.00	2009	0.65	0.48
Eleventh-grade private tutoring (10,000 KRW/month)						
Korean	234	12.00	24.29	1745	6.25	11.52
Math	246	34.38	29.48	1814	22.38	23.25
English	245	26.22	38.64	1803	18.16	22.41
Tenth-grade private tutoring (10,000 KRW/month)						
Korean	221	9.63	26.02	1635	6.47	11.36
Math	255	33.04	30.75	1855	22.56	23.49
English	242	25.69	30.74	1814	20.00	23.02

non-autonomous high school counterparts by about 0.6 ~ 0.9 standard deviations during their middle school years (i.e., their seventh, eighth, and ninth grades).

In terms of student characteristics, the proportion of female students is substantially lower in the autonomous schools (22%) than in the traditional non-autonomous schools (50%). This is largely because most of the autonomous schools in Seoul are male-only schools. Among the 25 autonomous schools operating in Seoul in 2013, 17 are male-only schools, five schools are co-educational, and only three schools are female-only schools. To account for this difference in gender composition between the autonomous schools and the non-autonomous schools, I control for gender and for the gender composition of the schools (i.e., male-only, female-only, and co-educational) in my regression analysis. In terms of family characteristics, students in the autonomous schools report higher parental income and educational attainment levels than their peers in traditional non-autonomous schools. Additionally, students in autonomous schools outspend their counterparts in traditional non-autonomous schools on private tutoring.

IV. Empirical Analysis

A. *Identification Issue*

Ideally, the causal effect of attendance at an autonomous school on student outcomes could easily be identified if admissions to autonomous schools were randomly determined. In fact, in 2013, the autonomous schools in Seoul admitted students *by lottery*. This indicates that, *among the participants in the applicant lottery,* admissions to the autonomous schools were randomly assigned. However, whether a student applied for entry into an autonomous school was clearly non-randomly determined. In 2013, only top 50 percent of students in terms of their middle school grades were able to apply for entry into an autonomous school. In addition, autonomous schools charged two to three times the tuition of non-autonomous regular high schools. These facts suggest that the identification of a causal effect of attending an autonomous school depends on how much one can control for middle school grades and the parental income of students as well as their preferences for an autonomous school.

B. *Empirical Model*

To address these concerns, I attempt to identify the autonomous school premium by controlling for ninth-grade (i.e., the third year of middle school) test scores of students along with other background characteristics that are likely to be correlated with their application decisions. Specifically, I estimate the following "value-added" model:

$$(1) \quad Y_{i,m,s,10th} = \beta_0 + \beta_1 Autonomy_s + \beta_2 Y_{i,m,s,9th} + X_{i,9th}\beta_3 + W_s\beta_4 + \rho_m + \varepsilon_{i,m,s,10th}$$

In equation (1), $Y_{i,m,s,10th}$ indicates the tenth-grade (i.e., first year in high school) test scores of student *i* in high school *s* who graduated from middle school *m*. $Autonomy_s$ is a dummy variable that takes a value of 1 if high school *s* is an autonomous school and 0 otherwise (i.e., a traditional non-autonomous school) as of the year 2013.[2] $Y_{i,m,s,9th}$ represents the ninth-grade (i.e., third year in middle school) test scores of student *i*. The lagged test scores are intended to capture the minimum required condition to apply for entry into an autonomous school (i.e., the top 50% in terms of middle school grades) and the difference in the pre-determined academic quality of students between the autonomous and the non-autonomous schools. $X_{i,9th}$ refers to the baseline characteristics of student *i* measured in the ninth grade when the student decided upon the high schools to which he would apply. Specifically, $X_{i,9th}$ includes variables on student characteristics (gender, disability, birth order) and family background (parental age, parental education, parental employment status, parental income, number of siblings, single parent). ρ_m represents middle school fixed effects, capturing any unobservable heterogeneity that students from the same middle schools have in common. In Seoul, nearly all students graduating from their elementary schools are assigned to their neighborhood middle schools. This suggests that ρ_m will also contain a substantial amount of information on the students' residential locations. W_s refers to the characteristics of high school *s*, such as the gender composition (co-educational, male-only, female-only) and the establishment type (private or public). Given that autonomous schools are all private and mostly single-sex schools, controlling for these characteristics is particularly important for distinguishing the effect of school autonomy. Finally, $\varepsilon_{i,m,s,10th}$ is an error term.

C. *Estimation Results*

I begin by estimating equation (1) with OLS, clustering standard errors at the middle school level. Table 3 shows the estimation results when the Korean, math, and English test scores are used as outcome variables. Column (1) of Table 3 shows the simple regression results without any covariate, which basically compares the average test scores of autonomous school students with those of non-autonomous school students. On average, students in autonomous schools achieve higher test scores than their peers in traditional non-autonomous schools by about 0.76, 1.00, and 0.97 standard deviations in Korean, math, and English, respectively. In column (2), I include student characteristics (gender, disability, first-born child), family characteristics (number of siblings, single parent, parental age, parental education, parental employment status, and parental income), high school characteristics (gender composition, establishment type), and dummies for the middle schools from which the students graduated (i.e., middle school fixed effects) as control

[2]Because all students in my estimation sample graduated from their middle schools in February of 2013 and entered their high schools in March of 2013, the year 2013 corresponds to the first year they were in high school.

TABLE 3—ESTIMATION RESULTS FOR THE TENTH-GRADE (FIRST YEAR IN HIGH SCHOOL) TEST SCORES

	(1)	(2)	(3)	(4)	(5)
Estimation	OLS	OLS	OLS	2SLS	2SLS
Covariates	None	Student, family, school characteristics	(2) + 9th grade test scores	Same as (3)	(4) + private tutoring expenditures
			A. Korean test scores		
Autonomous school	0.760***	0.688***	0.393***	0.176**	0.175**
	(0.057)	(0.065)	(0.057)	(0.080)	(0.088)
Observations	2,230	2,061	2,056	2,049	1,675
R-squared	0.062	0.265	0.500	0.308	0.299
First-stage F				453.4	327.9
			B. Math test scores		
Autonomous school	1.003***	0.877***	0.505***	0.320***	0.311***
	(0.059)	(0.068)	(0.059)	(0.076)	(0.073)
Observations	2,228	2,059	2,047	2,037	1,899
R-squared	0.107	0.252	0.551	0.419	0.432
First-stage F				694.9	625
			C. English test scores		
Autonomous school	0.969***	0.839***	0.527***	0.331***	0.346***
	(0.056)	(0.063)	(0.055)	(0.066)	(0.065)
Observations	2,232	2,063	2,061	2,052	1,862
R-squared	0.101	0.341	0.575	0.401	0.387
First-stage F				699.7	625.6

Note: Robust standard errors clustered at the middle school level are in parentheses (*** p<0.01, ** p<0.05, * p<0.1). Student, family, and school characteristics include gender, disability, first-born child, number of siblings, single parent, parental age, parental education, parental employment status, parental income, high school gender composition (male-only, female-only, co-educational), high school establishment type (private, public), and dummies for middle schools from which the students graduated (i.e., middle school fixed effects).

variables. When these characteristics are controlled, the estimated test score gap between the two groups is slightly reduced to about 0.69, 0.88, and 0.84 standard deviations in Korean, math, and English, respectively. These estimates are comparable to those reported in Kim and Namkung (2014) (0.6~0.7 and 0.7~1.0 standard deviations in Korean and math, respectively), who mainly estimated the impact of autonomous schools on the academic achievement of students by regressing test scores on school types after controlling for student, family, and school characteristics. The estimation results in column (2), in conjunction with the results in column (1), also indicate that the student, family, and school characteristics can only account for approximately 10% of the observed test score gap between autonomous school students and non-autonomous school students.

Exploiting the longitudinal structure of my data, in column (3) I include students' pre-determined academic quality levels as measured by their ninth-grade (i.e., third year in middle school) test scores as controls added to the list of controls used in column (2). As discussed in the chapter II, only the top 50% of students in terms of their middle school grades could apply for entry into autonomous schools. Hence, students in autonomous schools are likely to perform better than their peers in non-autonomous schools even before they enter high schools. Adding lagged test scores as an additional control could control for this differences in the pre-determined academic quality levels between the two groups. When the baseline test scores are further controlled, the estimated achievement gap between the two

groups is reduced substantially to about 0.39, 0.51, and 0.53 standard deviations in Korean, math, and English, respectively. These results suggest that Kim and Namkung (2014) likely overestimated the achievement effect of the autonomous schools by ignoring the differences in the pre-determined academic quality between the autonomous school students and the traditional non-autonomous school students.

In columns (1) to (3), I estimated equation (1) with the OLS method using different sets of covariates. Econometrically, however, estimating equation (1) with OLS will result in an inconsistent estimate when the error term $(\varepsilon_{i.m.h.10th})$ is serially correlated with its lagged term $(\varepsilon_{i.m.h.9th})$ because equation (1) includes a lagged dependent variable $(Y_{i.m.h.9th})$ as a regressor. To address this issue, I instrument the potentially endogenous ninth-grade test scores $(Y_{i.m.h.9th})$ with seventh-grade test scores $(Y_{i.m.h.7th})$. This allows the error term $(\varepsilon_{i.m.h.10th})$ to follow a "mild" serial correlation (i.e., AR(1) or AR(2) process) but not a "severe" one (i.e., AR(p) process with p≥3). Column (4) of Table 1, which is my most preferred specification, shows the two-stage least-square (2SLS) estimation results using seventh-grade test scores $(Y_{i.m.h.7th})$ as an instrument variable for ninth-grade test scores $(Y_{i.m.h.7th})$. The impacts of attending an autonomous school on Korean, math, and English test scores are estimated to be 0.18, 0.32, and 0.33 standard deviations, respectively. Comparing these 2SLS estimates with the OLS estimates reported in column (3) reveals that the serial correlation issue discussed above is indeed serious.

Table 2 shows that students in autonomous schools tend to spend more on private tutoring than their peers in traditional non-autonomous schools. To the extent that private tutoring may improve the academic achievement of students, as discussed in a number of recent studies (e.g., Kang, 2012; Ryu and Kang, 2013), the estimated autonomous school premium reported in column (4) of Table 3 could be spuriously driven by the differences in private tutoring investment. To check for this possibility, I add the amount of private tutoring expenditures for each subject as an additional control variable in column (5). Even after controlling for these non-school educational inputs, the estimated autonomous school premium remains similar, indicating that the estimates reported in column (4) are largely attributable to the type of high schools the students attend and not to differences in private tutoring expenditures. Finally, in Table 4, I repeat the above-mentioned analysis using eleventh-grade (i.e., second year in high school) test scores as an outcome variable. Specifically, I estimate equation (1) using $Y_{i.m.h.11th}$ as the left-hand-side variable instead of $Y_{i.m.h.10th}$. The results are roughly similar to those reported in Table 3. In terms of my preferred specification (column 4), attending an autonomous school improve tenth-grade test scores by 0.18, 0.32, and 0.33 standard deviations and eleventh-grade test scores by 0.24, 0.28, and 0.34 standard deviations in Korean, math, and English, respectively.

TABLE 4—ESTIMATION RESULTS FOR ELEVENTH-GRADE (SECOND YEAR IN HIGH SCHOOL) TEST SCORES

	(1)	(2)	(3)	(4)	(5)
Estimation	OLS	OLS	OLS	2SLS	2SLS
Covariates	None	Student, family, school characteristics	(2) + 9th grade test scores	Same as (3)	(4) + private tutoring expenditures
			A. Korean test scores		
Autonomous school	0.700***	0.655***	0.413***	0.242***	0.241***
	(0.062)	(0.070)	(0.061)	(0.075)	(0.079)
Observations	2,254	2,085	2,080	2,073	1,810
R-squared	0.053	0.235	0.424	0.278	0.295
First-stage F				464.5	409.3
			B. Math test scores		
Autonomous school	0.889***	0.730***	0.455***	0.280***	0.256***
	(0.076)	(0.082)	(0.076)	(0.091)	(0.091)
Observations	2,253	2,085	2,072	2,061	1,877
R-squared	0.084	0.186	0.373	0.239	0.252
First-stage F				661.3	643
			C. English test scores		
Autonomous school	0.911***	0.755***	0.500***	0.343***	0.370***
	(0.066)	(0.072)	(0.066)	(0.080)	(0.086)
Observations	2,251	2,082	2,080	2,070	1,875
R-squared	0.088	0.279	0.452	0.315	0.317
First-stage F				697.5	651.1

Note: Robust standard errors clustered at the middle school level are in parentheses (*** p<0.01, ** p<0.05, * p<0.1). Student, family, and school characteristics include gender, disability, first-born child, number of siblings, single parent, parental age, parental education, parental employment status, parental income, high school gender composition (male-only, female-only, co-educational), high school establishment type (private, public), and dummies for middle schools from which the students graduated (i.e., middle school fixed effects).

D. *Falsification Test*

Tables 3 and 4 show that students attending the autonomous schools outperform their peers attending non-autonomous schools after controlling for student, family and school characteristics (including middle school fixed effects), baseline academic performance, and private tutoring expenditures. However, whether the estimated achievement gap between the two groups of students reflects the *causal* effect of attending an autonomous school remains questionable. For example, it is still possible that the estimated achievement gap reflects unobservable differences in pre-determined academic quality levels between the two groups of students.

In order to determine whether equation (1) correctly identifies the causal effect of attending an autonomous school, I perform the following falsification test. Specifically, I estimate the impact of attending an autonomous school on the *pre-determined* academic performance of students. Specifically, I replace the outcome variable of equation (1) with eighth-grade test scores $(Y_{i.m.h.8th})$. Given that the eighth-grade test scores were determined *before* the students entered high school, the autonomous school attendance of students cannot causally affect their eighth-grade test scores.

Table 5 shows the falsification test results. Columns (1) to (3) report positive and

TABLE 5—FALSIFICATION TEST RESULTS FOR PRE-DETERMINED EIGHTH-GRADE (SECOND YEAR IN MIDDLE SCHOOL) TEST SCORES

	(1)	(2)	(3)	(4)	(5)
Estimation	OLS	OLS	OLS	2SLS	2SLS
Covariates	None	Student, family, school characteristics	(2) + 9th grade test scores	Same as (3)	(4) + private tutoring expenditures
	A. Korean test scores				
Autonomous school	0.504***	0.482***	0.210***	-0.050	-0.018
	(0.054)	(0.059)	(0.053)	(0.065)	(0.075)
Observations	2,279	2,109	2,104	2,097	1,444
R-squared	0.028	0.259	0.486	0.144	0.165
First-stage F				483.4	321.2
	B. Math test scores				
Autonomous school	0.753***	0.603***	0.269***	-0.028	-0.040
	(0.057)	(0.065)	(0.055)	(0.071)	(0.075)
Observations	2,279	2,109	2,096	2,085	1,905
R-squared	0.062	0.248	0.519	0.201	0.189
First-stage F				680.1	505.4
	C. English test scores				
Autonomous school	0.633***	0.465***	0.155***	-0.117*	-0.121*
	(0.054)	(0.057)	(0.050)	(0.065)	(0.066)
Observations	2,279	2,109	2,107	2,097	1,893
R-squared	0.046	0.305	0.567	0.222	0.218
First-stage F				725.9	640.5

Note: Robust standard errors clustered at the middle school level are in parentheses (*** p<0.01, ** p<0.05, * p<0.1). Student, family, and school characteristics include gender, disability, first-born child, number of siblings, single parent, parental age, parental education, parental employment status, and parental income, high school gender composition (male-only, female-only, co-educational), high school establishment type (private, public), and dummies for the middle schools from which the students graduated (i.e., middle school fixed effects).

statistically significant impacts of the autonomous high schools, indicating that the corresponding regression equations are likely to be misspecified. On the other hand, in columns (4) and (5), my preferred specifications, I do not find any statistically significant effect for Korean and math. These results suggest that the estimated achievement gaps in the Korean and math test scores reported in columns (4) and (5) of Tables 3 and 4 are not driven by model misspecifications but instead reflect the causal effects of attending an autonomous school. For English test scores, however, the estimates from the falsification test are marginally significant at the 10% level, suggesting that the estimation results for English test scores should be interpreted with caution.

E. *Subgroup Analysis*

As discussed in chapter II, autonomous schools can charge up to three times the tuition of traditional non-autonomous schools in exchange for receiving no governmental subsidies. This feature raises the serious public concern that the autonomous schools can only serve students from high-income families. In this

TABLE 6—ESTIMATED EFFECTS OF ATTENDING AN AUTONOMOUS SCHOOL BY INCOME LEVEL

	(1)	(2)	(3)	(4)	(5)	(6)
		Low-income			High-income	
Outcome variable	10th grade score	11th grade score	9th grade score (falsification)	10th grade score	11th grade score	9th grade score (falsification)
			A. Korean test scores			
Autonomous school	0.107	0.250**	-0.159	0.231**	0.264**	0.034
	(0.158)	(0.107)	(0.131)	(0.106)	(0.112)	(0.107)
Observations	777	832	683	898	978	761
R-squared	0.281	0.337	0.258	0.315	0.291	0.096
First-stage F	135.9	192.8	102.8	209	209	199
			B. Math test scores			
Autonomous school	0.410***	0.285**	-0.001	0.227**	0.221*	-0.027
	(0.099)	(0.138)	(0.123)	(0.095)	(0.120)	(0.089)
Observations	869	858	865	1,030	1,019	1,040
R-squared	0.457	0.241	0.157	0.410	0.248	0.222
First-stage F	274.7	278.4	242.3	364.7	303.9	331.5
			C. English test scores			
Autonomous school	0.282***	0.521***	-0.058	0.378***	0.281**	-0.181*
	(0.097)	(0.114)	(0.081)	(0.088)	(0.112)	(0.095)
Observations	854	861	863	1,008	1,014	1,030
R-squared	0.436	0.419	0.345	0.345	0.249	0.071
First-stage F	332.5	481.4	382.4	252.9	259.3	265.7

Note: Robust standard errors clustered at the middle school level are in parentheses (*** p<0.01, ** p<0.05, * p<0.1). Student, family, and school characteristics of gender, disability, first-born child, number of siblings, single parent, parental age, parental education, parental employment status, parental income, high school gender composition (male-only, female-only, co-educational), high school establishment type (private, public), and dummies for middle schools from which the students graduated (i.e., middle school fixed effects), with the amount of private tutoring expenditures controlled. The low-income group refers to students whose parental monthly income is below KRW 4,500,000. The high-income group refers to those whose parental monthly income is greater than or equal to KRW 4,500,000.

respect, it would be worthwhile to determine how the observed autonomous school premium varies across students' family backgrounds.

To address this issue, I divide the estimation sample into the two subgroups of a high-income sample and a low-income sample. The high-income sample consists of students whose parental income in 2012, when the students were enrolled in the ninth grade, or their third year of middle school, is greater than or equal to or the median (KRW 4,500,000). Accordingly, the low-income sample consists of students whose parental income in the ninth grade is below the median. For each subgroup, I estimate the value-added model of equation (1) using the 2SLS method after controlling for the amount of private tutoring expenditures. This specification is comparable to the regression model used for column (5) of Tables 3, 4, and 5.

Table 6 summarizes the estimation results. Overall, I do not find any clear evidence that the autonomous school premium varies according to students' family backgrounds. These results suggest that the benefits that accrue from school autonomy and incentives can be enjoyed by all students regardless of their family background.

V. Conclusion

In this paper, I estimate the causal effect of attending an autonomous private high school on the academic achievement of students. The autonomous private high school policy has been one of the most controversial educational policy issues in recent years. The 2014 local educational superintendent election sparked much heated debate about whether or not to abolish autonomous schools. The conflicts were followed by a series of lawsuits involving parents, autonomous schools, local offices of education, and the central government (i.e., the Ministry of Education). In spite of these serious conflicts, however, surprisingly little is known about how these autonomous schools affect students. Applying the value-added model by Todd and Wolpin (2007) to the longitudinal data at the student level, I find that autonomous schools more effectively improve the academic achievement of students by approximately 0.2 ~ 0.3 standard deviations relative to traditional non-autonomous high schools. A key feature of the autonomous schools is that they can operate free from governmental control but at the cost of foregoing financial subsidies from the government. Because autonomous schools do not receive governmental subsidies, they have a strong financial incentive to improve the quality of their education, as schools failing to do so will lose pupils and will likely close as a result. In this respect, the estimated autonomous school premium suggests that providing schools with adequate autonomy and incentives can induce them to become more productive.

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CEO Compensation and Concurrent Executive Employment of Outside Directors: A Panel Data Analysis of S&P 1500 firms

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In many advanced countries, most outside directors are executives, active or retired, at other firms; in other words, executives from other companies make executive compensation decisions. This situation may hinder the board of directors (BOD) in their efforts to optimize executive compensation levels objectively. Using a panel data analysis of the S&P 1500 companies, we provide supplemental evidence of whether, and to what extent, the concurrent executive employment of outside directors distorts the executive pay decisions at a given company. An unbiased fixed-effect estimation confirms that a $1.00 increase in CEO pay at outside directors' primary companies results in an approximate increase of $0.22 in CEO pay at the given company. From a policy perspective, this added agency problem — caused by the BOD and not by management — is noted as difficult to control; although a firm may establish board independence, the inherent concurrent employment of directors on a board continues to exist.

Key Word: CEO Compensation, Director-Agency Problem, Outside Directors, Board of Directors, Corporate Governance
JEL Code: M12, G34, G38

I. Introduction

Executive compensation has attracted significant attention from economists and business experts since the early 1990s. News stories have questioned the fairness of executive pay practices and articles have been published about inappropriate pay practices, spurring pressure for oversight and regulation

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(e.g., Boyd 1994; Hall and Murphy 2003; Jensen *et al.* 2004). As delegates of shareholders, board members are entitled to monitor and control managers to minimize agency costs. However, boards of directors (BOD) have not been very effective, partly because they have been captured by chief executive officers (CEO) and because of the directors' own agency problem; just as no reason exists to presume that managers automatically seek to maximize shareholder value, no reason exists to expect *a priori* that directors will also do so. Bebchuk and Fried (2003) refer to this phenomenon as the *director–agency problem*.

For instance, most outside directors have no significant vested interest in the firm and, therefore, need not be very careful when deciding on company expenditure amounts, including executive pay (Baker *et al.* 1988; Cyert *et al.* 2002; and Brenner and Schwalbach 2009). Furthermore, they may not want to hurt their personal or business relationships over stingy decisions about compensation. If outside directors are appointed owing to personal ties, they may even fear losing their board seats.

Additionally, it is noteworthy that most outside directors are executives, active or retired, at other firms in the United States and in many other advanced countries, including the United Kingdom, Australia and the Netherlands. In other words, executives from other companies make executive compensation decisions in those countries. This situation may significantly affect the executive compensation decision-making process. That is, directors' status as executives at other firms can be another source of the director–agency problem.

Numerous governance experts describe their concerns about this additional agency problem. For example, Jensen *et al.* (2004) recommend that boards limit the number of CEO-directors who are outside directors but serve as CEOs for other firms, as such outside directors tend to "subconsciously (if not consciously) view the board through the eyes of a CEO."[1] However, solid empirical support of their concerns is relatively rare, except for much of the "social network" literature that examines the influence of connections between board members and executives on the level of CEO compensation (e.g., Hwang and Kim 2009; Bizjak *et al.* 2009; Larcker *et al.* 2005; and Hallock 1997).

In contrast, business practitioners often claim that this concern could be imaginary and that their policy suggestions would cause over-regulation of BOD operations. Instead, they emphasize that having other top executives as directors on the board provides the BOD with numerous advantages related to its evaluation of business strategies, voting on major operational proposals, and the scrutinizing of financial and accounting reports. Their claims are partly supported by positive stock price reactions to director appointments when the appointee is an active CEO (Fich 2005).

The empirical work in this paper provides supplemental evidence of whether, and to what extent, the concurrent executive employment of outside directors distorts pay decisions at a given company. We collected a consecutive two-year compensation dataset of the S&P 1500 firms from Compustat Execomp. Using the Board Analyst database, each company listed in the S&P 1500 was matched to its

[1]Faleye (2011) argues that CEO-directors may overestimate the effort and skill requirements of the executive job and rationalize higher compensation packages for top executives.

outside directors' primary companies for which they serve (or served) as executives. We then calculated the average CEO pay of matched outside directors' primary companies. Based on this constructed sample, we tested whether the average level of CEO pay at directors' primary companies affects the level of CEO pay at a given company. If this finding is insignificant, we may not be able to state that the characteristics of outside directors' primary companies directly distort the executive pay decisions at a given company.

In our analysis, the test result was statistically significant. Using the fixed effects model with a panel data analysis, which is identical to the first-difference model in the given consecutive two-year sample, we observe a strong link between CEO pay at outside directors' primary companies and CEO pay at a given company; a $1.00 increase in CEO pay at outside directors' primary companies leads to an approximate $0.22 increase in CEO pay at a given firm. We also observe that a 1% increase in CEO pay at outside directors' primary companies is associated with a 0.13% increase in CEO pay at a given firm.

We obtained these test outcomes even after controlling for the size (and the entrenchment) associations between outside directors' primary companies and each given company. Taking all of the interlocking cross-directorship observations out of the samples made little difference in the test outcomes. The observed strong link may be generated by several other sources. For instance, the positive association between CEO pay levels is well supported by a psychological phenomenon known as the *anchoring effect* (Tversky and Kahneman 1974), which holds that outside directors may use the CEO pay levels at their primary companies as firsthand reference points when making CEO pay decisions at a given firm. It is also plausible that the homogeneity and cohesiveness of top executives, termed the *inner circle*, may affect the observed positive association between CEO pay levels.

Our empirical findings are noteworthy from a policy perspective because the additional agency problem within the supposedly "independent" BOD is difficult to control; despite the fact that we establish board independence through various imposed regulations and shareholder activism, this inherent characteristic of concurrent employment among board members continues to exist in the corporate governance systems of many advanced countries, unless the government imposes direct regulations on the board structure.

The paper is organized into the following sections. Section II describes the data sources and the procedure used to construct the dataset. Section III covers the methodologies used in the analysis. Section IV presents the test results. Section V discusses the implications of the findings, and Section VI concludes the paper.

II. Dataset and Variables

In this section, we describe the data sources and the process of constructing the dataset and explain the key variables used in the analysis. The Compustat Execomp (also known as ExecuComp) and Board Analyst databases were used as sources of the board and ownership characteristics and the economic information pertaining to the selected firms. Compustat Execomp, which began in 1994, provides data on compensation for the top five executives and basic economic performance metrics,

primarily for S&P 1500 companies. Board Analyst, which began in 2001, is a source of comprehensive, objective corporate governance and compensation information for more than 10,000 U.S. companies.

A. *Sample and Datasets*

The research in this paper covers the early 2000s because we use the first samples that the Board Analyst started to provide. More specifically, we constructed a dataset that contains two-year panel data (2002 and 2003) on 556 firm entries from the S&P 1500.[2] The most challenging aspect when constructing the datasets was to "match" the original company and the directors' primary company information. For each selected firm entry, background company information for at least two outside directors was entered successfully.

In detail, the 556 firm entries in the final sample were selected using the following procedure. First, we started with firm entries from the S&P 1500 in 2002 and 2003, after which we collected each firm's director information from "Directorships in Board Analyst" database, which included the names of the directors and their primary companies.

Second, CEO compensation and certain types of economic information for each firm were taken from "2002 and 2003 Compustat Execomp" database.

Third, to obtain CEO compensation information at the directors' primary companies, we matched the outside directors' primary companies against the "2002 and 2003 Compustat Execomp" database. Note that tiresome and time-consuming "hand matching" is required during this process because the names of the directors' primary companies contained in Board Analyst are displayed in a casual manner, whereas Compustat Execomp provides the companies' official names. Moreover, only outside directors' primary companies that are listed (or formerly listed) in the S&P 1500 index can be matched, as Compustat Execomp provides information exclusively on S&P 1500 companies.

Fourth, using the match success result, we selected firms that show at least two match successes (firms matched with at least two directors' primary companies).

Fifth, to control for the industry effect, we needed at least eight firm entries for each industry group identified using the two-digit Standard Industrial Classification (SIC) code. Thus, we eliminated firm entries with fewer than eight companies in the same industry group.

According to these criteria, 556 firm entries in total qualified for the test.[3] Larger firms are easier to match to their directors' primary companies listed in the S&P 1500. Thus, we find that most of the selected firms in the final sample are in the S&P 500. The match ratio is 2.78, indicating that each firm entry is matched to approximately three director companies.[4] The two-digit SIC code classification

[2]The empirical findings in this paper could be verified more concretely using a panel dataset that reflects extended periods.

[3]Samples with excessive CEO pay could hinder the determination of the general business-sector trend. Thus, we removed outliers reflecting more than $40 million in CEO compensation, either for each given firm or as the average CEO compensation of directors' primary companies.

[4]In detail, 301 firm entries have 2 matched director companies; 129 firm entries have 3; 88 firm entries have

shows that the sample contains 15 different industries.

B. *Variables*

The dependent variable is *CEO pay*, which is the CEO's total compensation, including their salary, bonus, restricted stock grants, stock options, and other benefits. The variable of primary interest is *director company CEO pay*, which represents the average CEO pay in the matched directors' primary companies.

Table 1 (Variable Definitions) and Table 2 (Summary Statistics) describe all of the control variables used in the test. Below, we define the key control variables and explain the correlation of these variables with the dependent variable, *CEO pay*.

The following eleven board and ownership characteristics were collected for the test: *independent directors*, *chairman CEO*, *compensation chair appointment*, *director appointment by CEO*, *busy directors*, *aged directors*, *board size*, *insider ownership*, *institutional ownership majority*, *entrenchment index*, and *CEO tenure*.

The first four variables represent the degree of board independence: *independent directors*, *chairman CEO*, *compensation chair appointment*, and *director appointment by CEO*.[5] The next three variables indicate the degree of board effectiveness: *busy directors*, *aged directors*, and *board size*.[6] Boards that are more independent and more effective are expected to be negatively correlated with the level of CEO pay (Core *et al.* 1999).

The next two variables, *insider ownership* and *institutional ownership majority*, represent the ownership structure. *Insider ownership* is the expected percentage of shares held by top management and directors. *Institutional ownership majority* indicates whether a majority of outstanding shares are held by institutions. Larger insider ownership levels result in fewer agency problems and are thus expected to be negatively correlated with the level of CEO pay (Jensen and Meckling 1976). Institutions are expected to participate more actively in corporate governance issues. Thus, majority ownership by institutional investors may prevent excessive CEO pay.

To reflect the overall entrenchment of the management team, we use the entrenchment index of Bebchuck *et al.* (2009), which is composed of four "constitutional" provisions, including a supermajority requirement for mergers and a staggered board, along with two "takeover readiness" provisions, specifically poison pills and severance agreements. Stronger entrenchment may imply a weaker governance structure and thus may result in higher CEO pay.

Finally, *CEO tenure* is expected to be negatively correlated with CEO pay. Jensen *et al.* (2004) argue that the BOD almost invariably pays "too much" for

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4; 26 firm entries have 5; 6 firm entries have 6; 4 firm entries have 7; and 2 firm entries have 8.

[5]*Compensation chair appointment* indicates whether the chairman of the compensation committee was appointed by the current CEO. *Director appointment by CEO* is the ratio of outside directors appointed by the current CEO.

[6]*Busy directors* refers to the ratio of directors who serve on more than four boards. *Aged directors* is the ratio of directors who are older than 70 years old.

TABLE 1—VARIABLE DEFINITIONS

Table 1 presents definitions of the variables used in this paper. [CE] indicates that the data source is Compustat Execomp and [BA] indicates that the data source is Board Analyst.

Variables	Definitions [Source]
Key Variables	
CEO Pay	Total compensation for the firm's CEO, including the following: salary, bonus, restricted stock granted, stock options granted. [CE]
Director Company CEO Pay	Average total compensation rewarded to CEOs of all matched companies, in which the outside directors work as executives. [collected manually, CE]
Ratio of CEO-Directors	Ratio of CEO-directors to the total number of outside directors. CEO-director is defined as an outside director who is a CEO in his/her primary company. [collected manually, BA]
Director Company Variables	
Director Company Market Cap	Average market cap of all matched companies, in which the outside directors work as executives. [collected manually, CE]
Director Company Entrenchment	Average *entrenchment index* of all matched companies, in which the outside directors work as executives. [collected manually, Lucian Bebchuk's Web]
Board and Ownership Characteristics	
Independent Directors	Ratio of the number of independent outside directors to the total number of outside directors. [BA]
Busy Directors	Ratio of the number of directors who are on more than four boards to the total number of outside directors. [BA]
Aged Directors	Ratio of the number of directors who are older than 70 years to the total number of outside directors. [BA]
Insider Ownership	Estimated percentage of shares held by top management and directors, as reported in the company's most recent proxy statement. [BA]
Chairman CEO	Indicates whether the current CEO is also the current chairman of the board. [BA]
Board Size	Total number of directors on the board. [BA]
CEO Tenure	Number of years of service of the current CEO. [BA]
Institution Ownership Majority	Indicates whether a majority of outstanding shares are held by institutions. [BA]
Compensation Chair Appointment	Indicates whether the chairman of the compensation committee was appointed by the current CEO. [collected manually]
Director Appointment by CEO	Ratio of the number of outside directors who were appointed by the current CEO to the total number of outside directors. [collected manually]
Entrenchment Index	Entrenchment index developed by Bebchuk *et al.* (2009). [Lucian Bebchuk's Web]
Economic Variables	
Market Cap	Closing price for the fiscal year multiplied by the number of common shares outstanding of the company. [CE]
Operating Income Change	Year-to-year percentage change in *operating income before depreciation*. [CE]
Operating Income Growth	Three-year least squares annual growth rate of *operating income before depreciation*. [CE]
ROE	*Net income before extraordinary items and discontinued operations* divided by *total common equity*. This quotient is then multiplied by 100. [CE]
Relative Performance	Indicates whether a company outperformed or underperformed its industry by comparing five-year returns [BA]
ROA	*Net income before extraordinary items and discontinued operations* divided by total assets. This quotient is then multiplied by 100. [CE]
Stock Market Return	Three-year total return to shareholders, including monthly reinvestments of dividends. [CE]
Stock Volatility	Standard deviation volatility calculated over 60 months. [CE]
SP Index Yearly Average	Yearly average of the S&P 500 index calculated using the average of 12 monthly averages. [Standard and Poor's]
Two-digit SIC Code	The first two digits from the four-digit Standard Industrial Classification (SIC) Code. [CE]

TABLE 2—SUMMARY STATISTICS

Table 2 presents summary statistics pertaining to the variables in the constructed dataset, which is a two-year panel dataset composed of firms with a fiscal year ending between June of 2002 and May of 2004. We provide only pooled statistics.

Variable	Obs.	Mean	Std. Dev.	Min.	Max.	Unit
Key Variables						
CEO Pay	547	6163.7	5790.9	467.6	36946.1	Thousands of dollars
CEO Pay (Log)	547	8.3	0.9	6.1	10.5	
Director Company CEO Pay	555	7419.1	5736.8	549.1	38612.1	Thousands of dollars
Director Company CEO Pay (Log)	555	8.6	0.8	6.3	10.6	
Ratio of CEO-Directors	537	36.2	16.4	6.3	100.0	Percentage
Director Company Variables						
Director Company Market Cap	556	18,820.4	24,747.7	49	15,4579	Millions of dollars
Director Company Entrenchment	533	2.6	1.0	0	6	Number
Board and Ownership Characteristics						
Independent Directors	537	77.0	13.5	25	94	Percentage
Busy Directors	537	10.8	12.1	0	58	Percentage
Aged Directors	537	5.7	10.6	0	100	Percentage
Insider Ownership	519	13.1	15.1	0	83	Percentage
Chairman CEO	537	0.7	0.4	0	1	Dummy
Board Size	537	10.8	2.8	5	23	Number
CEO Tenure	537	5.4	5.3	0	41	Number
Institution Ownership Majority	519	0.8	0.4	0	1	Dummy
Compensation Chair Appointment	531	0.2	0.4	0	1	Dummy
Director Appointment by CEO	514	36.0	30.0	0	100	Percentage
Entrenchment Index	518	2.8	1.2	0	6	Number
Economic Variables						
Market Cap	552	13,909.3	31,024.3	91	271,002	Millions of dollars
Operating Income Change	538	13.9	377.2	–5,899	4,655	Percentage
Operating Income Growth	523	4.5	29.2	–58	390	Percentage
ROE	543	13.2	41.8	–314	454	Percentage
Relative Performance	518	6.5	6.5	–12	20	Percentage
ROA	554	3.0	14.0	–207	60	Percentage
Stock Market Return	534	2.3	20.1	–76	75	Percentage
Stock Volatility	553	0.4	0.2	0	2	Number
SP Index Yearly Average	556	978.2	10.3	968	989	Number

newly appointed CEOs.[7]

In addition to the board and ownership characteristics, the following eight performance variables were devised for the test: *market cap*, *operating income change*, *operating income growth*, *ROE*, *relative performance*, *ROA*, *stock market return (three-year)*, and *stock volatility*. Among them, *market cap* is used to reflect the size of the company. *Operating income change*, which is the year-to-year percentage change in operating income, represents short-term performance. *Operating income growth*, which is the three-year growth rate of operating income, represents long-term performance. *ROE (ROA)*, which is net income divided by total common equity (total assets), represents the profitability of the firm's operations. *Relative performance* indicates the degree to which a company outperformed or underperformed its industry in the stock market over the five previous years. These performance variables are expected to be positively correlated with CEO pay.

[7]However, CEO tenure may also be a proxy for the CEO's job experience. If this is the case, tenure can be positively correlated with the wage variable, *CEO pay*.

Finally, to represent the macroeconomic condition in each year, the variable of the *SP index yearly average* is constructed, representing the average of 12 monthly averages of the S&P 500 index. To reflect the industry effect (Finkelstein and Hambrick 1995),[8] we also include the two-digit SIC code, consisting of the first two digits of the four-digit SIC code.[9]

III. Methodologies

In this section, we present the methodologies used for the panel data analysis. Primarily, we conduct the Breusch-Pagan Lagrange Multiplier (LM) test for random effects (Breusch and Pagan 1980). The null hypothesis of the test is that the variance of the unobserved firm-specific fixed effects is zero. Because the null is rejected in the test for various possible specifications, we confirm that a simple pooled OLS estimator is not efficient. (For instance, refer to the LM test result at the bottom of Table 3 for regression 3.) Either the random (or the fixed) effects models or a pooled OLS with cluster robust standard errors would be more appropriate.

A. *Regression Specification for the Hypothesis*

First, we search for suitable regression specifications to examine the relationship between CEO compensation in a given company and average CEO compensation in the directors' primary companies, presenting the empirical test outcomes for a pooled regression with cluster robust standard errors (Rogers 1993). The test outcomes are then compared with their alternatives — the random effects (RE) and the fixed effects (FE) models. We start with the following regression specification:

$$CEOpay_{it} = \alpha DirCEOpay_{it} + \gamma B_{it} + \delta E_{it} + \nu DI_i + \theta S_t + \upsilon DY_t + c_i + u_{it}. \quad (1)$$

In this regression, $CEOpay_{it}$ is the level of CEO compensation in the given company (*CEO pay*), $DirCEOpay_{it}$ is the level of average CEO compensation in the directors' companies (*director company CEO pay*), B_{it} is the set of the eleven board and ownership characteristic variables, and E_{it} denotes the set of the eight performance variables, including *market cap* and *ROA*. DI_i is an industry dummy obtained from the two-digit SIC code, which is used to control for the industry fixed effect. S_t is the S&P 500 index yearly average, which is included to reflect the macroeconomic conditions for each year.

The residuals for a given year may be correlated across different firms according to a concept known as "time effects" or "spatial correlation." To handle time effects, year dummies (DY_t) are included in the regression. On the other hand, the residuals for a given firm may be correlated across years, a concept referred to as

[8]Finkelstein and Hambrick (1995) argue that CEOs are frequently compensated in relation to CEOs in the same industry, with the empirical finding that a 1% increase in CEO pay within an industry is associated with a 1.37% increase in the pay of the focal CEOs in their sample.

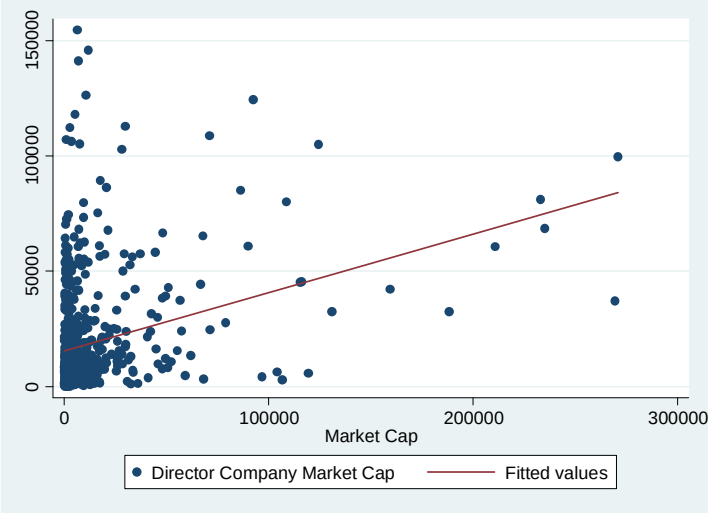
[9]The two-digit SIC code represents 83 industry categories in the United States.

TABLE 3—EFFECTS OF DIRECTOR COMPANY CEO PAY

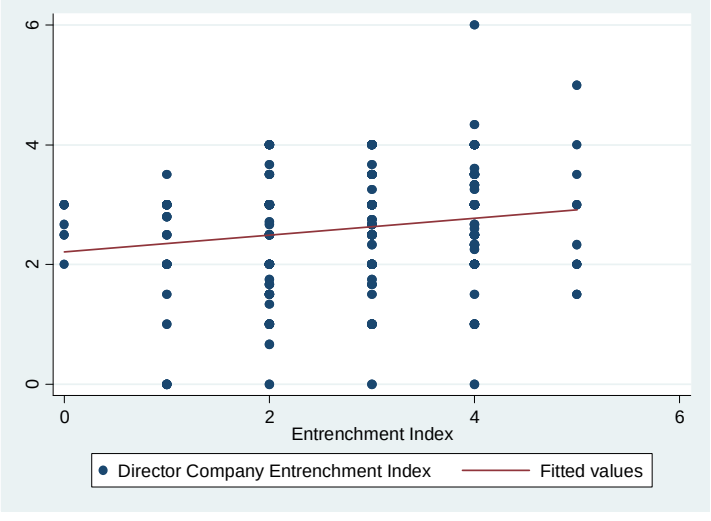
Table 3 presents the test results of the pooled regression with cluster robust standard errors (1–2), the random effects model (3), and the fixed effects model (4). The dependent variable is *CEO pay*. The key variable of interest is *Director Company CEO Pay*. The Breusch and Pagan LM test and Hausman's specification tests are reported at the bottom. In each regression, the following economic variables are controlled, together with *Market Cap* and *ROA*: *Operating Income Change*, *Operating Income Growth*, *ROE*, *Relative Performance*, *Stock Market Return*, *Stock Volatility*, and *SP Index Yearly Average.* Standard errors are reported in parenthesis. *, **, and *** indicate significance at the 10%, 5%, and 1% levels respectively.

Dependent Variable: CEO Pay	(1) Pooled OLS (Cluster)	(2) Pooled OLS (Cluster)	(3) Random Effects	(4) Fixed Effects
Director Company CEO Pay	**0.12*** **	**0.14*** **	**0.17*** **	**0.22*** **
	(0.0)	**(0.04)**	**(0.05)**	**(0.07)**
Director Company Market Cap		0.003	0.002	0.020
		(0.014)	(0.013)	(0.040)
Director Company Entrenchment		378.6	431.5	.
		(292.0)	(299.5)	.
Independent Directors	–25.0	–17.5	–10.2	–6.2
	(22.6)	(20.8)	(19.2)	(28.8)
Busy Directors	36.2*	40.9**	31.9	23.8
	(19.7)	(19.8)	(19.6)	(33.1)
Aged Directors	19.8	29.9	–12.8	–94.9***
	(72.2)	(73.2)	(22.7)	(35.0)
Insider Ownership	–30.6	–32.5*	–25.2	–16.7
	(18.8)	(19.5)	(17.8)	(21.5)
Chairman CEO	528.6	489.3	203.2	76.7
	(430.8)	(420.3)	(394.0)	(465.8)
Board Size	65.5	78.1	144	189.7
	(102.3)	(106.0)	(109.6)	(195.8)
CEO Tenure	17.4	39.3	80.7	347.1
	(77.2)	(76.8)	(87.6)	(212.0)
Institution Ownership Majority	453.4	253.1	525.4	893.4
	(698.1)	(712.7)	(597.1)	(819.9)
Compensation Chair Appointment	–2,428.2***	–2,261.4***	–1,104.2	935.6
	(797.7)	(790.1)	(699.5)	(988.5)
Director Appointment by CEO	13.1	13.2	0.33	–36.8
	(12.2)	(12.8)	(13.6)	(24.5)
Entrenchment Index	87.3	129.6	133.0	.
	(235.9)	(249.5)	(254.5)	.
Market Cap	0.082***	0.083***	0.078***	0.0087
	(0.011)	(0.012)	(0.009)	(0.034)
ROA	–4.44	20.1	102.1*	172.7*
	(69.5)	(67.1)	(58.3)	(89.6)
Economics Variables Controlled	Yes	Yes	Yes	Yes
Industry Dummy	Yes	Yes	Yes	Yes
Year Dummy	Yes	Yes	Yes	Yes
Observation	414	401	401	401
R-squared	0.52	0.54		
Breusch and Pagan LM Test			0.00	
Hausman p-value (FE/RE)				0.0001
Hausman p-value (FE/Pooled OLS)				0.0000

"firm-specific effects" or "temporal correlation." This unobserved firm-specific effect is denoted by c_i in the given specification. Finally, u_{it} represents the idiosyncratic error.



PANEL 1



PANEL 2

FIGURE 1—CHARACTERISTICS OF DIRECTORS' PRIMARY COMPANIES

Panel 1 presents a scatterplot of *director company market cap* versus *market cap* and its fitted line. Panel 2 presents a scatterplot of *director company entrenchment* versus *entrenchment index* and its fitted line.

As a primary estimation, we use a pooled OLS regression with cluster robust standard errors. Regression 1 in Table 3 presents the test result, which shows a significant coefficient for the variable of interest, $DirCEOpay_{it}$, where a $1.00 increase in CEO pay at the directors' primary companies is associated with an approximate $0.12 increase in CEO pay at the given firm.

The first concern with this initial specification is that the coefficient of the variable of interest $DirCEOpay_{it}$ may not capture the direct relationship between the two CEO pay variables. This concern arises out of the director selection procedure; large firms tend to hire outside directors from other large firms, and the size of director firms is strongly correlated with the level of CEO pay at the director firms. Panel 1 in Figure 1 indicates the potential seriousness of this concern: larger firms are more likely to hire directors from other larger firms. Thus, even without the strong direct effect of *DirCEOpay* on *CEOpay*, the coefficient for the variable of interest, α, could be significant in the test results. For example, *DirCEOpay* could simply be a proxy for the size of the directors' companies. To control for this possible bias, we include the size variable of director firms, *director company market cap*, as a control variable, representing the average market capitalization of all of the matched directors' companies.

Similarly, the coefficient for the variable of interest $DirCEOpay_{it}$ may reflect the positive correlation between the entrenchment in the directors' primary companies and that in the given company. Panel 2 in Figure 1 shows the potential seriousness of this concern: more strongly entrenched firms tend to hire outside directors from other entrenched firms. To control for this bias, we also include the entrenchment variable of director firms *director company entrenchment* as a control variable, representing the average entrenchment index of all of the matched directors' companies.

Thus, we arrive at the following adjusted specification, in which $DirComp_{it}$ represents a set of two additional control variables, *director company market cap* and *director company entrenchment*:

$$CEOpay_{it} = \alpha DirCEOpay_{it} + \beta DirComp_{it} + \gamma B_{it} + \delta E_{it} + \nu DI_i + \theta S_t + \upsilon DY_t + c_i + u_{it}. \tag{2}$$

The pooled OLS regression with the adjusted specification is displayed in the results for regression 2 in Table 3. This reconfirms the significant coefficient α for the variable of interest, $DirCEOpay_{it}$, which is estimated to be 0.14.

B. *Alternative Econometric Models*

We present the test results from other alternatives — the random and the fixed effects models — in regressions 3 and 4, respectively. Both regressions confirm the significant coefficients for the variable of interest, $DirCEOpay_{it}$, which are estimated to be 0.17 (RE) and 0.22 (FE).

The fixed effects model generates a consistent estimate even when the firm-specific effects are correlated with any of the independent variables.[10] Consequently, we conduct the Hausman (1978) specification test to compare the fixed effects model and several alternative estimators, shown at the bottom of Table

[10]According to Petersen's simulations (2009), the pooled regression with cluster robust standard error generates a consistent estimate when the firm-specific fixed effects vary over time. However, it does not make sense to assume time-varying fixed effects in the given two-period sample.

3, for regression 4, in which the null hypothesis is that the firm-specific effect (c_i) is not correlated with other regressors. Because the null is rejected in each test, both the random effects model and the pooled OLS model are biased.

Therefore, the fixed effects estimator used in regression 4, which is identical to the unbiased first-difference (FD) estimator owing to the two periods in the given sample, solely remains consistent. According to the fixed effects estimation, we conclude that a reliable estimate of the coefficient for the variable of interest, α, is 0.22.

IV. Empirical Analysis

In this section, we summarize the test results from the previously described panel data analysis. Table 3 indicates the test results for the given hypothesis using the pooled OLS (with cluster robust standard errors) and its alternative RE/FE models. In all of the regressions, the coefficient for *director company CEO pay* is positive and significant at the 1% level. However, the Hausman specification test demonstrates the unacceptability of the pooled OLS and random effects models against the fixed effects model. The unbiased FE(FD) estimation in regression 4 suggests that a $1.00 increase in CEO pay at the outside directors' primary companies results in an approximate $0.22 increase in CEO pay.

As a robustness check, we also report the elasticity estimates in Table 4 using the logarithm transformations of *CEO pay* and *director company CEO pay*. The three regressions present the elasticity of the two compensation variables in the pooled OLS, RE, and FE models. (The Breusch-Pagan LM test for random effects and Hausman's specification tests are reported at the bottom.)

The elasticities estimated using the three different regressions are found to be very close to each other, at approximately 0.12~0.13. According to the unbiased FE(FD) estimate, it is interpreted that a 1% increase in CEO pay at the outside directors' primary companies is associated with a 0.13% increase in CEO pay at a given firm. Except for the FE model, the coefficients of the variable of interest, *director company CEO pay (Log)*, are positive and significant at the 5% level. The coefficient of the FE model is significant at the 10% level when using a one-sided test, which is acceptable in the given analysis because we conjecture that the negative association between the two compensation variables is unreasonable.

TABLE 4—ROBUSTNESS CHECK (ELASTICITY)

Table 4 presents the test results of the pooled regressions with cluster robust standard errors (1), the random effects models (2), and the fixed effects models (3). The dependent variable is *CEO pay (Log)*. The key variable of interest is *Director Company CEO Pay (Log).* The Breusch and Pagan LM test and Hausman's specification tests are reported at the bottom. In each regression, the following economic variables are controlled together with *Market Cap* and *ROA*: *Operating Income Change*, *Operating Income Growth*, *ROE*, *Relative Performance*, *Stock Market Return*, *Stock Volatility*, and *SP Index Yearly Average.* Standard errors are reported in parenthesis. *, **, and *** indicate significance at the 10%, 5%, and 1% levels respectively. [*] indicates significance at the 10% level according to a one-sided test.

Dependent Variable: CEO Pay (Log)	(1) Pooled OLS (Cluster)	(2) Random Effects	(3) Fixed Effects
Director Company CEO Pay (Log)	**0.12****	**0.12****	**0.13[*]**
	(0.059)	**(0.057)**	**(0.088)**
Director Company Market Cap	1.2E–06	1.2E–06	–3.5E–09
	(0.000)	(0.000)	(0.000)
Director Company Entrenchment	0.056	0.059	.
	(0.056)	(0.054)	.
Independent Directors	–0.005	–0.004	–0.003
	(0.004)	(0.003)	(0.005)
Busy Directors	0.008**	0.005	0.000
	(0.003)	(0.003)	(0.005)
Aged Directors	–0.001	–0.005	–0.006
	(0.009)	(0.004)	(0.006)
Insider Ownership	–0.007**	–0.004	–0.002
	(0.004)	(0.003)	(0.003)
Chairman CEO	0.020	0.019	0.027
	(0.075)	(0.063)	(0.073)
Board Size	0.055***	0.042**	0.0031
	(0.019)	(0.019)	(0.031)
CEO Tenure	–0.002	0.000	0.006
	(0.018)	(0.015)	(0.033)
Institution Ownership Majority	0.180	0.220**	0.210
	(0.140)	(0.099)	(0.130)
Compensation Chair Appointment	–0.390***	–0.150	0.110
	(0.150)	(0.120)	(0.160)
Director Appointment by CEO	0.004	0.002	0.001
	(0.003)	(0.002)	(0.004)
Entrenchment Index	0.022	0.033	.
	(0.046)	(0.046)	.
Market Cap	8.1E–06***	8.9E–06***	3.4E–06
	(0.000)	(0.000)	(0.000)
ROA	0.005	0.017*	0.029**
	(0.011)	(0.010)	(0.014)
Economics Variables Controlled	Yes	Yes	Yes
Industry Dummy	Yes	Yes	Yes
Year Dummy	Yes	Yes	Yes
Observation	401	401	401
R-squared	0.46		
Breusch and Pagan LM Test		0.00	
Hausman p-value (FE/RE)			0.22
Hausman p-value (FE/Pooled OLS)			0.00

V. Discussion

The positive association between CEO pay at outside directors' primary companies and CEO pay at a given company may stem from either 1) the direct positive association between the size of the director firms and that of the given firm, or 2) the direct positive association between the degree of entrenchment in the director firms and that of the given firm; larger firms tend to hire outside directors from other large firms and more entrenched firms tend to hire outside directors from other entrenched firms. However, as discussed in Section III, the positive effect (observed in Tables 3 and 4) of the change in CEO pay at outside directors' primary companies on the change in CEO pay at a given company is not easily attributable to the direct positive associations between the director firms and the given firm because we include both *director company market cap* and *director company entrenchment* as control variables in the regression analysis.

On the other hand, some may argue that the observed positive effect is attributable to cross-directorship, in which outside directors favor CEOs who serve as outside directors at their own primary companies (e.g., Hallock 1997). However, it is noteworthy that while reciprocal CEO interlocks exist occasionally, they are not very common practice in the business world as it exists today.[11] In our limited sample of 556 firm entries, we observe only several interlocking relationships. Indeed, we find that removing those few entries from the sample makes little difference in the test outcomes. Therefore, without hesitation, we conclude that cross-directorship is not a major driving force behind the observed positive effect in the above analysis.

Among other possible sources that may generate the observed effect, we pay close attention to the following two phenomena: 1) the psychological phenomenon known as the *anchoring effect*, and 2) the societal phenomenon termed the *inner circle*. Tversky and Kahneman (1974) suggest that individuals make estimates by starting from an "initial value" (anchor) that is adjusted to yield the final answer. Because adjustments are typically insufficient, different starting points yield different estimates that are biased toward the initial value. According to this theory, CEO compensation in outside directors' primary companies may serve as a firsthand reference point for directors' decisions on CEO pay. Once their primary companies pay higher CEO compensation, they are more likely to choose a higher value for CEO pay in a given company and to accept a CEO's request to increase his or her salary. Thus, all else being equal, CEO compensation in a given firm may reflect the level of CEO compensation at the directors' primary companies. This anchoring effect may be a primary source of the observed sensitivity between the change in CEO pay at the directors' primary companies and the change in CEO pay at a given company.

Top corporate executives tend to be a relatively homogeneous, cohesive collection of individuals, which Useem (1984) calls the *inner circle*.[12] Outside directors who are top executives at other firms find it difficult to be fully objective

[11]For instance, even in Hallock's (1997) study, only 8% of CEOs are reciprocally interlocked.

[12]The term *inner circle* is defined as all corporate executives who serve on the board of directors of two or more big corporations.

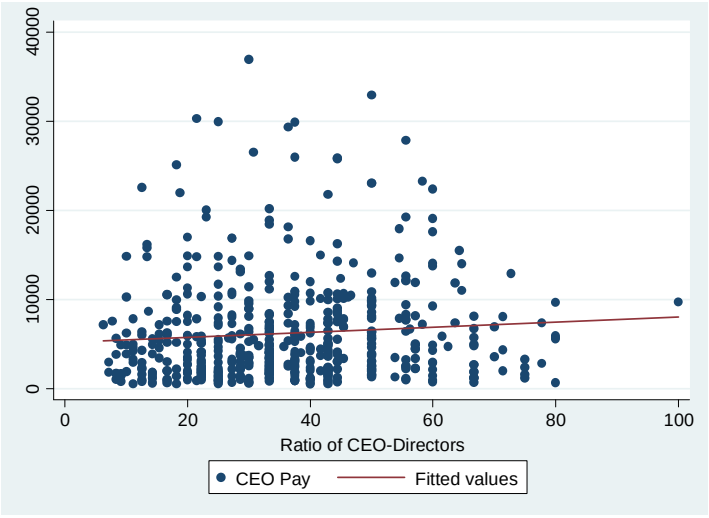


FIGURE 2. CEO PAY AND THE RATIO OF CEO-DIRECTORS

The figure presents a scatterplot of *ratio of CEO-directors* versus *CEO pay* with its fitted line.

about issues related to executive compensation. For instance, the pay consultant Watson Wyatt conducted a survey of two separate groups — directors and institutional investors — on the subject of the U.S. executive pay model.[13] The survey reveals that the two groups have very different views of the current pay system; 65 percent of directors and only 22 percent of institutional investors believe that the stock-based pay model in the U.S. has contributed to superior corporate performance; 90 percent of institutional investors think that executives at most companies are overpaid, compared to only 60 percent of directors. This homogeneity of the top executives and the concurrent executive employment of outside directors may contribute to the observed positive association between CEO pay levels as far as they can stimulate implicit collusion among executives. The more cohesive the top executives (outside directors) are in a specific industrial sector (or in a specific regional market), the more closely linked CEO pay levels of the relevant companies are.

As supplementary evidence to support this *inner circle* argument, we checked whether a higher number of outside directors who are active CEOs in other firms results in the board granting a higher level of CEO compensation. A higher number of CEO-directors on the board may help the CEO in a given firm propose suggestions that are favorable to his interests and position; CEO-directors are more likely to appreciate the CEO's efforts and contributions and, thus, willingly increase his salary. Let the variable *ratio of CEO-directors* represent the ratio of CEO-directors to the total number of outside directors on the board (c.f., Tables 1 and 2). At first glance, the scatterplot and the fitted line in Figure 2 imply a possible positive association between *CEO pay* and *ratio of CEO-directors*.

[13]In the survey, 55 institutions managing $800 billion in assets participated, along with 50 directors. The survey was conducted with institutional investors in 2005 and with directors in 2006 (Watson Wyatt, 2006).

TABLE 5—SUPPLEMENTARY ANALYSIS (RATIO OF CEO-DIRECTORS)

Table 5 presents the test results of the pooled regression with cluster robust standard errors (1), random effects model (2), and fixed effects model (3). The dependent variable is *CEO pay*. The key variable of interest is the *ratio of CEO-directors.* The Breusch and Pagan LM test and Hausman's specification tests are reported at the bottom. In each regression, the following economic variables are controlled, together with *Market Cap* and *ROA*: *Operating Income Change*, *Operating Income Growth*, *ROE*, *Relative Performance*, *Stock Market Return*, *Stock Volatility*, and *SP Index Yearly Average.* Standard errors are reported in parenthesis. *, **, and *** indicate significance at the 10%, 5%, and 1% levels respectively.

Dependent Variable: CEO Pay	(1) Pooled OLS (Cluster)	(2) Random Effects	(3) Fixed Effects
Ratio of CEO-Directors	**43.6*** **	**33.3*** **	**31.5* **
	(15.9)	**(12.9)**	**(17.9)**
Independent Directors	–25.0	–17.8	–3.0
	(21.1)	(19.3)	(29.6)
Busy Directors	33.5*	29.8	38.5
	(20.2)	(19.9)	(34.2)
Aged Directors	29.9	–13.3	–95.2***
	(69.0)	(23.2)	(36.4)
Insider Ownership	–27.6	–22.3	–14.1
	(18.7)	(18.0)	(21.6)
Chairman CEO	528.0	128.9	–200.4
	(429.6)	(398.5)	(467.6)
Board Size	125.6	200.1*	291.6
	(102.7)	(112.5)	(200.3)
CEO Tenure	16.0	44.1	274.5
	(79.6)	(83.8)	(198.0)
Institution Ownership Majority	453.8	689.8	929.0
	(689.7)	(610.6)	(835.5)
Compensation Chair Appointment	–2350.5***	–1318.0*	687.2
	(782.9)	(695.6)	(964.1)
Director Appointment by CEO	10.5	2.7	–22.9
	(11.9)	(13.2)	(22.3)
Entrenchment Index	–6.6	–2.8	.
	(225.5)	(246.4)	.
Market Cap	0.087***	0.082***	0.0002
	(0.011)	(0.009)	(0.034)
ROA	–39.0	59.9	166.3*
	(71.3)	(58.3)	(89.9)
Economics Variables Controlled	Yes	Yes	Yes
Industry Dummy	Yes	Yes	Yes
Year Dummy	Yes	Yes	Yes
Observation	415	415	415
R-squared	0.53		
Breusch and Pagan LM Test		0.00	
Hausman p-value (FE/RE)			0.0001
Hausman p-value (FE/Pooled OLS)			0.0000

Table 5 reports the estimation results of the pooled OLS (with cluster robust standard errors), the random effects model, and the fixed effects model. According to the fixed effects estimation, the unbiased estimate of the coefficient for the variable of interest is 31.5, indicating that an increase of one percentage point in the ratio of outside directors who are active CEOs in other firms results in an

approximate $31,500 increase in CEO pay.[14]

In the given sample, the average number of outside directors on the board is 9.3, indicating that one additional CEO-director on the board is equivalent to a 10.8% increase in the *ratio of CEO-directors*. Therefore, the replacement of a non-CEO outside director with a CEO-director may result in an approximate $0.34 million increase in CEO pay, a significant amount considering that the average CEO pay in the sample is approximately $6 million. This observation supports the argument that CEO-directors tend to view the BOD (either consciously or subconsciously) through the eyes of a CEO.

There may be other valid factors influencing the observed strong link between CEO pay at outside directors' primary companies and CEO pay at a given company. Further explorations are left for future research.[15]

VI. Conclusion

Since the early 1990s, many articles have been published on executive pay (Boyd *et al.* 2012). Some academic publications have focused on the optimal compensation contract and the pay–performance relationship when debating stock-based pay and soaring executive pay (e.g., Jensen and Murphy 1990a, 1990b; Hall and Liebman 1998; Hall and Murphy 2003; Aggarwal and Samwick 1999; Bebchuk and Fried 2003). Others have focused on the relationship between board characteristics and CEO compensation, concentrating on corporate governance weaknesses and BOD independence (e.g., Core *et al.* 1999; Jensen 1993). The primary concern in the literature has been connected to the agency problem caused by the management team, or what is termed a "captured" BOD.

However, the literature has not put much weight on the agency problem that may be caused purely by the BOD, including a concurrent employment issue of board members which may seriously affect their decision processes as related to the level of CEO compensation. As a complement to the "social network" literature that emphasizes the impact of cross-directorship on the level of CEO compensation, the empirical work in this paper provides a new approach to the issue of concurrent employment on a board.

The unbiased fixed effects estimation confirms a strong link between CEO pay at outside directors' primary companies and CEO pay in a given company, even after controlling for both the size (and the entrenchment) associations between the directors' primary companies and the given firm. Furthermore, we find that the

[14]We utilize the Breusch-Pagan LM test for random effects and present the test result at the bottom of Table 5 for regression 2. Because the null is rejected, a pooled OLS estimator is not efficient. Subsequently, we conduct the Hausman tests to compare the fixed and random effects models and to compare the fixed effects and pooled OLS models. The test results at the bottom of Table 5 indicate that the nulls are rejected. Thus, both the random effects and the pooled OLS estimators are biased, while only the fixed effects estimator remains unbiased and consistent.

[15]There may be a positive association between the CEO pay levels due to a CEO-director ability effect: When a talented CEO who is paid higher in firm A serves as an outside director in firm B, the CEO pay in firm B can be higher as well owing to the talented outside director's contribution to the productivity of firm B. However, the observed strong link in this paper cannot be attributable to this CEO-director ability effect because we report the test outcomes from the first-difference (FD) panel estimations. That is, the ability of CEO-director does not vary across the years and thus this omitted variable does not affect the estimation outcomes.

number of cross-directorship relationships is very small in the sample here; thus, the "social network" argument does not support the observed strong link. Instead, we do not rule out the possibility that the link results from unintended psychological bias, often known as the *anchoring effect*. It is also plausible that the pay inefficiency originates from the homogeneous and cohesive collection of individuals (the *inner circle*) who serve as top executives as well as outside directors.

The degrees of psychological bias and pay inefficiency will be greater when outside directors are more negligent or less considerate. Thus, these factors cannot be excuses for the violation of the directors' duty to care. Moreover, this may introduce improper incentives for CEOs, as they may be tempted to recruit outside directors who serve for companies paying greater CEO compensation or from their own *inner circle*.

This additional agency problem — caused by the BOD and not by the management team — will be difficult to control as long as concurrent executive employment on boards continues to exist. As suggested by some governance experts, including Jensen *et al.* (2004) and Brenner and Schwalbach (2009), this conflict may be alleviated through specific legal rules that make boards more accountable to shareholders, such as strengthening the procedural rights of shareholders who vote to elect directors.

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Misallocation and Manufacturing TFP in Korea

By JIYOON OH*

This paper analyzes the effects of allocative efficiency on productivity in the manufacturing sector of Korea following Hsieh and Klenow (2009). The results of this research indicate that the overall allocative efficiency declined from 1990 to 2012. Using the method of Oberfield (2013), which allows inter-industry resource movement as well as intra-industry reallocation, we confirm that intensified misallocation generally results from intra-industry allocative inefficiency. The potential loss from instances of worsening misallocation is estimated to be approximately 0.6% points for each year, which is considerable in terms of the overall TFP. In terms of the firm size distribution, initially large establishments are more likely to expand if distortions are removed in most countries. One notable feature in Korea is that this pattern is pronounced. This implies that subsidies to unproductive small-sized establishments are heavily implemented.

Key Word: Misallocation, Productivity Differences,
Misallocation by Establishment Size
JEL Code: O11, O47, O53

I. Introduction

Productivity differences have received attention as a principal source of the gap between rich and poor countries. There are two main directions of research on productivity. Traditionally, many studies compare the productivity growth of a representative (average) firm for each country. Growth accounting is a typical strategy used in this strand. Another approach focuses on the heterogeneity of firms or industries and pays attention to allocative efficiency. The distribution of firm productivity within the same industry is known to be highly dispersed; thus, resource redistributions toward more productive cases generates higher TFP growth throughout the economy, even if the productivity level of each firm does not change.

Emphasizing allocative efficiency has been a growing trend and is currently popular. Hsieh and Klenow (2009, HK henceforth) developed an empirical

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methodology for measuring TFP losses due to instances of resource misallocation. They compared the extent of misallocations for the U.S., China, and India using plant-level data from the manufacturing sector of each country. In the absence of misallocation, the U.S. can achieve an increase of output of approximately 30-43%, while China and India can enjoy rates of 87-115% and 100-128%, respectively. What they refer to as "misallocation" roughly alludes to inefficiency, which is measured as the ratio of current output relative to the potential output resulting from the removal of exogenous distortions. This potential output is hypothetically achieved in the absence of differences apart from productivity differences. Each firm/plant differs in terms of their technology or efficiency level when creating their products. Only in terms of productivity should the amount of factor inputs be distributed according to the relative order of firm productivity. In reality, this is not the case. All other factors affecting allocation except for productivity are regarded as distortions.

The HK methodology has several shortcomings; one of them is that it investigates allocative efficiency only within industries. In the HK method, the total amounts of capital and labor in each industry are fixed and do not very between industries. This weakness shuts down the flexibility of factor movement during industry-specific shocks. Oberfield (2013) developed a measure of allocative efficiency along the lines of Hsieh and Klenow (2009), allowing resource movement between industries. He found an interesting phenomenon during the 1982 crisis in the Chilean manufacturing sector. A severe contraction in output during this period is mostly accounted for by a falling Solow residual, and a decline in between-industry allocative efficiency contributes to the drop in TFP while the within-industry efficiency remains constant.

Following the HK methodology, observations from many countries were also reported. Using plant data from the French manufacturing sector, Bellone and Mallen-Pisano (2011) argued that allocative efficiency in France is comparable to that in the U.S. Particularly, Hosono and Takizawa (2012) reported that misallocation in the manufacturing sector in Japan has been deepening since the 1980s. The unique point in their paper is that it determined the time series of misallocation over time as well as the overall level of misallocation compared to those in other countries. For Japan, they showed that the improvement in the TFP would be 47% if misallocation disappears. Bartelsman *et al.* (2013), who developed their own criterion to quantify misallocation, showed that approximately 15% of TFP losses can be accounted for by distortions in resource allocation in the transition economies of Central and Eastern Europe.

In this paper, we quantitatively measure the extent of misallocation in the Korean manufacturing sector and compare it with those in other countries following the same methodology used by Hsieh and Klenow (2009). Although Midrigan and Xu (2009) have analyzed capital misallocation using the same data, their derivation of capital frictions is based on the assumption that other inputs such as labor and intermediate inputs are allocated efficiently without friction. The HK calculation considers labor misallocation as well as capital misallocation; thus, the efficient output differs from that of Midrigan and Xu (2009).

This study also shows how efficiency in this sector has evolved over the past 20 years. The time-series trend in allocative efficiency is another contribution of our

paper compared to the data results by Midrigan and Xu (2009), who mainly pool all year-plant observations while focusing on cross-sectional variations of capital productivity. With regard to robustness checks, we will investigate our results through the lens of the methodology developed by Oberfield (2013).

The main results here are as follows. Korea's average allocative efficiency is found to be lower than that of the U.S., higher than that of China, and similar to that of Japan. Allocative efficiency in the U.S. was 0.73 (highest) while that of China was 0.50 (lowest).[1] The rates for Korea and Japan are similar at 0.65 and 0.68, respectively. Meanwhile, the time series shows that Korea's allocative efficiency in the manufacturing has been trending downward since the 1990s. This downward trend in Korea is similar to that of Japan when it began to experience slowing growth in the 1990s. A downward trend consistently appears after allowing for inter-industry resource allocation.

The paper proceeds as follows. Section II describes the HK and Oberfield (2013) methodologies heavily relied on by the present paper. In Section III, we report the empirical results. Section IV briefly summarizes possible reasons for misallocation in the Korean manufacturing sector. Section V presents our concluding remarks.

II. Previous Methodology

In this section, we illustrate the backbone frames of Hsieh and Klenow (2009) and Oberfield (2013) which will be used for the empirical analysis here.

Hsieh and Klenow (2009) provide a quantitative methodology pertaining to the potential effect of resource misallocation on the aggregate TFP. In an economy with heterogeneous production units, aggregate TFP depends not only on the TFPs of individual production units but also on how the inputs are allocated across these production units. If firm productivity is the single factor of heterogeneity, it is natural for input resources such as capital and labor to be distributed according to productivity differences. However, if there are firm-level distortions which are not related to productivity, resource allocation may damage aggregate productivity via a process known as misallocation. In Hsieh and Klenow (2009), the reasons for distortions are not primarily studied, but the types of distortions noted are taxes and subsidies, financial friction, trade restrictions, capital rationing, a host of regulations associated with firm-size-dependent policies, industrial policies, and entry barriers. In other words, distortions are the reduced forms of all possible sources which generate resource misallocation other than productivity.

How can we measure distortions (wedge) in real data? Can we identify differences in firm productivity and firm distortions? The answers can be found based on a standard model of monopolistic competition with heterogeneous productivity. With the Cobb-Douglas production function and constant elasticity of substitution between differentiated goods, firm markup is fixed and equal to all firms. Thus, revenue productivity (the product of physical productivity and a firm's output price) should be equal across firms in the absence of distortion. To the

[1]This figure represents the ratio of current output relative to the potential output in the absence of misallocation.

extent revenue productivity differs across firms, a measure of firm-level distortion can be recovered. Although industry variance with regard to physical productivity exists, industry variance in revenue productivity vanishes in the absence of distortion.

The details of Hsieh and Klenow (2009) are presented below. Goods markets consist of final goods and industry goods. A final good (Y) is produced by a representative firm in a perfectly competitive final output market. Industry goods (Y_s) are aggregated by way of Cobb-Douglas production.

(1) $$Y = \prod_{s=1}^{S} Y_s^{\theta}$$

$$\text{where } \sum_{s=1}^{S} \theta_s = 1$$

Industry output Y_s is a CES aggregate of M_s differentiated products.

(2) $$Y_s = \left(\sum_{i=1}^{M} Y_{si}^{\frac{\sigma-1}{\sigma}}\right)^{\frac{\sigma}{\sigma-1}}$$

Because individual products are produced in a monopolistic competition market, price elasticity is always $\sigma(>1)$ regardless of the production volume, and the markup is $\sigma/(\sigma-1)$.

Firm production is determined with the Cobb-Douglas function with productivity A_{si}, capital K_{si} and labor L_{si}. Firm productivity A_{si} (physical productivity: TFPQ) differs across firms. Capital elasticity α_s, determining the income shares of capital and labor, is equal across firms within an industry but may not be equal across firms between industries.

(3) $$Y_{si} = A_{si} K_{si}^{\alpha_s} L_{si}^{1-\alpha_s}$$

A firm pursues profit maximization, as expressed by (4). $P_{si}Y_{si}$ denotes revenue in this case. w and R are the wage rate and capital rental rate, which are identical for all firms. Output distortion τ_{Ysi} and capital distortion τ_{Ksi} are unique in this setup. These types of taxes represent the reduced forms of distortion faced by firms. τ_{Ysi} alters the marginal productivity of capital and labor at the same proportion, but τ_{Ksi} interferes with the relative marginal productivity of capital over labor.

(4) $$\pi_{si} = (1-\tau_{Ysi})P_{si}Y_{si} - \omega L_{si} - (1+\tau_{Ksi})RK_{si}$$

From the perspective of first-order conditions, τ_{Ysi} and τ_{Ksi} can be viewed from several perspectives. As expressed by (5), we infer the presence of capital

distortion τ_{Ksi} when the ratio of labor compensation to the capital stock differs relative to what one would expect from the degree of output elasticity with respect to capital and labor. τ_{Ksi} is positive in a firm paying higher capital costs, and τ_{Ksi} is negative if the labor cost is relatively high. The ratio of the marginal productivity of labor relative to capital is different from the industry mean, and τ_{Ksi} is not equal to zero. Similarly, output distortion is measured when labor's share is different compared with what one would expect given the degree of industry elasticity of output with respect to labor. τ_{Ysi} represents the extent of the deviation of marginal revenue labor productivity $(1-\alpha)(P_{si}Y_{si}/L_{si})$ from the wage rate $\sigma/(\sigma-1)w$.

(5) $$1+\tau_{Ksi}=\frac{\alpha_s}{1-\alpha_s}\frac{\omega L_{si}}{RK_{si}}$$

(6) $$1-\tau_{Ysi}=\frac{\sigma}{\sigma-1}\frac{\omega L_{si}}{(1-\alpha_s)P_{si}Y_{si}}$$

Additionally, τ_{Ksi} and τ_{Ysi} in equations (5) and (6) converge to the industry mean once a hypothetical reform transpires. Note that the industry means of the distortions in the real data do not have to be zero. Meaningful interpretations of τ_{Ksi} and τ_{Ysi} refer to how much they differ from the industry mean instead of focusing on the degrees of the absolute deviation from zero.

Calculated values of τ_{Ksi} and τ_{Ysi} are reflected in the revenue productivity $TFPR_{si}$. All other factors are common, but this is not the case for the last instance in (7). Even firm productivity A_{si} does not make a difference with regard to $TFPR_{si}$. In the monopolistic competition frame, a firm with high productivity wants to lower its prices to raise its quantity of sales; thus, productivity itself is irrelevant with reference to the product of price and productivity. Only $(1+\tau_{Ksi})^{\alpha_s}/(1-\tau_{Ysi})$ is the single factor generating the variance of $TFPR_{si}$ within an industry.

(7) $$TFPR_{si}=P_{si}A_{si}=(\frac{\sigma}{\sigma-1})(\frac{R}{\alpha_s})^{\alpha_s}(\frac{\omega}{1-\alpha_s})^{1-\alpha_s}\frac{(1+\tau_{Ksi})^{\alpha_s}}{(1-\tau_{Ysi})}$$

The next hurdle is to measure the physical productivity A_{si}. This calculation is heavily dependent on the functional form of the CES aggregator of industry goods. With a strong assumption of the production form, physical productivity is recovered from the nominal output $P_{si}Y_{si}$. Because κ_s is common to firms

within an industry, it can be normalized as 1.

(8) $$A_{si} = \kappa_s \frac{(P_{si}Y_{si})^{\frac{\sigma}{\sigma-1}}}{K_{si}^{\alpha_s} L_{si}^{1-\alpha_s}} \quad \text{where} \quad \kappa_s = \omega^{1-\alpha_s}(P_s Y_s)^{-\frac{1}{\sigma-1}} / P_s$$

Industry TFP is given as in equation (9); it is produced by the sum of physical TFP(A_{si}) of each firm weighted by $\overline{TFPR_s} / TFPR_{si}$ (deviation from the average industry TFPR). When an individual firm's output and capital distortions (τ_{Ksi} & τ_{Ysi}) collapse to the average industry level, meaning a firm's $TFPR_{si}$ is equal to the industry $TFPR_s$, the industry TFP_s becomes equal to $\overline{A_s} = (\sum A_{si}^{\sigma-1})^{1/(\sigma-1)}$.

(9) $$TFP_s = \frac{Y_s}{K_s^{\alpha_s} L_s^{1-\alpha_s}} = (\sum_{i=1}^{M_s} (A_{si} \frac{\overline{TFPR_s}}{TFPR_{si}})^{\sigma-1})^{\frac{1}{\sigma-1}}$$

With (9) and the formula of $\overline{A_s}$, we can calculate the ratio of the current output (Y) to the efficient output ($Y_{efficient}$) via equation (10).

(10) $$\frac{Y}{Y_{efficient}} = \prod_s (\frac{TFP_s}{\overline{A_s}})^{\theta_s} = \prod_s (\sum_{i=1}^{M_s} (\frac{A_{si}}{\overline{A_s}} \frac{\overline{TFPR_s}}{TFPR_{si}})^{\sigma-1})^{\theta_s/(\sigma-1)}$$

We will make a comparison between the HK method and the reallocation effect. Before explaining reallocation effects in growth accounting, we assume that every firm has its own productivity level and that it is fixed. Essentially, the reallocation effect means that increasing the market share of a firm with higher productivity causes an economy-wide TFP improvement. The philosophy of the HK method is very similar to the basic concept of the reallocation effect except that it emphasizes firm heterogeneity caused not only by the firm's own productivity but also by distortion. When the government gives a particular firm a subsidy for its final products or provides access to inexpensive money, the marginal cost of the firm decreases as well as output increases relative to firms with identical productivity levels. Differing from the reallocation effect, the HK method suggests a quantitative degree of output loss caused by distortion, as explained above. In other words, the HK method presents the potential output if the allocative efficiency is maximized in the absence of distortion.

There are several limitations which apply to the HK method. First, misspecification error is one of the most crucial limitations, as the HK method relies fairly thoroughly on functional forms. If each firm has different levels of production elasticity of capital ($\alpha_{si} \neq \alpha_s$) caused by a technical gap in firms, τ_{Ksi} and τ_{Ysi} measured by the HK methodology consider the technical gap as a distortion. In fact, allocative efficiency is not clearly related to the technical gap or

to the capital elasticity of the firm. Even a social planner is not able to overcome this constraint of technological differences between firms. The assumption of identical markups for every firm also produces measurement error in reality; τ_{Ysi} represents this gap when the firm's markup differs according to size.

Second, as mentioned above, the HK method is available for measuring the efficiency of only intra-industry firms. Inter-industry reallocations of capital and labor are not allowed during an industry-specific shock. Depending on the industrial policy of the government, inter-industry resource allocations can be affected even when the efficiency of intra-industry firms is not influenced. The HK method cannot detect this type of inter-industry resource allocation. Considering that many emerging countries implement industrial policies, this shortcoming may be a crucial limitation when this method is applied to emerging countries.

Third, the HK method does not take entries or exits into account explicitly. According to the literature, the quantitative size of the extensive margin (entries and exits) on TFP is not negligible and is more important than the extent of the intensive margin (incumbent firms). Entries and exits are frequent in the manufacturing sector, and fluctuation in the TFPR variance can be affected by entries and exits. For instance, if increasing demand in a particular industry causes more firms to enter it, the TFPR variance can increase and result in an increase in misallocations in the HK method.

Oberfield (2013) partly overcomes these shortcomings of the HK methodology. Although Oberfield (2013) depends on functional forms and does not consider the effects of entries and exits, his methodology relaxes the strong assumption that each firm has the same elasticity of capital. An important improvement by the methodology of Oberfield (2013) is that the decomposition of intra-and inter-industry misallocation can be accomplished by allowing free movement of inputs between industries. The key is that input reallocation between firms depends not only on firm productivity A_{si} but also on the elasticity of capital α_{si}. This implicitly makes input movement between industries possible.

III. Empirical Results

In this section, we use survey data from the mining and manufacturing industries supplied by the Korea statistics office. This survey contains every establishment with more than 10 employees and reports value-added, sales, employment, and several types of capital stocks and investment. Data starting in 1990 is available, but 2010 is excluded due to serious numbers of missing values for capital stocks.

Before calculating the gains from hypothetical liberalization from distortions, we trim the 1% tails of $\log(TFPR_{si} / \overline{TFPR_s})$ and $\log(A_{si} / \overline{A_{si}})$ across industries. We then recalculate the variables used as components. In the periods analyzed, the standard industrial classification in Korea was revised three times (1991, 1998, and 2006); hence, the time-series discontinuities at the industry level may affect the results. Thus, we unify the industry classification for all periods at the three-digit level with the eighth industrial classification. A unit of this data is an establishment

TABLE 1—DISPERSION OF TFPQ VALUES

	Korea				
	1992	1998	2001	2008	2012
S.D.	0.98	1.03	0.96	0.97	0.94
p75-p25	1.29	1.38	1.22	1.14	1.10
p90-p10	2.46	2.65	2.43	2.37	2.27
N	45,334	40,697	50,736	55,401	59,967
	The U.S.				
	1977		1987		1997
S.D.	0.85		0.79		0.84
p75-p25	1.22		1.09		1.17
p90-p10	2.22		2.05		2.18
N	164,971		173,651		194,669
	China				
	1998		2001		2005
S.D.	1.06		0.99		0.95
p75-p25	1.41		1.34		1.28
p90-p10	2.72		2.54		2.44
N	95,980		108,702		211,304

TABLE 2—DISPERSION OF TFPR VALUES

	Korea				
	1992	1998	2001	2008	2012
S.D.	0.55	0.62	0.54	0.57	0.55
p75-p25	0.70	0.80	0.66	0.65	0.63
p90-p10	1.42	1.60	1.37	1.38	1.31
N	45,334	40,697	50,736	55,401	59,967
	The U.S.				
	1977		1987		1997
S.D.	0.45		0.41		0.49
p75-p25	0.46		0.41		0.53
p90-p10	1.04		1.01		1.19
N	164,971		173,651		194,669
	China				
	1998		2001		2005
S.D.	0.74		0.68		0.63
p75-p25	0.97		0.88		0.82
p90-p10	1.87		1.71		1.59
N	95,980		108,702		211,304

and not a firm in this survey. For convenience, we use firm and establishment interchangeably and only connote a difference between them when necessary.

Table 1 and Table 2 show that variances of TFPQ and TFPR for Korea, the U.S. and China. The data for the U.S. and China are quoted from Hsieh and Klenow (2009), but observations of Japan are not available. TFPQ indicates the physical productivity of a firm. This variable shows that the variance for Korea is greater than that of the U.S. and similar to that of China. However, the dispersion of TFPQ is not directly related to distortions in resource allocation, as technological differences between firms can be endogenous.

The variance of TFPR is closely related to misallocation. The standard deviation and the interquartile range of the Chinese data are largest in both cases, while those in the U.S. data are the smallest. From Tables 1 and 2, we know that the variance of the TFPR of China is larger than that of Korea despite the fact that the variance of the TFPQ of Korea is larger than that of China for 1998, indicating less distortion

TABLE 3—OUTPUT DISTORTION AND CAPITAL DISTORTION IN KOREA AND JAPAN

	Korea		Japan	
	$\tau_{Y_{si}}$	$\tau_{K_{si}}$	$\tau_{Y_{si}}$	$\tau_{K_{si}}$
S.D.	0.584	25.656	1.63	14.17
N	1,154,534	1,154,534	3,565,341	3,565,341

TABLE 4—TFP GAINS AFTER EQUALIZING TFPR WITHIN INDUSTRIES

	Korea	Japan	The U.S.	China
No output distortion & No capital distortion				
TFP gap, $Y / Y_{efficient}$	0.6482	0.679	0.733	0.502
TFP gain, $Y_{efficient} / Y - 1$	54.27%	47.18%	36.60%	99.17%
No output distortion				
TFP gap	0.816	0.810	N.A.	N.A.
TFP gain	22.52%	23.40%	N.A.	N.A.

in the Korean manufacturing sector compared to that in China. The variance of the TFPR of China converges to a level similar to that of Korea; thus, distortion in the Chinese manufacturing sector appears to be relieved. Caution is needed when making a comparison of the results of each country because differences in sampling can affect the results.

Table 3 shows the standard deviations of τ_{Ysi} and τ_{Ksi} from the Korean data from 1990 to 2012 and from the Japanese data from 1981 to 2008. Korea has a smaller standard deviation of τ_{Ysi} than Japan, but Korea's variance of τ_{Ksi} is two times higher than the Japanese case. Literally, Korea has greater distortion in the relative price of capital to labor than that in Japan.

The ultimate purpose of this paper is to calculate the ratio of realized output to efficient output, as in (10). Table 4 shows these results by country. Keeping in mind that a careful comparison is required due to the different periods and sampling methods, a descending order in terms of allocative efficiency is as follows: U.S., Japan, Korea and China. In the absence of output and capital distortion, there is no significant difference between Korea (0.648) and Japan (0.679). If output distortion is eliminated, the ratios of current to efficient output are 0.816 for Korea and 0.810 for Japan, also indicating that distortion in either output market or capital market does not contribute solely to the results.

We also investigated changes in allocative efficiency in the time-series data. Figure 1 indicates that $Y / Y_{efficient}$, remaining at 0.68 until the mid-1990s, decreased sharply during the Asian Financial Crisis during the years 1997 and 1998. In the early 2000s, it recovered to 0.65 but declined toward 2008 (the year the financial crisis struck), after which it showed a minor increase in 2012.

The downward trend of allocative efficiency is still valid in plants that survive for all of the years and remain in the sample. In the sample, the entry rates are between 12 and 27% and the exit rates range from 11% to 26%. In order to check if the time series of allocative efficiency ($Y / Y_{efficient}$) was mainly driven by changes in the productivity distribution caused by entering and exiting plants, we construct a

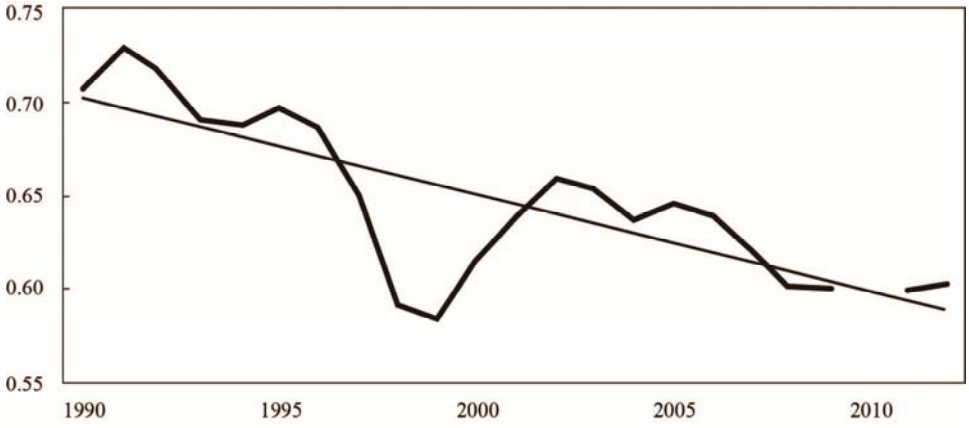

FIGURE 1. CHANGES IN THE ALLOCATIVE EFFICIENCY $Y / Y_{efficient}$ OF KOREA'S MANUFACTURING FIRMS

Note: 1) Korea's 2010 Survey of Mining and Manufacturing is a complete enumeration survey. However, the capital items account has been removed due to omitted data to ensure continuity.

Source: Author's calculation of the Survey of Mining and Manufacturing by Statistics Korea.

TABLE 5—TRENDS OF ALLOCATIVE EFFICIENCY IN KOREA AND JAPAN

	Korea		Japan		
	90-99	2000-12	80-90	90-99	2000-08
No output distortion & No capital distortion					
TFP gap	0.6742	0.6264	0.696	0.676	0.666
TFP gain	48.32%	59.64%	43.7%	47.9%	50.2%
No output distortion					
TFP gap	0.8445	0.7925	0.830	0.809	0.792
TFP gain	18.41%	26.18%	20.5%	23.6%	26.3%

balanced panel. It was found that the time-series pattern of allocative efficiency does not change much compared to the original exercise. This finding implies that the fluctuation of misallocation is not greatly influenced by dropping unproductive plants and adding productive ones in Korea.

Table 5 summarizes the trend changes in Korea and Japan. By calculating the TFP gain ($Y_{efficient} / Y - 1$), which indicates the potential advantage of eliminating misallocations, we find that the inefficiency of Japanese manufacturing firms increased through the 1980s to the 2000s. The gain from hypothetical liberalization in Japan is 43.7% in 1980, 47.9% in 1990 and 50.2% in 2000 (until 2008). In Korea, the TFP gain becomes 59.64% in the 2000s, and it stood at 48.32% in 1990, meaning that instances of misallocation has deepened at a rapid speed compared to that in Japan. When only the output market distortion is removed, the TFP increases by 18% in the 1990s and by 26% in the 2000s. Although both Korea and Japan undergo worsening degrees of capital distortion, the speed of the increase in inefficiency in the capital market is much faster in Korea.

This result may change depending on the data treatment for outliers. In this paper, we cut the top and bottom of 1%, following Hsieh and Klenow (2009). When the threshold increases by 2%, the level of inefficiency slightly falls off as variance of $TFPR_{si}$ decreases. However the increasing trend of inefficiency remains apparent. Extreme values of the upper tail and lower tail do not control the empirical results.

The aggregation of industrial classification can affect the results. We calculate $Y/Y_{efficient}$ with the two-digit industrial classification. In this case, the variance of $TFPR_{si}$ increases; thus, $Y/Y_{efficient}$ falls compared to the use of three-digit aggregation. However, the degree of intensified misallocation does not make a major change at all. Moreover, the pattern of the decreasing trend in efficiency remains valid even when the two-digit classification scheme is used. Therefore, aggregation of the industrial classification does not alter the results significantly.

Productivity studies show that the contribution of a firm's entry and exit is quantitatively considerable. Although the method does not consider extensive margins, entries and exits occur frequently in the actual data. A balanced panel was established to observe the effects of entries and exits on $Y/Y_{efficient}$. We constructed the balanced panel of data in two parts. The first consists of establishments surviving from 1990 to 1999, and the second contains firms which were operating from 2000 to 2012. If only firms who survive for the entire period (1990-2012) are selected, a limited number of firms remain, leading to survivorship bias. The results from the two parts of the balanced panel depict time-series trends which are quite similar to that in the original result. The effects of entries and exits appear to be limited for this measure of allocative efficiency.

Oberfield (2013) overcomes the limitations of the HK method by allowing that ① firms within an industry can have different degrees of capital elasticity to output, and ② factors can move between industries. Following Oberfield (2013), we can confirm that the technological differences between firms do not determine the main result of the decreasing trend of allocative efficiency. An important point from the results in Oberfield (2013) is that within-industry misallocation is the main contributor to inefficiency in the Korean case, in contrast to the Chilean manufacturing sector.

Figure 2 suggests that the overall allocative efficiency, M_{both}, has trended downward from the 1990s onward. M_{both} consists of M_W and M_B, which denote the extent of misallocation for intra-industry and inter-industry cases, respectively. As shown in the graph, intra-industry allocative efficiency drives M_{both} down, while inter-industry misallocation does not show significant changes over time. Note that the level of M_W, the counterpart of allocative efficiency in the HK method, is slightly higher than that in the HK method because M_W allows individual capital elasticity, which was regarded as a form of distortion by Hsieh and Klenow (2009).

Table 6 presents TFP decomposition caused by the technological advances of

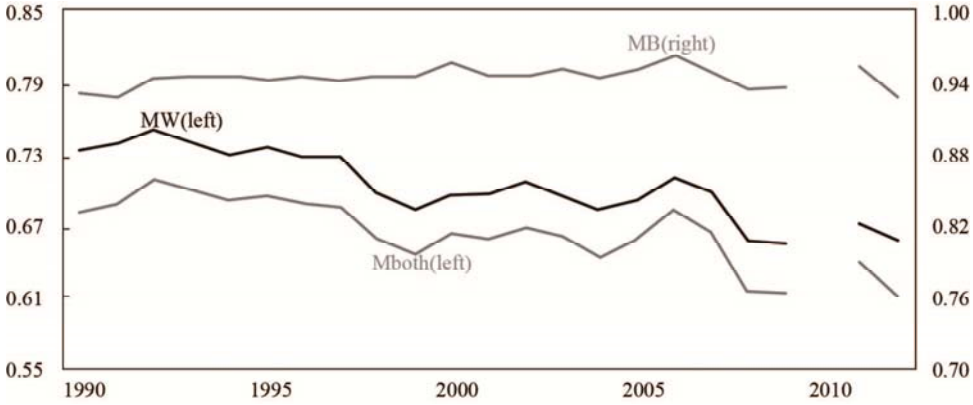

FIGURE 2. CHANGES IN INTRA-INDUSTRY AND INTER-INDUSTRY ALLOCATIVE EFFICIENCY LEVELS

TABLE 6—DECOMPOSITION OF TFP

(UNIT: %)

	TFP	Resource allocative efficiency	Technological progress
1991~1999	5.90	-0.60	6.50
2000~2009	5.39	-0.53	5.92
Total	5.63	-0.57	6.19

Note: 1) The author adopted Oberfield's (2013) method and calculated the TFP using the original data from the Survey of Mining and Manufacturing. Results regarding TFP growth rates may differ depending on the calculation method used.

a representative firm and the changes in misallocation. In the Korean manufacturing sector, the annual TFP (Solow residual) growth rate was approximately 5.6% on average during 1990-2009 and the allocative efficiency ($d\ln M_W + d\ln M_B$) worsened by nearly 0.6% on an annual basis. In other words, the TFP growth rate would have reached 6.2% if there had been no degeneration of allocative efficiency. "0.6%" is not a minor amount considering that increasing TFP growth by 1% cannot easily be done with artificial policies.

IV. Analysis of Changes in Allocative Efficiency

In order to analyze the factors behind the changes in allocative efficiency, this section investigates at the efficiency distributions of firms by size and age.

$Y / Y_{efficient}$ can be calculated at the level of each establishment. The current output of a firm is proportional to its productivity and distortion in (11). The percentage deviations of the current output $P_{si}Y_{si}$ from the efficient output $P_{si}^{*}Y_{si}^{*}$ are expressed by (12). If $\tau_{Y_{si}}$ is smaller and $\tau_{K_{si}}$ is larger than the industry mean, it is beneficial for the aggregate output for this plant to increase its production

because $P_{si}^{*}Y_{si}^{*}$ is greater than $P_{si}Y_{si}$. If there is a negative correlation between $\tau_{Y_{si}}$ and $\tau_{K_{si}}$, $P_{si}^{*}Y_{si}^{*} / P_{si}Y_{si}$ would be amplified.

(11) $$P_{si}Y_{si} \propto \frac{(1-\tau_{Ysi})^{\sigma-1} A_{si}^{\sigma-1}}{(1+\tau_{Ksi})^{\alpha_s(\sigma-1)}}$$

(12) $$\log(P_{si}^{*}Y_{si}^{*}) - \log(P_{si}Y_{si}) = (\sigma-1)\log(\overline{\frac{1-\tau_{Ys}}{1-\tau_{Ysi}}}) - \alpha_s(\sigma-1)\log(\overline{\frac{1+\tau_{Ks}}{1+\tau_{Ksi}}})$$

Table 7 shows how the size of the initially large vs. small plants would change if the TFPR were equalized in each country. The entries are the unweighted shares of plants. The rows are the initial (actual) plant size quartiles, and the columns are the bins of the efficient plant size relative to the actual size: 0%–50% (where the plant should shrink by half or more), 50%–100%, 100%–200%, and 200+% (where the plant should at least double in size). Although the aggregate output increases in the optimal case when distortions are removed, many establishments of all sizes would shrink. In the U.S., China, and India in Hsieh and Klenow (2009), initially large plants are less likely to shrink and are more likely to expand. The remarkable feature in Korea is that this pattern is much more pronounced. For the top size quartile in Korea, the share of plants which produce less than their efficient level is 16.3% (100%-200% + 200+%), while this rate is 10.6% in the U.S. and 12.0% in China. In contrast, the share of plants which should shrink in terms of efficient output is higher in Korea than in other countries. For the plants located in the bottom quartile, the share of plants which produce more than their optimal output is 19.1% (0%-50% + 50%-100%), while it is 16.7% in the U.S. and 16.4% in China.

TABLE 7—PERCENTAGES OF PLANTS: ACTUAL SIZE VS. EFFICIENT SIZE

(UNIT: %)

	Korea (1990-2012)			
	0-50%	50-100%	100-200%	200%
Bottom quartile	10.4	8.7	4.1	1.8
3rd quartile	5.5	9.1	6.8	3.7
2nd quartile	3.8	7.7	7.2	6.3
Top size quartile	2.5	6.2	7.4	8.9
	The U.S. 1997			
	0-50	50-100	100-200	200
Bottom quartile	4.7	12.0	4.3	4.1
3rd quartile	4.5	9.8	5.4	5.4
2nd quartile	4.4	9.6	5.8	5.1
Top size quartile	4.4	10.0	6.7	3.9
	China 2005			
	0-50	50-100	100-200	200
Bottom quartile	10.5	5.9	4.5	4.2
3rd quartile	8.5	6.0	5.2	5.4
2nd quartile	7.3	5.9	5.3	6.6
Top size quartile	7.0	6.1	5.4	6.6

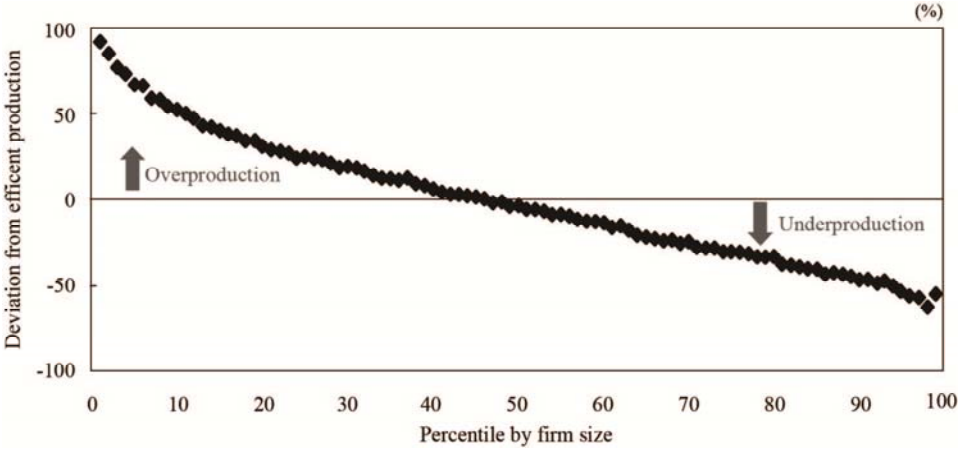

FIGURE 3. DEVIATION OF ACTUAL OUTPUT AGAINST EFFICIENT OUTPUT BY FIRM SIZE

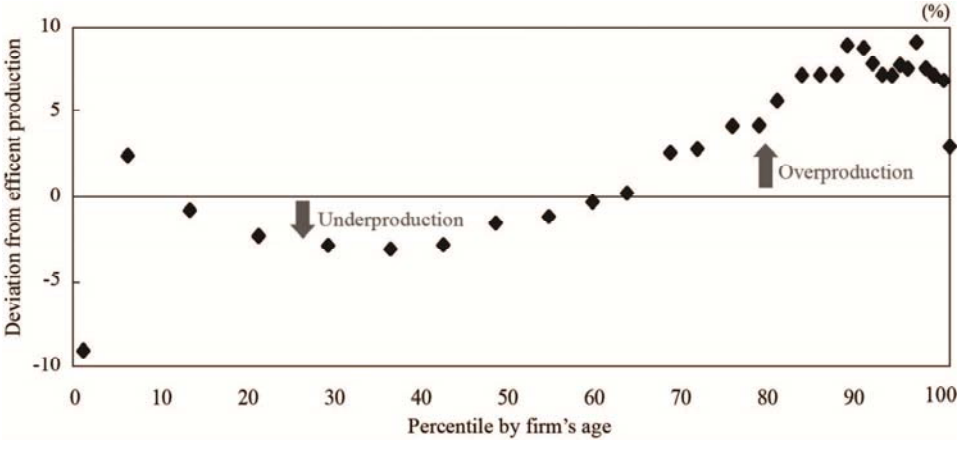

FIGURE 4. DEVIATION OF ACTUAL OUTPUT AGAINST EFFICIENT OUTPUT BY FIRM AGE

Figure 3 shows the related average $100 \times \log(Y_{si} / Y_{efficient,si})$ of each percentile with reference to the plant size. This shows that the larger the firm, the less the firm produces compared to its efficient level, with the degree of the tendency quantitatively intensifying. This figure reinforces the results shown in Table 7.

The results also imply that many small-sized establishments are uncompetitive in the Korean manufacturing sector. In other words, small-sized establishments exploit factor inputs which exceed the optimal amounts. Thus, it is favorable for the entire economy when unproductive small-sized establishments reduce their production and more productive, large plants take resources which had been held by these small establishments. Indirectly, it also implies that unselective support for all small-sized establishments may impede the healthy circulation of exits by unproductive plants and entries by productive ones.

Figure 4 displays the patterns of production compared to the efficient level in terms of the plant age. Generally, it was found that the younger the firms, the less they produced compared to the efficient production level and the older the firms,

the more excessively they produced. The older the firm, the further the deviation of the actual output from the efficient level above zero, yet the extent of deviation does not exceed 10%.

Financial friction and adjustment costs are commonly mentioned as the main culprits behind instances of capital misallocation. Financial constraints are considered as an important difference between rich and poor countries, leading to productivity gaps by country (Banerjee and Duflo 2005). However, Midrigan and Xu (2009) point out that financial friction contributes little to account for the variance of the average product of capital using Korean manufacturing survey data identical to ours.[2] Hsieh and Klenow (2009) examine how much differences in adjustment costs explain the extent of misallocation for the U.S., China, and India. The variations in the adjustment costs explain only a modest amount of the overall dispersion in the TFPR. Midrigan and Xu (2009) also mention the role of adjustment costs and conclude that the quantitative effect of adjustment costs on misallocation is marginal.

V. Concluding Remarks

Allocative efficiency in the Korean manufacturing sector (0.65) showed a downward trend from the 1990s to 2012. On average, allocative efficiency is approximately 0.65, which is lower than that of the U.S. (0.73), similar to the level in Japan (0.68) and higher than that in China (0.50). If allocative efficiency does not decrease in the case of Korea, the manufacturing TFP growth rate is estimated to climb by an additional 0.6%p on an annual basis.

As the productivity gap between large and small plants in Korea is higher than those in other countries, this paper empirically presents that the over-production of small-sized plants and the under-production of relatively large plants are noticeable. This implies that one type of distortion which makes the market less efficient may be extensive subsidies to small-sized plants. However, caution is needed when interpreting this empirical result regarding productivity differences by size. First, even within the same industry which is narrowly defined, products by small and large plants may not be homogeneous. In such a case, a unilateral comparison of productivity by size is not appropriate because goods markets are different. Second, there is a possibility that unfair vertical relationships enjoyed by large firms with small subcontractors may contain a measurement error of productivity. The productivity of large firms may be overestimated and that of small firms may be underestimated. As such, there may be deviations in estimations of efficient production rates. Third, we should be more careful when devising firm-related policy implications using the results, as our analysis is dependent on plant-level data and not firm-level data. Even bearing these possibilities in mind, the clear pattern of under- and over-production by size, as in Figure 3, stresses that this consistent pattern by size likely does not derive only due

[2]Midrigan and Xu (2014) note that the potential effect of financial frictions can be large because it impedes the entry of productive plants without enough money into the market.

to heterogeneous products or from abuse by the market dominance of large firms.[3] From this perspective, this paper recommends that it is desirable to concentrate more on selective support for younger and smaller firms with consideration of their growth potential rather than on providing unilateral support to all small-sized firms.

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[3]Specifically, large firms' partners in vertical relationships are usually medium-sized firms. However, nearly 77% of plants (establishments) in the manufacturing sector are in fact firms (a firm has one plant), and 87% of them are plants with fewer than 20 employees. This suggests that the left side of Figure 3 may not be related to large contractors.

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Household Over-indebtedness and Financial Vulnerability in Korea: Evidence from Credit Bureau Data†

By YOUNG IL KIM, HYOUNG CHAN KIM, AND JOO HEE YOO*

Financial soundness in the household sector matters for financial stability and for the real economy. The level of household debt in Korea raises concern about the financial soundness of the household sector due to its size, growth rate and quality. Against this backdrop, we assess the financial vulnerability of borrowers based on an analysis of credit bureau (CB) data, in which the actual credit activities of most individuals are recorded at a high frequency in Korea. We construct over-indebtedness indicators from the CB data and then assess the predictability of forthcoming defaults. Based on the over-indebtedness indicators, we show how borrowers are distributed in terms of over-indebtedness and how the over-indebted differ from average borrowers in terms of their characteristics. Furthermore, we show how the aggregate credit risk in the household sector would change under macroeconomic distress by analyzing how each borrower's credit quality would be affected by adverse shocks. The findings of this paper may contribute to assessing household debt vulnerability and to enhancing regulatory and supervisory practices for financial stability.

Key Word: Household debt, Over-indebtedness, Credit Risk, Stress Test, Financial Stability
JEL Code: D12, D14, G21

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I. Introduction

Financial soundness in the household sector matters for financial stability as well as for the real economy. The global financial crisis of 2007-09 has shown how the financial soundness of the household sector can affect both the financial system and the real economy. The level of household debt in Korea raises concern over financial stability due to its large size, high growth rate and compositional quality. The amount of household debt relative to GDP (or income) is large compared to those of OECD countries, as shown in Figure 1. Household debt has been growing faster than household income, suggesting that the debt-repayment ability of households has weakened. Furthermore, the share of loans with high-interest rates from non-bank financial institutions has increased rapidly, as shown in Figure 2,

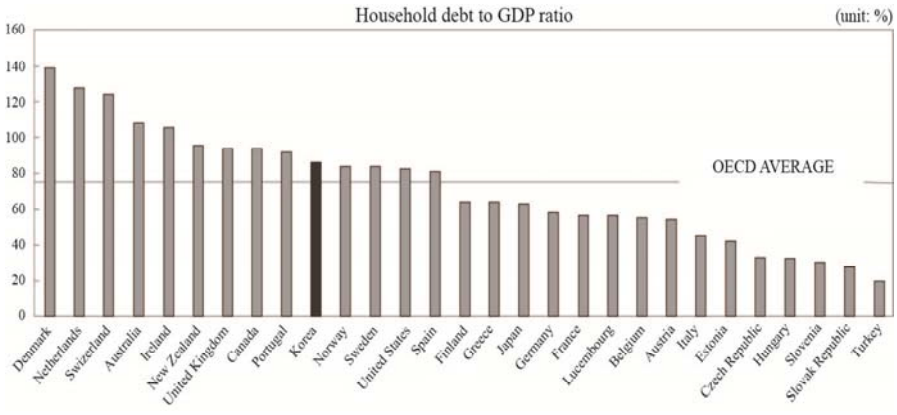

FIGURE 1. HOUSEHOLD DEBT: CROSS-COUNTRY COMPARISONS

Source: OECD, Euro Stat., FRED (2012).

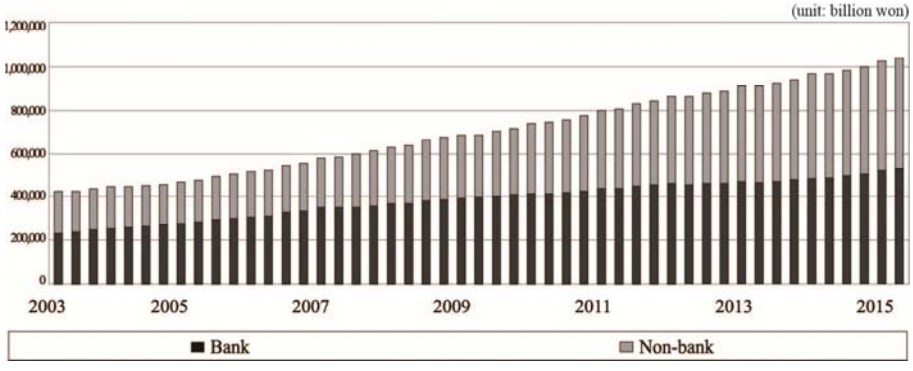

FIGURE 2. HOUSEHOLD DEBT BY FINANCIAL SECTOR: BANK VS. NON-BANK

Note: Non-bank financial institutions in the figure include nonbank depository institutions and other financial institutions such as insurance companies, pension funds and public financial institutions.

Source: Bank of Korea.

while the share of balloon-payment loans is relatively large as well. In short, these phenomena give rise to concerns about household debt vulnerability. Against this backdrop, we investigate household over-indebtedness and its associated financial vulnerability based on an analysis of credit bureau (CB) data, which contain the actual credit activities of most individuals in Korea.

The vulnerability of household debt at the aggregate level may be determined by how much debt is held by the overly indebted with weak repayment abilities. In other words, the larger the amount of debt held by borrowers bearing high credit risks, the larger the share of debt that could be vulnerable. In order to evaluate this issue, we need criteria by which to assess which debtors are exposed to higher credit risks. The current study constructs several common indicators of over-indebtedness from the CB data and assesses the predictability of forthcoming defaults. The CB data employed in this study cover most individuals in Korea and contain their actual credit activities, including the amount of debt and debt services, the types of debt contracts and lending institutions used, delinquencies, their credit score, and more. In order to understand who the overly indebted are, we identify them using over-indebtedness indicators and characterize the vulnerable indebted in comparison with average borrowers. Furthermore, we show how both the distribution of borrowers in terms of the credit score dimension and the aggregate default risk would be affected by sudden changes in the macroeconomic environment.

The concept of over-indebtedness may be ambiguous. Hence, it may not be easy to define what it means in practice – see Betti, Dourmashkin, Rossi, and Yin (2007) and D'Alessio and Lezzi (2013) for discussions. Nevertheless, it may be reasonable to relate the concept to a condition of difficulty in debt repayment. Over-indebtedness indicators can be assessed and compared with one another with regard to their ability to predict defaults in the forthcoming year. We examine common over-indebtedness indicators - such as the credit score, debt service ratio (DSR), the DSR of unsecured debt holders, the loan to income (LTI) ratio, and the number of credit commitments – in terms of their performance as predictors of defaults. We show that the credit score is a dominant predictor of defaults, while the number of credit commitments may have some additional explanatory power for forthcoming defaults. The DSR may have weak predictability of defaults in the forthcoming year, though it loses its explanatory power if the credit score and the number of credit commitments are controlled.

Based on the over-indebtedness indicators, vulnerable borrowers can be identified and analyzed. We document the characteristics of the over-indebted to shed some light on who they are and how they differ from average borrowers in several dimensions. Based on such primary indicators as the credit score, the number of credit commitments, and the DSR, the overly indebted are characterized. We show that most over-indebtedness indicators recount common characteristics of the over-indebted. For instance, all over-indebtedness indicators show that over-indebted borrowers tend to depend heavily on non-bank financial institutions as the sources of their loans. In addition, those classified as overly indebted in terms of one indicator tend to be classified as overly indebted by other indicators as well. However, different indicators appear to elucidate idiosyncratic characteristics of the over-indebted. For example, the over-indebted with poor credit scores tend to

show lower DSRs than average, while the over-indebted with multiple credit commitments show much higher DSRs than average. Thus, caution may be needed when using those indicators for purposes of risk monitoring or supervision.

We are often asked about how the share of borrowers at risk would be affected by a sudden change in the macroeconomic environment (e.g., GDP, interest rates, asset prices). The over-indebtedness indicators can be used to assess borrowers' credit risks at the individual as well as at the aggregate level by relating them to default probabilities. The stress test conducted in the current study may guide us to assess the financial vulnerability of the indebted at the aggregate level due to macroeconomic distress. In the current study based on CB data, we show how macroeconomic distress would affect the credit quality of borrowers as well as the aggregate default rates in household credit markets. If the historical scenarios of past financial crises such as the Asian financial crisis ('97~99) and the global financial crisis ('07~09) were assumed to reoccur, the aggregate default risk would then increase dramatically, with the larger impact resulting from the stress scenario of the Asian financial crisis, the worse macroeconomic condition. It should be noted that the stress test proposed in the current study shows uniquely how each borrower's credit quality would deteriorate under macroeconomic distress, which in turn would change the aggregate default risk.

It is also important to note that the current study analyzes CB data, whereas many previous studies of household debt vulnerability used household survey data. The survey data are subject to certain limitations in their assessments of financial vulnerability in spite of their strengths regarding various kinds of information about household characteristics and activities, as they are subject to errors and psychological biases and use insufficient representations of their populations (Lusardi and Tufano 2009; Karlan and Zinman 2008; Zinman 2009).[1] The CB data employed in the current study cover most individuals and financial institutions in Korea and contain their actual credit activities at a high frequency, with the information used by financial institutions during their loan generation processes. Thus, the CB data may provide an advantage when assessing and monitoring the levels of actual credit risk for the purpose of financial regulation and supervision relative to the household survey data. For instance, we can assess the over-indebtedness indicators for their predictability of forthcoming defaults because the CB data keep track of defaults on loans, whereas the survey data provide insufficient information on default events.

In order to assess and monitor the financial vulnerability of household debt, which may be important for prudential regulation and supervision, we often rely on stress tests that assume extremely adverse macroeconomic conditions. The stress test proposed in the current study is distinct from those in previous studies in terms of the intermediary mechanism through which macroeconomic shocks affect the aggregate credit risk of household debt. First, we use the credit score, a dominant predictor of forthcoming defaults, as a part of an intermediary channel through which macroeconomic distress affects the aggregate credit risk. Second, we shed some light on how the relationship between the over-indebtedness indicator and

[1]In addition, the CB data are updated at a higher frequency, up to the most recent period, in contrast to the survey data; hence, this data may be a useful source of information for risk monitoring and supervisory practices.

default rates would be affected by macroeconomic fluctuations. It should be noted that the relationship between over-indebtedness measures and default rates are not constant but may change as macroeconomic conditions fluctuate. For instance, the default rates jumped dramatically for borrowers even with constant debt servicing burdens during the global financial crisis of 2007-09 (Kim and Byun 2012). These distinct features of the stress test proposed in the current study may contribute to enhancing the assessments and monitoring of the aggregate credit risk emerging from household debt.

The remainder of this paper is organized in the following order. Previous studies are reviewed in relation to the current study in section II. Data and descriptive statistics are shown in section III. Indicators of over-indebtedness are presented and assessed for their ability to predict defaults in the near term future in section IV. Based on the preceding analysis, the characteristics of the over-indebted are discussed in section V. Stress tests are then conducted to assess the financial vulnerability of the indebted against macroeconomic distress in section VI. Concluding remarks are made in section VII.

II. Relation to Previous Studies

The current study assesses the financial soundness of the household sector by analyzing CB data based on measures of over-indebtedness and a stress test, motivated by the rising level of concern over financial stability. This issue has received a great amount of attraction from central banks and financial supervisory bodies since the global financial crisis in 2007-09, leading to regular and irregular assessments of the household sector from the perspective of financial stability. Nonetheless, there is still room to improve the tools and related databases. Our approach with regard to this issue is to assess household sector credit risk at the aggregate level from an analysis of microeconomic data, as this dataset suitably represents the population and is updated at a high frequency. As argued by Mian and Sufi (2010), an analysis of such microeconomic data can shed some light on household liability issues and provide guidance to policymakers.

Household debt, the liability side of the household sector, has been somewhat neglected in the literature compared to the asset side of the balance sheet (Zinman 2015). As noted by Zinman (2015), “the neglect of household debt is pronounced relative to its cousin literatures on corporate debt.” Concerning the credit risk assessment of interest in the current study, we find a large volume of literature on corporate debt, whereas there is much less to be found on household debt. For example, we address how macroeconomic conditions would affect the credit quality of borrowers by showing how the obligors would migrate in the dimension of credit ratings. With regard to this question, we find much work on corporate debt in the literature (e.g., Trück 2008; Koopman *et al.* 2009; Bangia *et al.* 2002; Carling *et al.* 2007; Bonfim 2009; Altman and Rijken 2004), whereas we find few similar works on household debt in spite of the rising demand from policy circles to our knowledge. This situation may have arisen because credit risk assessment in the household sector had not received much attention until its importance was recognized in the face of the historically unprecedented large debt accumulation

followed by the subprime mortgage crisis of 2007-09, while microeconomic data such as CB data have only become available in the recent years. Against this backdrop, we aim to contribute to the assessment of financial stress in the household sector.

The global financial crisis provided a critical moment for extensive reforms on financial regulation and supervision with an emphasis on macro-prudential policies (Galati and Moessner 2013; IMF 2011a; IMF 2011b; FSB, IMF, and BIS 2011; Lim *et al.* 2011; Bernanke 2011). An important element for effective financial supervision is to monitor and to assess emerging risks in the financial system in relation to the real economy. It should be noted that the most frequently used macro-prudential instruments in practice have been caps on LTV and DTI ratios that are aimed at borrowers from the household sector among the wide range of policy tools (Claessens 2014; Darbar and Wu 2015); suggesting that the assessment and monitoring of risks emerging from the household sector may be in high demand from regulatory and supervisory bodies. Motivated by the rising interest from policy circles, we aim to assess the credit risk emerging from the household sector in relation to macroeconomic conditions.

Various indicators and measures of household over-indebtedness or financial vulnerability have been discussed in recent studies (e.g., Bankowska *et al.* 2014; Bryan, Taylor, and Veliziotis 2010; Civic Consulting 2013; Disney, Bridges, and Gathergood 2008; D'Alessio and Iezzi 2014; Shubhasis 2008). We find that common indicators have been used in the form of debt levels (or debt servicing burdens) out of debt payment abilities, such as income, assets or consumption spending. For example, the ratios of debt to assets, debt to income, and debt service to income were used as indicators of household debt vulnerability by the ECB (2013) and in Costa and Farinha (2012) and Ehrmann and Ziegelmeyer (2014). Other indicators have also been explored as an extension of these common indicators. These measures include information about the cash flow aspects of income-consumption-debt repayment streams or the asset-side liquidity of balance sheets as used to assess financial vulnerability, particularly in Albacete and Lindner (2013) and Ampudia *et al.* (2013), among others. In addition, self-reported household-level information about overdue debt, the number of credit commitments and financial difficulty has been used to measure the seriousness of debt problems. As concerns are raised about household debt in Korea, recent studies have assessed the vulnerability of Korean household debt based on household-level survey data. Several of those studies utilized indicators of financial vulnerability based on household survey data (e.g., Kim and Yoo 2013; Karasulu 2008; Kim *et al.* 2014).

Most of the earlier studies on household over-indebtedness or vulnerability are based on an analysis of household survey and self-reported data, which may be subject to errors or psychological biases while the data are released with significant time lags of a few years. In contrast to these previous studies, we use credit bureau (CB) data, which contain the actual credit activities and transactions recorded in most financial institutions and are used by financial institutions for the loan generation process. Based on the CB data, indicators of over-indebtedness are constructed for each borrower. The over-indebtedness indicators of interest in the current study are the credit score, the debt-service-to-income ratio (DSR), the DSR

of unsecured loan holders, the loan-to-income ratio (LTI), and the number of credit commitments. These indicators are studied for their properties and for their ability to predict defaults in the subsequent year.

The aforementioned indicators may be used to identify the over-indebted or to study the associated vulnerability levels. Once the indicators of over-indebtedness are constructed, we then analyze who the overly indebted are in terms of their characteristics based on these indicators. There are several studies that analyzed the characteristics of the over-indebted based on household survey data (e.g., Bryan, Taylor, and Veliziotis 2010, Civic Consulting 2013; Disney, Bridges, and Gathergood 2008; D'Alessio and Iezzi 2014). Because survey data contains various details about households, the data may provide very useful information about the characteristics of the over-indebted. In comparison with survey data, CB data may provide a somewhat limited range of information about borrowers' characteristics. Nevertheless, the CB data in the current study cover most individuals and financial institutions participating in the credit market in Korea and provide accurate information about the liability side of borrowers. It is also updated at a high frequency.

We conduct stress tests on household debt to assess how the credit risk of the household sector would be affected at the aggregate level by adverse shocks to macroeconomic environments. We can find evidence in previous studies that macroeconomic conditions serve as determinants of the default risk of retail loans in banks' portfolios, though they do not show how obligors may migrate in terms of certain aspects of the credit risk, such as credit ratings, in response to adverse shocks (e.g., Mayer *et al.* 2009; Agarwal and Liu 2003; Rinaldi and Sanchis-Arellano 2006; Louzis *et al.* 2012; Büyükkarabacak and Valev 2010). We can also find studies on stress tests for an assessment of household debt vulnerability but based on household survey data (e.g., Kim and Yoo 2013; Karasulu 2008; Albacete and Fessler 2010; IMF 2012; Shubhasis, Djoudad, and Terajima 2008). These studies use information about cash flows and balance sheet positions to identify vulnerable households. They examine how the share of vulnerable households and their debt holdings change in response to fluctuations in macroeconomic conditions. In contrast, we use prominent indicators of debt vulnerability based on CB data and examine how the credit quality of borrowers would change if the macroeconomic environment changed dramatically. It should be noted that the stress tests using household survey data in previous studies may have limitations in timely credit risk assessments due to both insufficient information about borrowers' default events and a few years of lags until their release. As concerns about household debt have increased in Korea, several recent Korean studies have used CB data to analyze its vulnerability (e.g., Hahm, Kim, and Lee 2010; Kim, Chang, and Choi 2012; Lee *et al.* 2014). However, these previous studies which rely on CB data ignored how macroeconomic conditions would affect the dominant predictors of defaults, instead focusing on indicators which only weakly predict defaults. It is also important to note that the current study sheds some light on how the relationship between the over-indebtedness measure and default rates would change in the face of adverse macroeconomic conditions. These features of the stress test conducted in the current study may enhance the risk assessments and monitoring of household credit risk at the aggregate and at the obligor level.

III. Data and Measures of Over-indebtedness

We use data from the National Information and Credit Evaluation (NICE) CB to analyze the vulnerability of Korean household debt. The CB dataset contains actual credit activities from most financial institutions and covers most household loans in Korea.[2] It contains information such as individual characteristics, debt contracts, delinquencies, types of lending institutions, and estimated incomes. We sample nearly one million individuals from the CB data in the analysis; hence, they are assumed to constitute a nationally representative random sample of individuals with credit records. Table 1 presents descriptive statistics pertaining to the sample data set, including the sample sizes, estimated incomes, ages, credit scores, debt amounts and DSR and LTI values. We note that high-income groups tend to have more borrowers with larger amounts of debt and debt service levels but better credit scores and lower default rates as well in comparison with other income groups.

As a criterion to define over-indebtedness and to assess the financial vulnerability of debt holders, we use the likelihood that borrowers will not repay their debt or interest. In particular, the likelihood of being in arrears for more than 90 days during the subsequent year is used as a criterion to assess the extent of over-indebtedness. In other words, the more borrowers are exposed to default risks, the more they are considered to be overly indebted. We examine such indicators as the credit score, the DSR, the DSR of unsecured loan holders, the LTI, and the number of credit commitments, after which we assess their ability to predict defaults in the forthcoming year.

Regarding the credit score, borrowers are divided into ten groups from the lowest to the highest credit quality based on the NICE credit scoring system. The debt service to income ratio (DSR) is computed as the debt-servicing burden out of income, estimated based on the CB data, as recommended by the Financial Supervisory Commission (FSC). DSR ratios are also computed for unsecured loan holders. Loan to income ratios (LTI) are computed as the ratio of outstanding debt

TABLE 1— DESCRIPTIVE STATISTICS OF THE SAMPLE DATA (2014Q4)

Income quintile	1	2	3	4	5	Total
Sample size (No. of individuals)	156,332	247,279	209,873	193,111	188,742	995,337
Median age	32	38	45	47	49	45
Average estimated income (\10,000) [a)]	1,308	2,174	2,940	3,700	5,438	3,114
Median credit score	5	5	4	2	2	4
Default rates (%) [b)]	3.5	3.7	3.3	2.7	1.4	2.7
Share of borrowers (%)	32	24	40	54	69	43
Average debt of borrowers (\10,000)	2,710	3,850	4,090	5,268	11,290	6,367
Median DSR of borrowers (%)	20	21	19	19	26	22
Median LTI of borrowers (%)	66	65	61	64	109	74

Note: a) Incomes in the table are on an annual basis. b) Default rates are the ratio of those who are not currently in default but experience a default within one year.

[2]These encompass nearly every type of financial institution operating in Korea, including domestic banks, branches of foreign banks, securities companies, insurance companies, savings banks, credit card leasing and finance companies, agricultural and fisheries cooperatives credit unions, and community credit cooperatives.

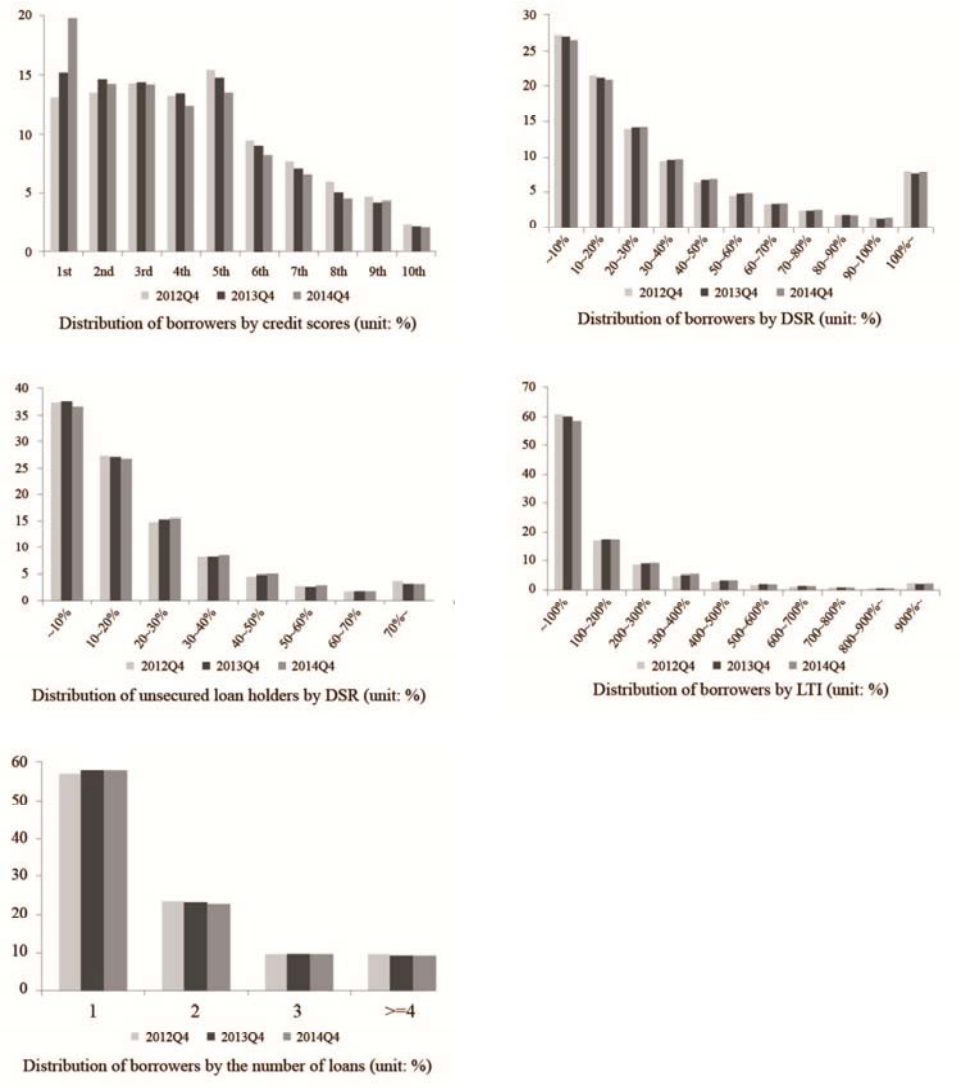

FIGURE 3. DISTRIBUTION OF BORROWERS BY OVER-INDEBTEDNESS INDICATORS

out of the estimated individual incomes. The number of credit commitments refers to how many financial institutions each borrower is indebted to at the moment.

Based on the indicators of over-indebtedness, we note how borrowers are distributed on the spectrum of each measure, as shown in Figure 3. Regarding the credit score, nearly 70% of borrowers are rated at equal to and above the fifth credit score, while the remaining 30% of borrowers are rated equal to or below the sixth credit score. We also find that the share of borrowers in each group of credit scores decreases as credit scores deteriorate for the remaining 30% of borrowers. Regarding the DSR, approximately 70% of borrowers have DSRs below 40%, while the remaining 30% have DSRs higher than 40%. We also observe that the number of borrowers decreases as the DSR increases. With respect to the DSR of

unsecured debt holders, about 80% of borrowers with unsecured loans have DSRs below 30%, with the remaining 20% having DSRs higher than 30%. Again, the share of borrowers with unsecured loans decreases as their DSR increases. Regarding the LTI, nearly 80% of borrowers have LTIs lower than 200%, and the number of borrowers decreases as the LTI increases. Approximately 80% of borrowers have loans from less than or equal to two financial institutions while the remaining 20% of borrowers have loans from more than or equal to three financial institutions.

IV. Over-indebtedness and the Likelihood of Default

We analyze how over-indebtedness indicators are related to default rates in the near-term future in order to assess how well they reflect borrowers' default risk. Indicators that show a strong and clear relationship with the default rate in the forthcoming year may be preferred over indicators that do not show such a relationship. In addition, an indicator of over-indebtedness can be compared with other indicators in terms of their ability to predict a default. In the current study, a state of default is defined as being in arrears for more than 90 days. By computing the frequency of defaults in the forthcoming year corresponding to different values of over-indebtedness, we can quantify the relationship between the over-indebtedness indicators and the rates of default in the subsequent year. We assess over-indebtedness indicators such as the credit score, the DSR, the DSR of unsecured debt holders, the LTI, and the number of credit commitments for their association with default rates in the forthcoming year.

In the following, we analyze how well the over-indebtedness indicators predict the frequency of defaults in the subsequent year.[3] Borrowers credit-rated from the sixth to the tenth levels show higher than average default rates in the forthcoming year, as shown in Figure 4, while borrowers credit-rated from the eighth to the tenth record default rates above 20%, much higher than the average default rate. Borrowers with DSRs above 40% record default rates in the subsequent year higher than average, as shown in Figure 4. The default rates of borrowers tend to increase very slowly as the DSR increases. Changes in the DSR do not significantly influence changes in default rates in the subsequent year. In fact, the association between changes in default rates and changes in the DSR appear much weaker as compared to the credit score. With respect to the DSR of unsecured loan holders, those with DSRs higher than 30% record default rates in the forthcoming year higher than average, as shown in Figure 4. However, the overall pattern of default rates associated with the DSR of unsecured loan holders is similar to that of the general indebted. Regarding the LTI, borrowers with LTIs above 600–700% record default rates in the next year higher than average, as shown in Figure 4. The default rates of borrowers tend to increase very slowly in the LTI range above 300–400% as the LTI increases. However, changes in the LTI do not appear to show a significant association with changes in default rates in the subsequent year.

[3]The frequency of defaults within the subsequent year is computed as the ratio of those who are not currently in default but experience a default within the next one year.

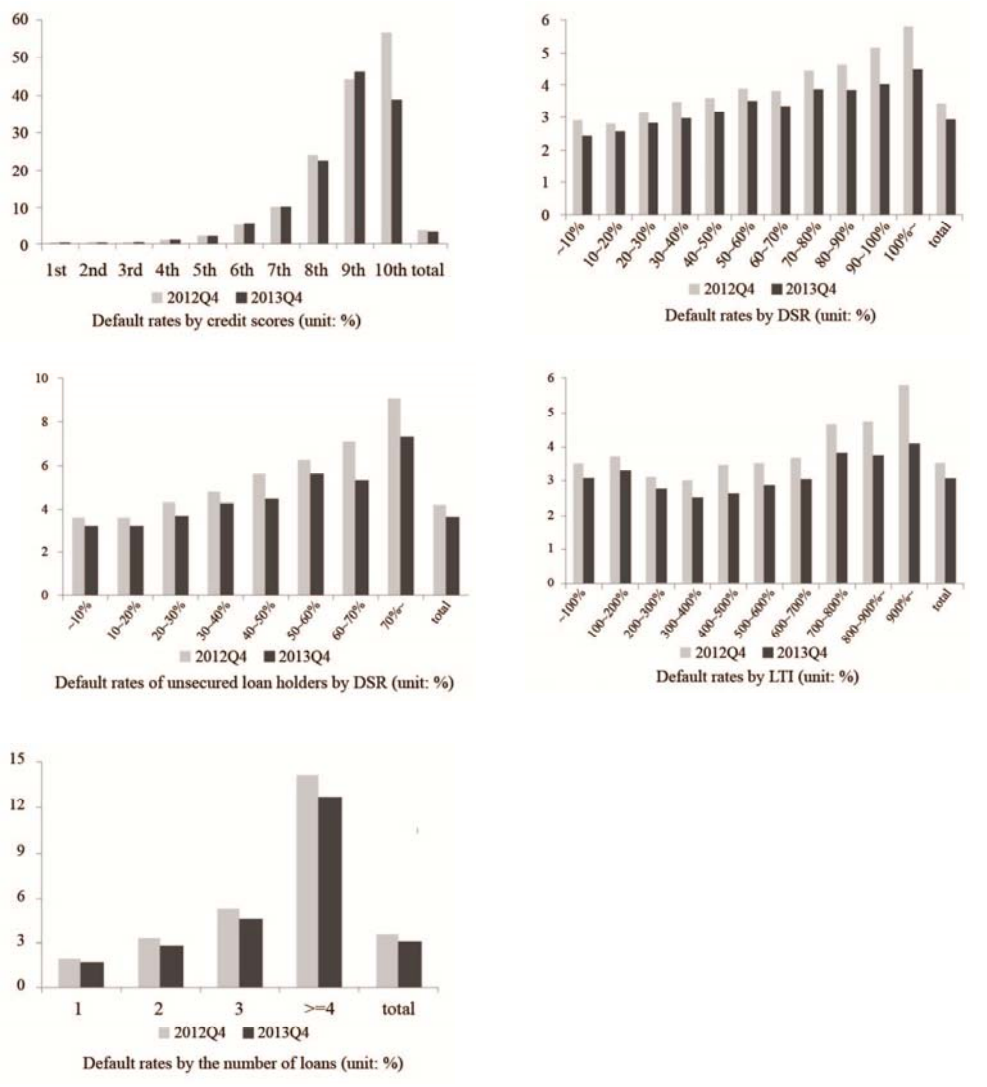

FIGURE 4. OVER-INDEBTEDNESS AND DEFAULT RATES IN THE FORTHCOMING YEAR

For instance, default rates do not change much in the LTI range up to 700%, as shown in Figure 4. Multiple credit commitments are shown to be strongly associated with default rates in the near future. Borrowers with loans from up to two financial institutions show default rates lower than or similar to the average in the subsequent year, as depicted in Figure 4. However, their default rates become higher than average if the number of credit commitments reaches three. Default rates become much higher than average for borrowers with loans from four or more financial institutions. Thus, the number of credit commitments may serve as a useful predictor of default in the near future.

The credit risk of borrowers may be better characterized by a combination of over-indebtedness indicators than by a single indicator. Thus, it may be valuable to analyze how borrowers are distributed and how are associated with default rates in

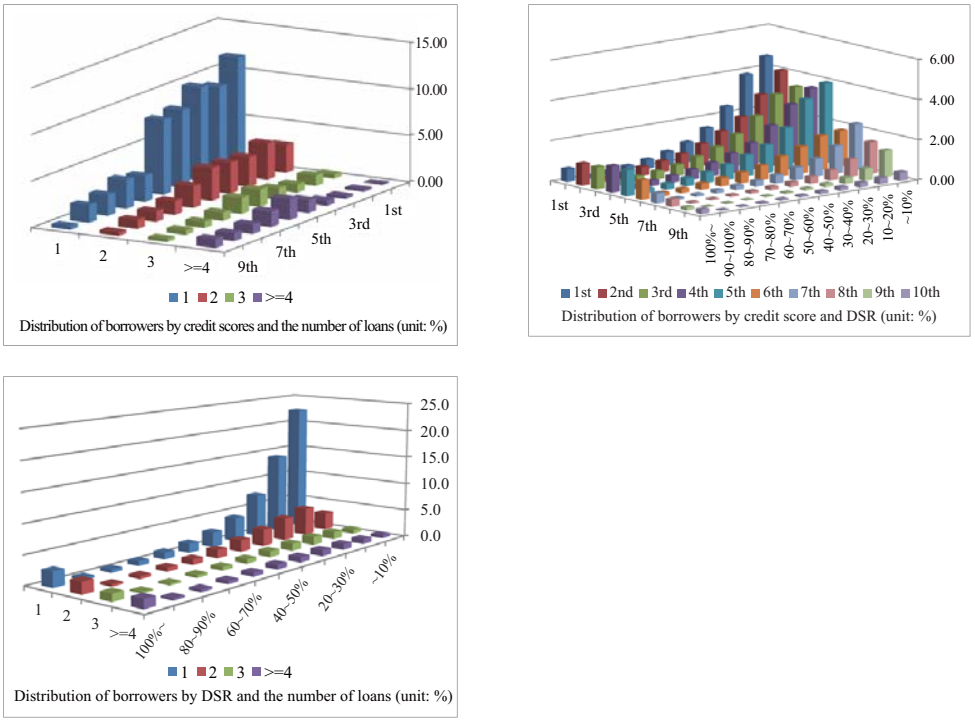

FIGURE 5. DISTRIBUTION OF BORROWERS IN MULTIPLE DIMENSIONS OF OVER-INDEBTEDNESS

multiple dimensions of the over-indebtedness indicators. With respect to the multidimensionality of both credit scores and credit commitments, a large share of borrowers can be found among those with relatively better credit scores and fewer credit commitments, as shown in Figure 5. With respect to the multidimensionality of both credit scores and DSRs, more borrowers exist in a region of relatively better credit scores and lower DSRs, as shown in Figure 5. Regarding the multidimensionality of both DSRs and the number of credit commitments, a large proportion of borrowers are found among those with fewer credit commitments and relatively lower DSRs.

The likelihood of defaults may also be evaluated in multiple dimensions combining multiple indicators of over-indebtedness. The default rates in the forthcoming year appear most strongly associated with the credit score out of the indicators of the credit score, the DSR, and the number of credit commitments, as shown in Figure 6. If the DSR and the number of credit commitments are compared with regard to their ability to predict default, the number of credit commitments is more strongly associated with the default rate. Note that the number of credit commitments may still contain some explanatory power for defaults in the near future, even with credit scores controlled, as shown in Figure 6. In contrast, the DSR does not show a clear pattern in relation to default rates if the credit scores are fixed, as shown in Figure 6. In short, the credit score can be used as a dominant predictor of the default rates in the forthcoming year, while the number of credit commitments has some additional explanatory power with regard to default rates even when credit scores are taken into account.

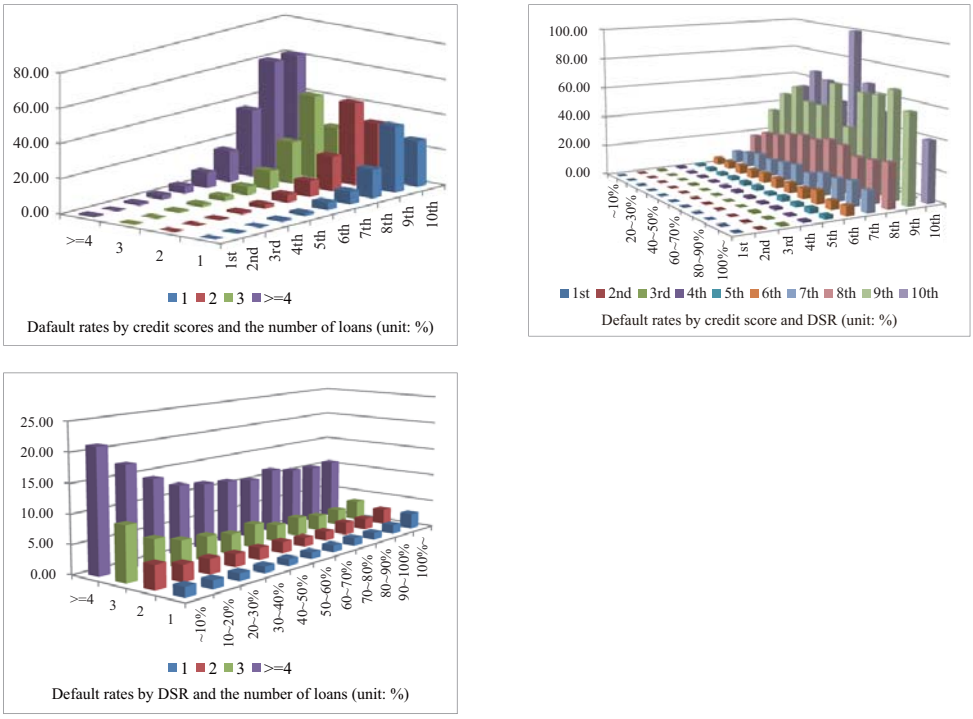

FIGURE 6. DEFAULT RATES IN IN MULTIPLE DIMENSIONS OF OVER-INDEBTEDNESS

Thus, the credit score may be the dominant predictor of default in the forthcoming year, while the number of credit commitments may have some additional predictability. The DSR may carry some information about default risk; however, it loses its explanatory power to predict a default in the near future if better predictors such as the credit score and the number of credit commitments are taken into account. Table 2 shows how the over-indebtedness indicators are correlated with one another. Credit scores show a positive association with the number of credit commitments, implying that borrowers with poorer credit rates tend to borrow from a larger number of financial institutions. Meanwhile, the debt burden ratio indicators of the DSR and LTI are not clearly related to credit scores, whereas they are weakly and positively correlated with the number of credit commitments. In other words, the amount of debt or the debt service burden relative to income may not necessarily provide information about their credit score, whereas higher debt-burden ratios in terms of the DSR and LTI are positively associated with the number of credit commitments. The DSR and LTI are highly correlated with each other, implying that the information delivered by the respective indicators may overlap somewhat.

We estimate simple logistic regression models in order to assess the predictability of the over-indebtedness indicators for defaults in the forthcoming year. The regression results from (1) to (4) in Table 3 show that all indicators are statistically significant, although the suitability of each indicator differs significantly from one another. Regression (1) shows that the credit score alone has

TABLE 2—CORRELATIONS AMONG OVER-INDEBTEDNESS INDICATORS (2014Q4)

Correlations	Credit score	DSR	LTI	No. of loans
Credit score	1			
DSR	0.0249	1		
LTI	-0.0431	0.9071	1	
No. of loans	0.3676	0.1195	0.1281	1

Note: p-values of all correlation coefficients are less than 0.0001.

TABLE 3—LOGISTIC REGRESSION RESULTS (2013Q4)

	(1)	(2)	(3)	(4)	(5)
Credit score	0.4503				0.4053
	(0.0040)				(0.0045)
DSR		0.0006			0.0017
		(0.0001)			(0.0002)
LTI			0.0001		-0.0006
			(0.0000)		(0.0001)
No. of loans				0.3814	0.2108
				(0.0041)	(0.0045)
Constant	-5.9231	-3.5642	-3.5239	-4.3657	-6.1620
	(0.0040)	(0.0062)	(0.0098)	(0.0148)	(0.0295)
No. of obs.	420,619	405,313	413,364	413,209	405,253
C-statistic	0.823	0.548	0.504	0.687	0.844

Note: Numbers in parentheses are standard errors.

excellent predictability with regard to a default in the subsequent year according to the C-statistic (above 0.8).[4] The regression results of (2) and (3) show that the DSR or LTI may not be useful predictors by themselves according to the C-statistic (close to 0.5), while LTI is virtually uninformative about forthcoming default events. In addition, the LTI coefficient estimate is even unstable across prediction models, as depicted by (3) in comparison with (5), for example. Regression (4) indicates that multiple credit commitments show good predictability of defaults, although the predictability is weaker than that of the credit score. Regression (5), including all of the indicators, shows improved predictability compared to all other regression models, though it is only slightly better than model (1). In short, we find that the credit score is the dominant predictor of default in the near future. Including additional over-indebtedness indicators such as the number of credit commitments and the DSR and LTI in the regression specification may improve the predictability, but only slightly. We also find that the regression results appear consistent with the pattern of default rates, as shown above in the various dimensions of the over-indebtedness indicators. It should be noted that the assessment of the indicators is for their short-term (one year) predictability of defaults. The predictability of DSR and LTI, which are considered to reflect fundamental repayment capability of debtors, may improve over a longer time horizon.

[4]The C-statistic can be used to evaluate model predictability. It ranges from 0.5 to 1. If the C-statistic is close to 0.5, the model is interpreted as not useful for predictions.

V. Characteristics of the Over-indebted

It is important to understand who the overly indebted are in order to design policies to deal with household debt problems. The findings above demonstrate that the over-indebtedness indicators can predict near-term defaults but with different strengths of predictability. Credit scores are strong in terms of their ability to predict defaults, while the number of credit commitments and DSR may provide some additional information on the likelihood of a default. Below, we analyze how the over-indebted in terms of the credit score, the multiple credit commitments and the DSR are characterized in comparison with average borrowers. Over-indebted borrowers are analyzed on several metrics, including their age, income, debt level, geographical residence area, lending institutions, and on other metrics relevant to the overly indebted.

First, borrowers with poor credit ratings are compared with borrowers with better credit ratings and with average borrowers, as shown in Table 4. There does not appear to be a significant difference in age between borrowers in the different credit ratings groups. Borrowers rated below the seventh credit score represent about 19% of all borrowers. They tend to have smaller income and debt levels than those with better credit ratings. Their residences are generally equally distributed between the capital area and other regions, while borrowers with better credit ratings are concentrated somewhat more in the capital area than in other regions.

TABLE 4—CHARACTERISTICS OF THE OVER-INDEBTED BY CREDIT SCORE, NO. OF LOANS, AND DSR (2013Q4)[5]

Borrower characteristics		Credit score		No. of loans		DSR		TOTAL
		1st–6th ratings	7th–10th ratings	< 3	>=3	< 60%	>= 60%	
Median age		46	45	46	44	45	49	46
Share of borrowers (%)		81	19	81	19	84	16	100
Average income relative to the total (%)		104	82	99	104	99	107	100
Average debt relative to the total (%)		108	65	86	161	54	341	100
Income quintile (by median income)		4	3	4	4	4	4	4
Share of borrowers in	Capital (%)	53	51	52	55	53	52	52
	Non-capital (%)	47	49	48	45	47	48	48
Share of borrowers from	Only banks (%)	49	12	50	8	45	24	42
	Only non-banks (%)	33	66	40	34	39	39	39
	Both banks and non-banks (%)	19	22	10	59	16	37	19
Average number of loans		1.8	2.6	1.3	4.1	1.7	2.5	1.8
Median credit score		3	8	3	6	4	4	4
Median DSR (%)		21.9	16.9	17.5	43.7	16.4	94.4	21.1
Median LTI (%)		77. 5	46.1	58.2	134.2	52.7	386.3	70.7

[5]The cut-off levels for over-indebtedness in the above analysis are the seventh rating for the credit score, 3 for the number of credit commitments, and 60% for the DSR. Those with a credit score equal to or less than the seventh credit rating tend to record high default rates, while they may have difficulty to obtain an unsecured loan from a bank. Those with loans from three or more financial institutions are often categorized as 'borrowers with multiple credit commitments' and tend to show high default rates, as shown in Figure 4. The DSR is related to default rates, but it may not be as good a predictor as the credit score or the number of loans. Nevertheless, DSR may represent potential risk on a longer time horizon, while those with high DSRs beyond 50-60% may have difficulty obtaining a mortgage.

Nearly two thirds of borrowers with low credit ratings borrow only from non-bank financial institutions, while 12% of them borrow only from banks. In contrast, about half of borrowers with better credit ratings borrow only from banks, while one third of them borrow only from non-bank financial institutions. Borrowers with low credit ratings tend to have more credit commitments than those with better credit ratings. Nevertheless, they show much lower DSRs and LTIs than those with better credit ratings due to their much smaller debt amounts.

Second, borrowers with multiple credit commitments are compared with those who have fewer credit commitments in Table 4. There appears to be a trivial difference in age between those with more credit commitments and those with fewer. Borrowers with three or more credit commitments represent nearly 19% of all borrowers. They tend to have levels of income similar to that of the average borrower but much higher levels of debt than the average borrower. Thus, their DSRs and LTIs are much higher than average. They are concentrated slightly more in the capital area than in other regions. They show a very high dependence on non-bank financial institutions in comparison with the average borrower; 34% of them borrow only from non-bank financial institutions and 59% of them borrow from both banks and non-bank financial institutions. Borrowers with multiple loans tend to have worse credit ratings compared to the average borrower.

Third, borrowers with heavier debt burdens in terms of the DSR are compared with borrowers with lighter debt burdens in Table 4. Borrowers with heavier debt burdens are four years older than those with less debt. Borrowers with DSRs above 60% comprise approximately 16% of all borrowers. They tend to have levels of income similar to that of the average borrower but a much higher level of debt. Thus, their DSRs and LTIs are much higher than average, even multiple times higher. Their residences are distributed slightly more in the capital area than in other regions. There does not appear to be a significant difference in terms of residential region between borrowers with a high DSR and the average borrower. High-DSR borrowers show high dependence on non-bank financial institutions when they borrow. It was found that 39% of them borrow only from non-banks and 37% of them borrow from both banks and non-banks. In addition, they tend to borrow from more financial institutions. Nevertheless, they do not show a significant difference in terms of their credit scores when compared with those less burdened with debt, indicating that high-DSR borrowers may not necessarily be riskier than the average borrower according to their credit scores.

We have documented the characteristics of the over-indebted to shed light on who the over-indebted are and how they differ from average borrowers in several dimensions. We find that most over-indebtedness indicators recount the common characteristics of the over-indebted. For example, over-indebted borrowers tend to depend heavily on non-bank financial institutions as sources of their loans. In addition, those classified as overly indebted in terms of one indicator tend to be classified as overly indebted by other indicators as well. However, it is important to note that different indicators appear to elucidate the idiosyncratic characteristics of the over-indebted. For instance, the over-indebted with poor credit scores tend to have lower DSRs than average, while the over-indebted with multiple credit commitments have much higher DSRs than average. Thus, special caution needs to be exercised regarding the appropriate use of over-indebtedness indicators for

financial supervision or risk-monitoring purposes.

VI. Financial Vulnerability of the Indebted: Sensitivity Analysis

We are often asked about how adverse shocks to the economy would affect the amount of credit risks to which borrowers are exposed at the aggregate level, which would then affect the soundness of financial institutions with the risky loans in their asset portfolios. In order to answer this question, we may conduct stress tests, which may be a useful tool for assessing financial vulnerability from a forward-looking perspective. By relating the degree of over-indebtedness to the probability of default (PD), we may be able to compute borrowers' exposure to default risk at the aggregate level and analyze how they are expected to change in response to shifts in the macroeconomic environment. For example, we may compute the frequency of defaults associated with each credit score (CS_i); hence, borrower i's credit score (CS_i) can indicate his/her probability of default (PD_i). If each borrower i were assigned a probability of default (PD_i) based on the relationship between their over-indebtedness and its corresponding default rates, the average likelihood of default would then be computed by $(\sum_i PD_i)/N$, where N is the total number of borrowers. This expression indicates that the aggregate credit risk would increase if more borrowers were associated with higher probabilities of default. To assess the financial vulnerability of the indebted, we conduct the following stress test. First, stress scenarios reflecting macroeconomic distress are juxtaposed against baseline scenario. Stress scenarios may reflect hypothetical changes in macroeconomic environments (e.g., the GDP, interest rates, and asset prices, among others). They may be based on specific historical events (e.g., the Asian financial crisis, the global financial crisis) or on a distribution of macroeconomic variables of interest. Next, a shift in macroeconomic conditions changes the distribution of borrowers in terms of over-indebtedness. Because over-indebtedness is associated with the probability of default, the distributional change in the over-indebtedness dimension would reassign a new probability of default (PD_i) to each borrower i. The new PD_i reassigned to each borrower i may allow us to predict the average default rates under the stress scenario.[6] If more borrowers were associated with more over-indebtedness, the share of borrowers exposed to higher levels of credit risk would then increase. In short, a shift in macroeconomic conditions may affect the aggregate credit risk in terms of the average probability of default by changing the distribution of borrowers in terms of over-indebtedness.

Because the credit score is the dominant predictor of default in the near future among the over-indebtedness indicators discussed above, we analyze below how the distribution of borrowers would change in the credit score dimension if the

[6]In turn, we may assess how much bank assets are exposed to the default level of risk, although we do not assess the impact on the soundness of bank assets in this paper. That is, the change in borrowers' credit risk would affect the balance sheets of the banking sector in terms of exposure at default (EAD) and expected loss (EL).

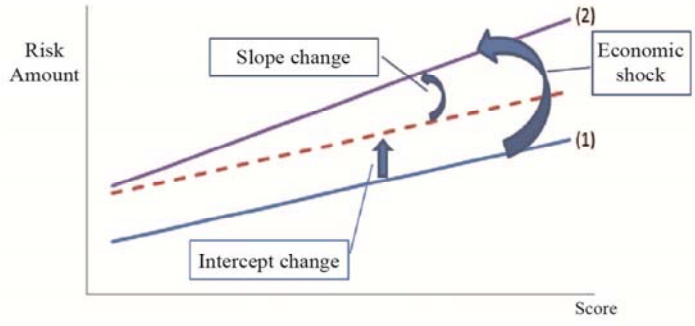

FIGURE 7. ECONOMIC DISTRESS AND
THE RELATIONSHIP BETWEEN CREDIT SCORES AND RISK AMOUNTS

economy were hit by an adverse macroeconomic shock.[7] Suppose that the credit score is related to the amount of risk in a straight-line relationship, as depicted in Figure 7, where the risk amount may be measured by $ln(odds) = ln((1-p)/p)$, where p denotes the borrower default rate. Macroeconomic distress may change the relationship between the credit score and the risk amount (RA). We may decompose the change in the relationship into a change in the intercept and a change in the slope. Figure 7 shows how the change in the relationship from (1) to (2) may be decomposed into the intercept change and the slope change. A change in the intercept may reflect equal changes in the amount of risk at each credit score, while a change in the slope may reflect unequal changes in the amount of risk across different credit scores - see Capuzzo (2011) for a conceptual discussion and some applications. We may capture the change in the relationship by changes in credit scores in order to keep each credit score associated with a certain amount of risk. In other words, macroeconomic distress would change the distribution of borrowers in the credit score dimension, i.e., credit migration, if we would like to keep default rates associated with each credit score fixed.

The baseline scenario and the stress scenario are denoted here as t_0 and t, respectively. The relationship between the credit score (CS) and $\ln(odds)$ under the baseline scenario (t_0) may be written as $\ln(odds)_{t0} = \beta_{t0} CS_{t0} + \alpha_{t0}$, while the relationship between the credit score (CS) and $\ln(odds)$ under the stress scenario (t) may be written by $\ln(odds)_t = \beta_t CS_t + \alpha_t$. If the macroeconomic environment shifted from the baseline to the stress scenario, the relationship between the credit score (CS) and $\ln(odds)$ would change and could then be represented by a change in α $(\Delta\alpha \equiv \alpha_t - \alpha_{t0})$ and a change in β $(\Delta\beta \equiv \beta_t - \beta_{t0})$. Because we would like to keep the risk amount associated with each credit score fixed, the change in the risk amount should be captured by the risk-equivalent change in the credit scores. We can compute the risk-preserving credit score change

[7]Note that the credit score alone is similar in terms of predictability to the prediction model including all the indicators together, as shown in Table 3.

$(\Delta \mathrm{CS} \equiv CS_t - CS_{t0})$ by solving $\beta_{t0} CS_{t0} + \alpha_{t0} = \beta_t (CS_{t0} + \Delta CS) + \alpha_t$. This equation can be rearranged for ΔCS, which can be expressed in terms of $\Delta\alpha$ and $\Delta\beta$ as shown below.

$$\Delta CS = -\left(\Delta\beta / \beta_t\right) CS_{t0} - \left(\Delta\alpha / \beta_t\right) \tag{1}$$

This equation can also be written as $CS_{t0} = \left(\beta_t / \beta_{t0}\right) CS_t + \Delta\alpha / \beta_{t0}$ to show how the credit score at $t\,(CS_t)$ is related to the risk-equivalent credit score at t_0 (CS_{t0}). Equation (1) indicates how changes in the intercept $(\Delta\alpha)$ and in the slope $(\Delta\beta)$ may lead to changes in credit scores (ΔCS).

If α_t and β_t were given as functions of the macroeconomic variables (X_t), i.e., $F_\alpha(X_t)$ and $F_\beta(X_t)$, respectively, the shift in the macroeconomic condition would then change the credit score according to equation (1). In order to compute the change in the credit score, α_t and β_t may need to be estimated as functions of the macroeconomic variables (X_t). Note that $F_\alpha(X_t)$. and $F_\beta(X_t)$ can be estimated for mortgage borrowers and other borrowers separately in order to take differences in risk characteristics into account. Figure 8 shows the time series of α_t and β_t for mortgage borrowers. In this paper, α_t and β_t are estimated as functions of macroeconomic conditions (X_t) consisting of the misery index (unemployment rate + inflation rate - GDP growth rate), corporate bond yields, and stock returns, among other factors.[8] Below are the estimated models for α_t and β_t for mortgage borrowers.[9]

$$\widehat{\alpha_t} = -0.01758\left(ur + cpi - gdp\right)_{t-14} - 0.0616\left(\mathrm{int}\,s\right)_{t-7} + 0.341\left(stock\right)_{t-4} + 5.3948 + 0.276\left(dummy\right)$$ [10]

$$\widehat{\beta_t} = 0.00003514\left(trend\right) + 0.00003462\left(ur + cpi - gdp\right)_{t-7} - 0.000056\left(\mathrm{int}\,s\right)_{t-12} + 0.0095$$ [11]

According to the estimation results of α_t, the misery index and interest rates tend to decrease $\ln\left(odds\right)$ while stock returns tend to increase $\ln\left(odds\right)$, consistent with our sense of the credit risk, noting that $\ln\left(odds\right)$ is inversely related to the level of default risk.

[8]Similar macroeconomic variables have been used to explain default rates of retail loans in BOKST-07, one of stress test models of the Bank of Korea – see Moon (2008) for details.

[9]In the estimation models, ur, cpi, gdp, ints, and stock denote unemployment rates, CPI inflation rates, GDP growth rates, interest rates (nominal), and stock returns (nominal) at a monthly frequency, respectively. GDP growth rates at a monthly frequency are computed by interpolating the quarterly GDP growth rates.

[10]R^2=0.48. The model includes a dummy variable to reflect the upgrade of the credit scoring system at 2010.

[11]R^2=0.92

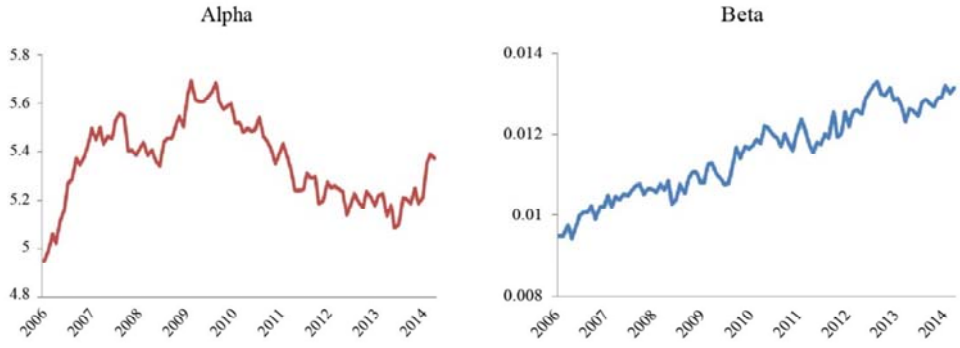

FIGURE 8. INTERCEPT (α) AND SLOPE (β) OF THE RELATIONSHIP BETWEEN $\ln(odds)$ AND CREDIT SCORE

If α_t and β_t were given as estimated functions of the macroeconomic variables, respectively, the change in the macroeconomic condition $(X_{t0} \to X_t)$ would then bring about changes in the credit score $(CS_{t0} \to CS_t)$ according to equation (1). As each borrower's credit score changes, the distribution of borrowers in terms of their credit scores would then change. If borrowers were redistributed in this credit score dimension, each borrower i would be reassigned a new PD $(PD_{i,t})$ corresponding to their new credit score $(CS_{i,t})$. We can then compute the average PD under the stress scenario by $(\sum_i PD_{i,t}) / N$, as discussed above. In short, we may be able to predict how the distribution of borrowers would change in terms of the credit score and how the aggregate default risk would change in terms of the average PD if macroeconomic conditions worsened.

We show the stress test results for the aggregate credit risk below while assuming a reoccurrence of historical events, such as past financial crises. The stress scenarios of interest may be macroeconomic conditions during the period of past crises such as the Asian financial crisis of 1998-99 (AFC) and the recent global financial crisis of 2008-09 (GFC).[12] Figure 9 shows how the distribution of borrowers would change in terms of the credit score dimension against the baseline scenario if each historical stress scenario reoccurred. We find that macroeconomic distress would redistribute borrowers from better credit scores to worse credit scores. It was also found that more migration would occur from the upper to the middle credit score range while less migration would occur from the middle to the lower credit score range. In addition, the impact of stress scenarios on the credit migration would be much stronger for the case of the Asian financial crisis, the relatively worse macroeconomic condition. As borrowers migrate into credit score regions of higher default probability (PD) levels, the overall credit risk is expected to increase. Figure 9 shows that the aggregate credit risk in terms of the average

[12]As stress scenarios, the GDP growth (%), unemployment (%), inflation (%), KOSPI returns (%) and 3yr. corporate bond yield change (%p) are assumed to be -0.5, 4.7, 6, -21 and 7.8 for the AFC, and 1.4, 3.3, 4.3, -32 and 1.2 for the GFC, respectively. The historical stress scenarios correspond to the periods that maximize the PD.

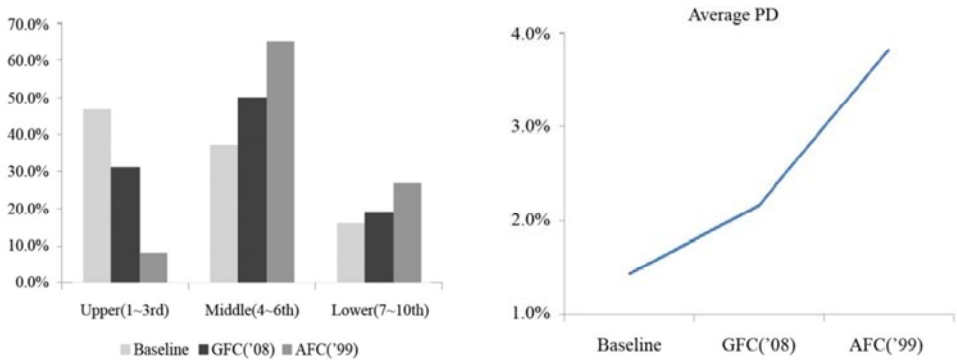

FIGURE 9. DISTRIBUTION OF BORROWERS AND AVERAGE PD BY STRESS SCENARIOS

PD would increase if macroeconomic conditions shifted from the baseline to each historical stress scenario of the past crises.

It is important to note that the stress test proposed in the current study utilizes the dominant predictor of default as an intermediary channel through which macroeconomic conditions affect the aggregate default rates. In the current stress test model, macroeconomic shocks affect the aggregate credit risk level by changing the distribution of borrowers in terms of their credit scores, which is strongly associated with default rates. In addition, we take into account the impact of macroeconomic fluctuations on the relationship between the over-indebtedness indicator and default rates. These features of the current stress test are distinct from those in previous studies that also used CB data but that utilized DSRs or LTVs as main variables in the intermediary channels, which are weakly associated with default rates, as noted in Table 3.[13] Changes in such indicators as the DSR, LTI, or LTV may only partially explain the movement of default rates, as also noted by Kim and Byun (2010) and Choi and Park (2015), among others. In addition, the relationship between such over-indebtedness indicators and subsequent defaults was often assumed to be constant in earlier work, thus resulting in little change in the default rate even against severely depressed macroeconomic scenarios.

VII. Concluding Remarks

The current study assesses the vulnerability of household debt based on an analysis of obligor-level information from CB data. We construct over-indebtedness indicators from the CB data and assess their capability to predict defaults in the near future. Based on the over-indebtedness indicators, we show how borrowers are distributed in terms of over-indebtedness and how the over-indebted differ from average borrowers in terms of their characteristics. Furthermore, we conduct a stress test on household debt to assess the vulnerability

[13]See Hahm, Kim, and Lee (2010); Kim, Chang, and Choi (2012); and Lee, Jun, Chung, and Byun (2014) for related previous studies which use CB data.

of borrowers and show how the aggregate credit risk would change under severe macroeconomic distress such as that which occurred in past financial crises (the global financial crisis and the Asian financial crisis).

The global financial crisis of 2007-09 has shown us that the rising risk from burgeoning household debt in association with real estate bubbles could damage the financial stability with large disruptions to the real economy and lead therefore to extensive reforms on financial regulation and supervision afterwards. In order to implement effective policies for financial stability, the collection of information and the monitoring of emerging risks have been strongly emphasized, as discussed above. The findings of this paper may contribute to this end by providing practical guidance for assessing the vulnerability associated with the burgeoning household debt, which is a notable risk factor challenging the financial stability of Korea. It is also important to note that the CB data used in the current study cover most individuals and financial institutions in Korea and are updated at a high frequency, thus allowing for timely assessments of credit risks. The stress test conducted based on the CB data allows us to assess the household credit risk at the aggregate level while showing us how the distribution of obligors in terms of the credit risk (ratings) dimension would change in response to adverse shocks. This distributional feature of the risk assessment measure may also help financial institutions with the retail loans in their portfolios for their risk management practices.

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The Effects of Financial Support Policies on Corporate Decisions by SMEs†

By CHANGWOO NAM*

This paper investigates the effectiveness of public credit guarantee programs and interest-support programs for SMEs (small and medium enterprises). First, assuming that there is an imperfect information structure in the SME loan market, we analyze how SME support financial programs affect the corporate decisions made by SMEs with regard to default or loan sizes. In addition, this paper theoretically computes the optimal levels of credit guarantee amounts and the interest-support spread under equilibrium with imperfect information in a competitive loan market. Second, the paper empirically analyzes the continuous policy-treatment effect with the GPS (generalized propensity score) method. In particular, we consider the ratio of guaranteed debt to the total debt as a continuous policy treatment. The empirical results show that marginal effects of a credit guarantee on SMEs' productivity, profitability, and growth potential decrease with the ratio of guaranteed debt to the total debt. In addition, the average effect of a credit guarantee is maximized when this ratio is at 50% to 60%.

Key Word: Small and Medium-sized Enterprises,
Information Asymmetry, Loan Market,
Credit Guarantee, Generalized Propensity Score
JEL Code: G14, G18, G21, G28

I. Introduction

Recently, concerns have been raised that financial support for Korean SMEs could delay the restructuring of SMEs and reduce the productivity of the Korean economy overall despite the fact that the Constitution of the Republic of

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† This paper is based upon "The Effectiveness Analysis of Financial Support Policies for SMEs: Based on Firm's Dynamic Decision Model," Korea Development Institute, 2014, by the same author.

Korea guarantees the incubation and training of Korean SMEs. The magnitude of public credit guarantees compared to the economy is relatively large as compared to that in other OECD countries. In fact, this issue has been constantly raised, but studies that develop theoretical models about financial friction in relation to limited financing for SMEs are rare. Stiglitz and Weiss (1981) in their partial equilibrium model for the SME loan market show the possibility of the existence of credit allocation (credit rationing). However, they do not provide a balanced model of the SME lending market which reflects policy effects on production and consumption in the economy to provide a basis that can be comprehensively judged by policymakers.

This study seeks to identify the optimal level of SME financing policies by simulating how corporate decision-making, including bankruptcy decisions and bank interest rate decisions in the lending market, can be affected by government finance policies (public credit guarantees and interest support) in the presence of information asymmetry (imperfect information) for SMEs, a balanced fiscal policy, and the current tax regime. The basic purpose of SME financing policies is to alleviate financial friction and information asymmetry in order to realize the optimal allocation of resources. However, the existence of information asymmetry leads to a fiscal policy under which consumers should pay taxes for SMEs. Therefore, the amount of macro-financial assistance is determined at an optimal combination of income taxes paid by small businesses and final consumers.

Moreover, this paper proposes more concrete policy measures to improve the credit guarantee policy on the basis of the presented theoretical discussions, in particular by analyzing the performances of SMEs. Previous studies focused mainly on exposure (or a lack of it) to credit guarantee policies rather than on the impact of the degree of exposure to a credit guarantee policy on the performance of SMEs. Therefore, this study empirically examines the effect of the ratio of guaranteed debt to total debt as a continuous policy treatment on SME performance outcomes with the GPS (generalized propensity score) method.

Briefly, this study finds the following. The equilibrium model based on bankruptcy and lending decisions by SMEs shows that the current scale of public credit guarantees is higher than the optimal level of policies because the social cost is beyond the optimal level. This suggests that the government should consider gradually reducing the amount of public credit guarantees to maximize social welfare. Second, the performance analysis shows that the marginal effect of credit guarantee policies on their ratio of credit-guaranteed debt to total debt is decreasing for the SMEs. This suggests that it is necessary to limit the ratio of credit guarantees to total debt.

This paper is organized as follows. Section II describes the current status of SME financial support programs and compares with other country-specific financial support policies in OECD countries. Section III develops the equilibrium model and Section IV conducts a social welfare analysis in accordance with a counter-factual economic model. Section V analyzes the policy effects of the credit guarantee program with firm-level data. Finally, Section VI proposes directions for improvement of the financial support programs for SMEs.

II. SME Financial Policies

A. *Credit Guarantee and Interest Support Programs*

A public credit guarantee refers to a type of financial support program for SMEs that deficient collateral capacity. It offers a warranty for payment to banks upon the bankruptcy or liquidation of the borrowing SME. This program offers a higher degree of financial support in terms of the policy scale, and is typical of SME financing policies. Regarding the overall amount of Korean public credit guarantee funding, KODIT,[1] KIBO,[2] and KOREG[3] recorded a value of 75.5 trillion won at the end of 2013, which accounted for 15.4% of all SME loans, i.e., 488.9 trillion won. Compared with 2007 before the global financial crisis, the magnitude of public credit guarantees has increased by nearly 70%.

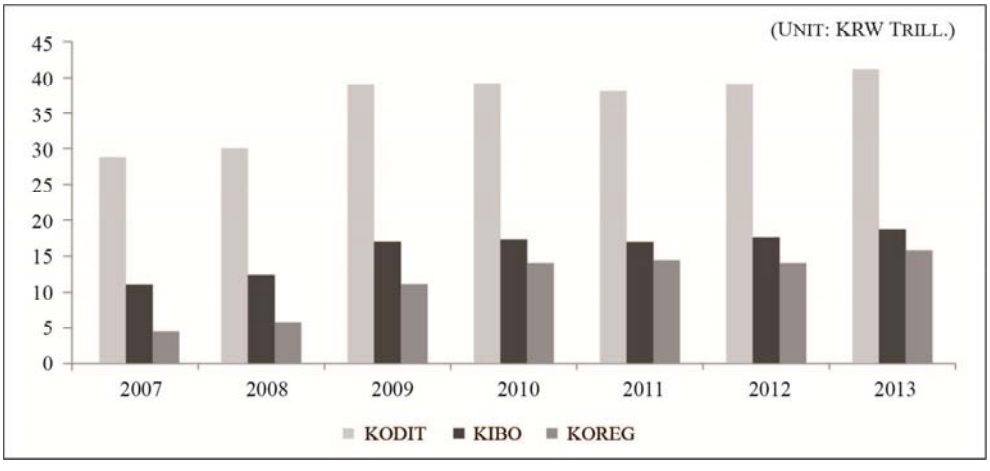

FIGURE 1. TRENDS OF CREDIT GUARANTEE PROGRAMS

TABLE 1—CREDITS FOR INDEMNIFICATION OF KODIT (UNIT: KRW TRILL., %)

	2007	2008	2009	2010	2011	2012	2013
Guaranteed Loans	28.5	30.3	39.2	38.7	38.4	39.2	40.6
Subrogation	1.1	1.4	1.8	1.8	1.9	1.9	1.7
(Subrogation Rate)	4.1	4.8	4.7	4.7	5.0	4.9	4.4
Indemnity	3.1	3.1	3.0	2.9	3.0	3.2	3.3
Recovered Credits	0.6	0.6	0.7	0.8	0.7	0.7	0.5
Recovery Rate	17.2	18.8	23.1	25.8	24.5	23.5	16.6

Note: we use average amounts per year to compute the subrogation and recovery rates.

[1]Korea Credit Guarantee Fund.
[2]Korea Technology Finance Corporation.
[3]Korea Federation of Credit Guarantee Foundations.

Credit guarantee institutions subrogate payments of bankrupt SMEs to lenders and obtain indemnity from bankrupt SMEs. KODIT (the Korea Credit Guarantee Fund) had an average subrogation rate of 4.7% from 2007 to 2013, and its average recovery rate of credits for indemnification was 21.3%. However, the average recovery rate of commercial banks is approximately 25%, implying that public credit guarantee funds may be more benevolent than the private sector.

An interest-support program means a financial policy that provides SMEs with a portion of the loan interest rate when SMEs satisfy certain conditions. The interest-support program is typically implemented by local governments. For example, the Seoul metropolitan government provided SMEs with interest support from 1%p to 3%p according to CD rates, and most local governments, such as the Busan metropolitan government, offer supports ranging from 2%p to 5%p of the interest spread when they lend working capital to small businesses. In recent years, the fiscal expenditures of the central government for interest-support programs appear to be expanding, but it is not clear whether small business loans are actually supported.

In addition, KIBO, as a technology credit bureau, operates an interest-support program offering up to 3%p for credit loans issued. In particular, the 2014 budget of KIBO for this program was 3.75 billion won, which supported nearly 375 billion won of credit loans at an average interest-support spread of 1%p.

B. *Policy Comparison with OECD Countries*

Given the recent doubts about the economic efficiency of public credit guarantee programs, how much the government should provide public credit guarantees is at issue. In particular, the amounts of SME loans and the sizes of public credit guarantees in OECD countries are very important references in setting policy goals.

Table 2 shows the proportions of SME loans to all business loans in OECD countries. Korean SME loans held a ratio of 74.7% of all business loans in 2012, but exceeded 80% of all business loans from 2007 to 2010. SME loans in most other countries accounted for no more than 50% of all business loans, except in Portugal and Switzerland, which recorded similar levels, at 74.7% and 78.8%, respectively, in 2012.

TABLE 2—SME LOANS OVER BUSINESS LOANS OF OECD COUNTRIES (UNIT: %)

	2007	2008	2009	2010	2011	2012
Canada	17.4	15.6	17.9	17.5	17.5	15.7
Chile	16.7	15.2	17.5	18.2	18.2	18.5
Korea	**86.8**	**82.6**	**83.5**	**81.5**	77.7	74.7
Mexico	13.0	12.3	12.0	13.0	13.4	16.1
Portugal	78.3	77.7	77.4	77.3	77.1	74.7
Switzerland	81.2	81.1	80.1	79.9	78.8	78.8
Turkey	40.1	33.8	31.7	35.6	35.7	37.5
U.K.	19.6	18.0	19.9	21.2	21.2	21.8
U.S.A	30.1	27.7	27.6	29.0	26.5	23.7

Note: OECD, Financing SMEs and Entrepreneurs 2014: An OECD Scoreboard.

TABLE 3—GUARANTEED LOANS OVER SME LOANS OF OECD COUNTRIES (UNIT: %)

	2007	2008	2009	2010	2011	2012
Canada	1.4	1.6	1.4	1.5	1.4	1.3
Chile	3.0	2.5	6.5	10.2	9.4	15.0
Finland	3.6	3.7	4.8	5.4	6.3	5.2
Korea	**12.0**	**11.6**	**15.2**	**16.0**	**15.4**	**15.4**
Mexico	0.9	1.1	1.9	1.9	2.1	1.6
Netherland	1.9	1.9	2.4	5.7	5.4	3.2
Portugal	0.9	1.7	5.3	7.4	6.9	7.2
Spain	1.3	2.0	2.2	3.4	4.3	3.4
Switzerland	0.0	0.0	0.1	0.1	0.1	0.1
Turkey	0.1	0.3	0.7	0.7	0.7	0.6
U.K.	0.2	0.2	0.7	0.5	0.3	0.3
U.S.A.	3.0	2.3	2.2	3.4	3.1	3.9

Note: OECD, Financing SMEs and Entrepreneurs 2014: An OECD Scoreboard.

Table 3 shows the shares of SME loan credit guarantees of SME loans in OECD countries. The table also shows that Korea has the highest proportion of public credit guarantees of SME loans among OECD countries, followed by Chile, which guaranteed 15% of SME loans during the financial crisis.

However, Portugal and Switzerland recorded SME loans of 7.2% and 0.1%, respectively, at the end of 2012, showing relatively low rates compared to Korea. In addition, the table confirms that public credit guarantees in developed countries such as Canada, the United Kingdom, and the United States do not exceed 5% of all SME loans. Additionally, in developed countries, capital market financing is more accessible than loan financing to SMEs; therefore, the intervention of those governments in the loan markets is less likely to occur than in other countries.

III. Equilibrium in the SME Loan Market

A. *Theory of the SME Loan Market with Imperfect Information*

It has been a long time since the issue of corporate lending through financial intermediation, particularly the SME loan market with information asymmetry, has been discussed. Stiglitz and Weiss (1981) reviewed the possibility of credit rationing as provoked by information asymmetry. Specifically, a single collateral rate and a single loan interest rate in the economy were presented in their work and demonstrated to distort the supply curve of SME loans through adverse selection, with the supply curve not increasing monotonically at a certain level of interest rate because SMEs with riskless businesses exit the loan market, whereas SMEs running risky businesses stay in the loan market. Eventually banks offer interest rates that lead to excess demand such that the credit rationing by the government can exist.

This argument is quite controversial, and the possibility of credit rationing is refuted in various papers as a result. Typical papers taking this line include Bester (1986) and Arnold and Riley (2009). Bester (1986) argued that credit assignment occurred in the economy of Stiglitz and Weiss (1981) due to the single-rate loans with a single collateral rate. If a bank can offer various sets of collateral rates and

interest rates, the loan market can identify the riskiness of medium and small businesses and thus provide several financial products with different collateral rates and interest rates. In other words, the bank can offer a wide range of loan products with collateral requirements and interest rates and can force companies to reveal their degree of riskiness.

Arnold and Riley (2009) contradict the possible existence of credit rationing using the same economic structure in Stiglitz and Weiss (1981). They argue that credit rationing may occur due to a disconnection in the demand curve rather than an issue with the supply curve, as in Stiglitz and Weiss (1981). If the bank revenue curve shows a U-shaped line, the optimal curve of the interest rate to maximize a bank's expected return will have a disconnection at a certain point, at which demand exceeds supply. However, Arnold and Riley (2009) explain that this is likely to arise only when the tail distribution of a low-risk business is extremely low. Recently, however, Nam (2013a) demonstrated the possibility of credit rationing by calculating invariant measures of SMEs in a repeated game structure with bankruptcy decisions. We essentially utilize the economic structure of Stiglitz and Weiss (1981). However, there are different assumptions in this paper as compared to those in Stiglitz and Weiss (1981). The government presents a credit guarantee ratio for a loan, and banks offer optimal interest rates to SMEs under imperfect information. Moreover, SMEs determine loan amounts depending on their productivity states, unlike in Stiglitz and Weiss (1981). Finally, the government finances the social cost of credit guarantees by collecting tax on consumer deposits.

Decision-making by firms in the economy is much more complicated than the consumer decision-making structure. First, a company should optimize its capital and investment amounts depending on the current investment opportunities and should determine wage prices and other production factors, including labor. In addition, firms should allocate net income optimally into investments and dividends according to economic conditions and should determine their use of indirect financing or direct financing in their corporate finance strategy.

A few studies of corporate decisions about investments, dividends and corporate finance use the Bellman equation (or dynamic programming). Although many recent studies are notable, the present study mainly references Hopenhayn (1992); Chatterjee, Corbae, Nakajima, and Rios-Rull (2007); and Arellano, Bai, and Zhang (2012).

Hopenhayn (1992) for the first time showed that there is a general equilibrium state in a firm's entry-exit structure, proving that there is an invariant distribution of companies in the market in accordance with the steady-state rates of entry and exit. Chatterjee, Corbae, Nakajima, and Rios-Rull (2007) analyze the social welfare effects of changes in credit policies, especially when consumer debt is not fully guaranteed. Arellano, Bai, and Zhang (2012) demonstrate that the corporate development growth rate and the size of indirect financing depend on the degrees of financial development and financial friction. However, a firm's exit value in the economy is exogenous, as in Hopenhayn (1992), and a firm automatically exits from the market if its operational value is lower than its exit value.

This paper utilizes the economic structure of Hopenhayn (1992). A firm decides to exit the market through bankruptcy, or it can borrow money from a bank, as in

Arellano, Bai, and Zhang (2012). There are financial friction and an adjustment cost of indirect financing. Finally, the present paper calculates the invariant distribution of firm bankruptcies with partial government guarantees, as in Chatterjee, Corbae, Nakajima, and Rios-Rull (2007). One of the main features in this model is that the liquidation value of a SME is determined endogenously such that the firm's entry rate is determined in accordance with its bankruptcy decision.

This paper refers to many earlier works about such dynamic decision models. Zhang (2005); Cooper (2006); Li, Livdan, and Zhang (2009); Nikolov and Whited (2009); and Livdan, Sapriza and Zhang (2009) are the main references. These all posit various dynamic designs of corporate decision models and discuss how they are affected by uncertainty in the macro-economy or by idiosyncratic shocks. In addition, these corporate decision models analyze the impact of dividends and investments on stock prices. In particular, Clementi and Hopenhayn (2006) explain firm dynamics in terms of dynamic contract theory under information asymmetry. However, they do not take into account the distribution of equilibrium among firms, and they only focus on the optimal conditions of contract theory. Moreover, they do not endogenously address market entry and exit rates.

The present model is in general new compared to those in earlier works. First, capital from the loan market is included as an operating profit function, whereas most existing dynamic models separate external financing from production capital. Second, our model endogenizes the default decisions of SMEs such that the firm's default value and entry value are endogenously determined by a default decision. Third, in our model, the government intervenes in the loan market with financial policies such as credit guarantee and interest support programs. Earlier firm dynamic models only consider the relationships between firms and financial institutions and not policy-intervention efforts by the government. Thus, our model is unique in terms of proactive policy intervention in the loan market.

B. *SME Dynamic Decision Model*

1. *Operating Firms' Decisions*

The structure of corporate cash flow is defined as

$$\pi(k=1,b,z)=z(k+b)^{\alpha},$$

in which π is the operating income function with z as the exogenous shock unknown to the government and banks, as they know only the transition probability of z. Moreover, z is defined as a first-order autoregressive process with i.i.d. shock. In this model, z is the main factor to induce information asymmetry into the SME loan market, k is the capital normalized to one, and b is the loan size that is also the leverage ratio owing to the normalization of capital. α is a parameter which denotes the capital share in the operating income function.

The current operating company, that is, a SME without a credit history of default has dynamic decisions defined as

$$V(b,z,h=0)=max\begin{cases}V_0(b,z,0)\equiv \max\limits_{b'\in[0,\bar{b}]} C_0+\beta E\left[V(b',z',0)|z\right],\\ V_1(b,z,0)\equiv C_1+\beta E\left[V(0,z',1)|z\right]\end{cases}$$

in which h is defined as the credit history of default, taking a value of one if a SME became bankrupt in the last period or zero otherwise. C_0 and C_1 are defined as

$$C_0\equiv\pi-(q-r_g)b+(1-f_c c)b'-\Phi(b',b), C_1\equiv\pi 1_{\{z<0\}}-rc\cdot cb.$$

Here, V is the value of the operating company without a history of default, and V_0 is the value of the operating company when it decides not to default and obtain a new loan, b', from a bank. V_1 is the firm value of a company that decides to declare bankruptcy at the present time. q and r_g are defined as the loan interest rate and the interest-support spread, respectively. f_c and c are the credit guarantee fee to pay to the government and the coverage ratio of the credit guarantee, respectively. β is a parameter pertaining to the time discount preference, and rc is the recovery rate of credit for indemnification after subrogation of the government instead of a bankrupt company. $E[\cdot|z]$ is the expectation operator given z. $\Phi(b',b)$ is defined as the adjustment cost function of the loan size, as in the equation

$$\Phi(b',b)=\frac{\gamma}{2}(b',b)^2 1_{\{(b'-b)>0\}},$$

in which γ is a parameter linked to the adjustment cost function, which is represented as a quadratic function in order to prevent companies from borrowing money excessively. Finally, $1_{\{\cdot\}}$ is an indicator function having a value of 1 if the statement in $\{\}$ is true, and 0 otherwise. The symbol $'$ over the variables denotes the next period. If the firm decides not to declare bankruptcy at the present time, it chooses the optimal size of a loan for the next period, b', in $[0,\bar{b}]$, and proceeds to the decision of the next period, $V(b',z',0)$. Moreover, if the firm decides to declare bankruptcy at the present time, the firm should give any positive operating income to the lender and pay the recovery rate multiplied by the guaranteed loan size (cb) back to the government, and then move to $(0,z',1)$.

2. *Bankrupt Firms' Decisions*

A firm with a history of default $(h=1)$ has the following decision structure,

$$V(0,z,1) = 0 + \beta(1-\delta)E\left[(1-\lambda)V(0,z',0) + \lambda V(0,z',1) \mid z\right],$$

in which δ is defined as the probability of business liquidation after bankruptcy, and δ is exogenously given. The cash flow of a bankrupt firm is zero because it is assumed to have ended its operations during its default history. λ is defined as the probability of bank account suspension of the bankrupt company, which is 1/2. In this model, the business liquidation rate and the duration of the default history determine endogenously firms' default rates and entry rates into the market. Therefore, if a firm decides to default in the previous period, then in the present period, the firm stays in the market without operating income, and the firm will close its business in the next period with the probability of δ or will stay in the market with or without a history of default according to the probability[4] of λ.

3. *Firms' Invariant Distributions*

The state-mapping function is defined according to the state variable vector, (b,z,h),

$$H(b,z,h'=0) = \begin{cases} 1 & if\ d=0\ and\ h=0 \\ \gamma & if\ h=1 \\ 1 & if\ d=1 \end{cases}$$

$$H(b,z,h'=1) = \begin{cases} 1 & if\ d=o\ and\ h=0 \\ 1-\lambda & if\ h=1 \\ 1 & if\ d=1 \end{cases}$$

and the transition function of corporate policy is defined as

(1) $$G(b,z,h'=0,S) = \oint_z 1_{\{b' \in B\}} H(b,z,h'=0) f(dz' \mid z)$$

(2) $$G(b,z,h'=1,S) = \oint_z 1_{\{b \in B\}} H(b,z,h=0,h'=1) f(dz' \mid z)$$
$$\text{or } \oint_z 1_{\{b=0\}} H(b,z,h=1,h'=1) f(dz' \mid z)$$

in which $\oint$ is the matrix operator, $f(\cdot \mid \cdot)$ is the transition probability of z, and S is defined as the compact space of the state variables. The entry function of newborn firms is defined as

[4]The Korea Financial Telecommunications & Clearings Institute (KFTCI) suspends the checking acco unts of bankrupt companies for investor protection for at least two years.

(3) $$N(b,z,1,S) = \oint_z 1_{\{(b',h')=(0,0)\}} g(dz'),$$

in which $g(\cdot)$ is the probability of z. Finally, we define the transition function of firm as

$$G^*(b,z,S) = \delta(1-\lambda)N(b,z,1,S) + G(b,z,0,S) + (1-\delta)G(b,z,1,S).$$

Additionally, given (c,q), the distribution of the corporate state-vector (b,z,h), μ, is defined using the operator Y as

$$\left(Y_{(c,q)}\mu\right)(B \times Z) = \int G^*(b,z,S)\,d\mu$$

Theorem (Unique Existence of Invariant Measure)
For any $(c,q) \in C \times Q$, and for the measurable selection of the optimal policy correspondence, the unique and invariant distribution $\mu_{(c,q)} \in M(B \times Z, S)$ satisfies $\mu_{(c,q)} = Y_{(c,q)}\mu_{(c,q)}$..

Proof
We use the proof of Theorem 2 in Chatterjee, Corbae, Nakajima, and Rios-Rull (2007).

Here, μ represents the distribution of default decisions as well as the loan sizes according to the state variables. Specially, μ is defined as banks' belief function with respect to d, b, and z, which are information inaccessible to banks. We can then calculate the default probability and the conditional default probability $\mu(d=1)$ and $\mu(d=1|b)$ (Athreya, Tam, and Young 2012). In addition, $\mu(d=1)$ is the point-mass value because d represents a discrete choice of the default decision.

C. *Equilibrium Model*

1. *Bank Interest Rates for Loans*

Banks do not know the state of the SME, z, but are only aware of the transition probability of z. A SME is assumed to repay the bank loan, b, after which it will come back to the bank with a new contract, b'. However, the bank does not

know the SME's history, which means that the bank only knows b', not b.[5] This is the mechanism of information asymmetry. Therefore, if banks have a cumulative distribution function of belief (μ) regarding a SMEs' decision to declare bankruptcy, the expected return of b is then defined as

$$\int_b R_b(c,q_b|b)b\mu(db)-i\int_b b\mu(db),$$

in which i denotes the bank's financing cost, that is, the interest rate of consumer deposits, which is assumed to be exogenously given, and $\mu(db)$ is the probability density function with respect to b. We then define the expected return of b apart from the financing cost as follows:

$$R_b(c,q_b|b)=r(c,q_b|b)-\underbrace{f_g}_{bank's\ contribution\ rate\ of\ credit\ guarantee\ fund},$$

We assume that the loan market is perfectly competitive such that the expected profit is defined as

$$r(c,q_b|b)-f_g-i\equiv 0, \tag{4}$$

where q_b is the interest rate for b. Additionally, r is defined as

$$r(c,q_b|b)\equiv\underbrace{\int_Z q_b\mu(d=0,\,dz|b)}_{expected\ profit\ of\ non-default}+\underbrace{\int_Z\left[\frac{\pi(b,z)1_{\{z>0\}}}{b}+c\right]\mu(d=1,dz\,|\,b)}_{expected\ profit\ of\ default}, \tag{5}$$

in which $\mu(d=0,dz\,|\,b)$ and $\mu(d=1,dz\,|\,b)$ are the conditional beliefs in non-default or default with respect to z given b, and q_b is expressed with (4) and (5) in the following form:

$$q_b=\frac{i-\left[\int_Z\left[\frac{\pi(b,z)1_{\{z>0\}}}{b}+c\right]\mu(d=1,dz|b)-f_g\right]}{\mu(d=0\,|\,b)}>i+f_g. \tag{6}$$

The determination of equation (6) is due to information asymmetry, i.e., how much the banks believe that the loan (b) which pays interest (q_b) will default. Particularly, if banks believe that $\mu(d=1|b)=0$, $q_b=i+f_g$ is obvious.

[5]The bank actually can review the history of SME loan cases. Our assumption can then be stronger than reality but can also be interpreted to be a gadget creating information asymmetry that can be update d in the model.

2. *Government's Balanced Budget*

In this paper, the government supports SMEs' financing operations through a public credit guarantee fund and an interest support program, and the government finances this from SMEs' guarantee fees, banks' contributions, and consumer taxes. Thus, the government's budget constraint is defined as

$$\int_b \left[f_g - \underbrace{\left[r_g + (1-rc)c\mu(d=1|b) \right]}_{government\ expenditure} \right] b\mu(b) \geq (-f_c c - \tau)\int_{b'} b'\mu(b'),$$

in which τ is the tax rate. If the government balances its budget, the tax rate is then solved, as follows:

$$\left[r_g + (1-rc)c\frac{\int_b b\mu(d=1,b)}{\int_{b'} b'\mu(b')} \right] - (f_g + f_c c) = \tau.$$

It is important to note that τ increases with the interest-support spread and the ratio of default loans to all loans. The ratio of default loans to all loans is determined endogenously by the proposed decision model for firms. The guarantee coverage ratio simultaneously affects government expenditures for SME policies and government revenue, but it impacts tax rates differently according to the ratio of default loans to all loans. In addition, τ decreases with the recovery rate of default loans, the bank contribution rate, and the credit-guarantee fee rate. However, the bank contribution rate can negatively influence the total credit size in the economy, and the credit-guarantee fee rate can affect the operating cash flow of SMEs such that corporate decisions may be distorted. Moreover, the recovery rate can affect the value of a bankrupted company.

3. *Consumer Utility Problem*

This paper assumes one representative consumer in the economy, with the following utility problem:

$$v(D,B) = \max_C U(C) + \beta E[v(D',B')].$$

She has a budget constraint for consumption (C) which is defined as

$$C = D + iB - (1+\tau)B',$$

in which D, B, and B', are the dividend, deposit at the previous time, and new deposit for the next period, respectively, in the aggregation such that

$$\mathrm{D} \equiv \int_{z,b} \left[C_0 1_{\{d=0\}} + C_1 1_{\{d=1\}} \right] \mu(dz, db),$$
$$\mathrm{B} \equiv \int_b b\mu(db), \text{ and } B' \equiv \int_{b'} b'\mu(db').$$

In particular, B and B' should be equal according to the market clearing condition.

4. *Bayesian Equilibrium*

The Bayesian equilibrium is defined under information asymmetry as in Athreya, Tam, and Young (2012). Athreya, Tam, and Young (2012) simultaneously analyze equilibrium without information asymmetry, but our study focuses on the optimal levels of financial support in equilibrium with imperfect information.

Definition (Bayesian Equilibrium)
The Bayesian equilibrium in the SME loan market consists of (a) a SME's loan $\mathrm{b}^*: \mathrm{S} \to \mathrm{R}$ and default decision $\mathrm{d}^*: \mathrm{S} \to 0,1$, (b) a bank's loan interest rate $\mathrm{q}^*: \mathrm{R} \times \mathrm{M} \to [1/\beta, \bar{\mathrm{q}}] \equiv \mathrm{Q}$, (c) the government's budget $\tau^*: \mathrm{R} \times \mathrm{Q} \to [0,1)$, and (d) given μ^*, the bank's belief about the SME's loan and bankrupt decision satisfies the following:

① SME solves the optimization problem of b^* and d^* given $\mathrm{q}_{\mathrm{b}}^*$.

② Banks offer $\mathrm{q}_{\mathrm{b}}^*$ as a mixed Nash equilibrium under the price competition given the SME's b^* and their belief about default, $\mu^*(\mathrm{b})$.

③ The government balances the budget by adjusting τ^* given b^*, b^*, q^*, and μ^*.

④ A consumer chooses C^* and B^* given $\mathrm{b}^*, \mathrm{q}^*$, μ^* and τ^*.

D. *Assessment of the Equilibrium Model*

In this chapter, we define the SME dynamic model, bank interest rate decisions, the government's balanced budget constraint, and the abovementioned consumer's utility problem. However, discussion about the compatibility of the model is critical for the interpretation of the policy evaluation. For this reason, we assess the equilibrium model and limit the analysis of the effects of SME financial policies through a simulation.

SMEs' optimal decisions are the most important element in the model. In particular, SMEs in the model are assumed to use only loans rather than the capital markets. This is reasonable because the majority of SMEs as unlisted companies

use indirect financing according to available data. However, while analyzing the substitutability of direct financing and indirect financing via the financial markets, it becomes necessary to expand our model to the capital markets. Second, in this model, we do not take into account the industrial characteristics of SMEs. Our model basically assumes that SMEs are only manufacturing companies.

Third, the model assumes a perfectly competitive market for loans, which is not unrealistic. The paper seeks to analyze the interactions between SMEs' dynamic decisions and bank interest rates such that perfect competition is not decisive with regard to theoretical results. The fourth important feature of our model is that macro-prudential measures such as the capital ratio to risk assets are not regulated explicitly because reductions of credit amounts by banks are endogenously adjusted through changes in the loan interest rates.

Fifth, one representative consumer is assumed to own all of the SMEs. However, the SMEs' optimal decisions are determined individually. In other words, the final consumer exists only for the social welfare analysis with the economic variables of the SMEs, the banks, and the government. Finally, a price structure for the products made by the SMEs is not present because the purpose of our study is to analyze steady states of firm distributions according to market entry and exit rates rather than dynamic transitions of product prices induced by unexpected shocks.

IV. Policy Simulation in the Equilibrium Model

A. *Simulation Methodology*

1. *Computational Methodology*

The heterogeneous agent model in a state of Bayesian equilibrium can be computed by dynamic programming. Specifically, the discrete decision model regarding the bankruptcy decision can be processed by the methodology in Adda

TABLE 4—SIMULATION PARAMETERS AND RESULTS

	Parameters	Values
α	Income share of capital	0.33
β	Annual time discount rate	0.98
γ	Adjustment cost	4.0-5.5
δ	Liquidation of bankrupt SMEs	70%
$1/\lambda$	Average duration of default history record	2
c	Average guarantee coverage ratio	90%
i	Average gross deposit rate	4.1%
f_c	Average credit guarantee fee rate	1.1%
f_g	Bank's contribution rate to credit guarantee fund	0.38%
rc	Average recovery rate from indemnity	20%
Leverage ratio	Base model ($r_g \in [0.000,0.045]$)	141%~308%
Leverage ratio	KIS DB from FY 2000 to FY 2011	293%~356%
Subrogation rate	Base model ($r_g \in [0.000,0.045]$)	4.68%~10.6%
Subrogation rate	KODIT from FY 2007 to FY 2013	4.1%~5.0%

and Cooper (2003).[6]

2. *Simulation Parameters*

Before simulating the model, the outside parameters should be determined given the economic context. Table 4 shows the values of the outside parameters. First, we use the value from Park (2012), 0.33, as the capital income-share parameter. The annual time discount rate is set to 0.98 considering the interest spread and deposit rate, and the parameter of adjustment cost function used ranges from 4.0 to 5.5 such that we have a range of simulation results. The business liquidation rate of bankrupt SMEs is 70%, close to the value in Kang (2004), and the average coverage ratio is assumed to be 90%. The average deposit rate is the average interest rate for new deposits from 2001 to 2013 at BOK ECOS, 4.1%. The average credit guarantee fee rate and recovery rate of indemnity are respectively 1.1% and 20% from Kang (2005). The bank contribution rate to credit guarantee funding takes its value from Kang et al. (2014), i.e., 0.38%. The interest-support spread ranges from 0.00% to 4.5% at increments of 50bp. We then calculate the average leverage ratios and average subrogation rates.

In Table 4, the leverage ratio of manufacturing SMEs from FY 2000 to FY 2011 in the KIS database ranges from 293% to 356%. From the simulation, we compute the leverage ratio of SMEs as ranging from 141% to 308%. In addition, the KODIT subrogation rate from FY 2007 to FY 2013 has a range of 4.1% to 5.0%, and the range of the simulation is from 4.68% to 10.6% according to the range of the interest-support spread.

B. *Equilibrium Simulation Results*

1. *Consumer Social Welfare in Equilibrium*

Figure 2 shows the final tax rate of the consumer according to the coverage ratio and interest-support spread under a state of equilibrium. Without the credit guarantee and interest support programs, the consumer will have the lowest tax rate, SMEs will not pay the guarantee fee, and banks will not contribute to public guarantee funds. In addition, the tax rate does not increase monotonically, as it shows an $\cap$-shaped decline at a coverage ratio of 60%. The growth of bankruptcies by SMEs responds non-monotonically to policy variables and the expected return curves of banks.

Figure 3 shows the final consumption at equilibrium according to the coverage ratio and the interest-support spread. Surprisingly, consumption is relatively high at lower coverage ratios and interest-support spreads. This phenomenon has two causes. First, the higher guarantee coverage ratio and interest-support spread imply

[6]Find steps of computation in the appendix.

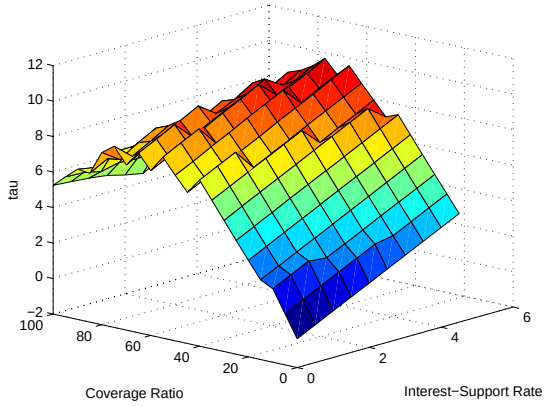

FIGURE 2. CONSUMER'S TAX RATE AT EQUILIBRIUM (UNIT: %)

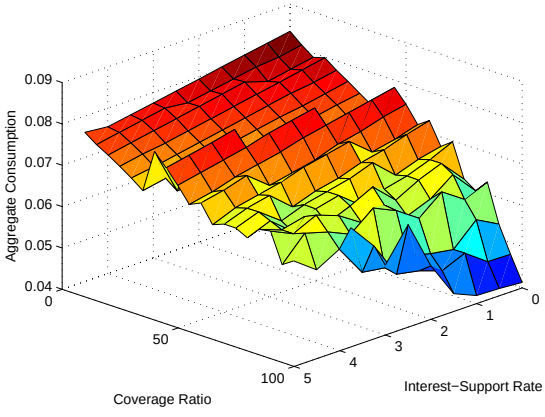

FIGURE 3. CONSUMPTION AT EQUILIBRIUM (UNIT: %)

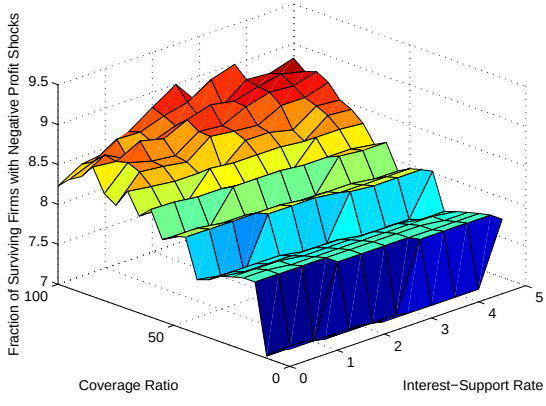

FIGURE 4. PROPORTIONS OF MARGINAL SMES IN THE MARKET AT EQUILIBRIUM (UNIT: %)

that the consumer pays more tax, causing the social cost to increase. Second, marginal SMEs' exits are delayed due to the credit guarantee and interest support programs, after which the overall profitability of the economy declines as the proportion of marginal companies in the market increases. In Figure 4, although the proportion of operating SMEs with negative profits is close to 7% at equilibrium without the credit guarantee and interest support programs, the proportion at equilibrium with these programs increases to 9%. Finally, the figure shows that the financial policies for SMEs are likely to delay marginal SMEs' exits and thus lower the final consumption rate. Thus, the cost of the risk sharing of SMEs' bankruptcy and the reduced consumption level can lower social welfare overall.

2. *Optimal Levels of Financial Support and Social Welfare*

Table 5 shows the optimal levels of the coverage ratio and interest-support spread according to the recovery rates of indemnity as determined via a simulation.[7] The optimal levels of the coverage ratio at recovery rates of 20% and 40% are 8.75% and 16.25%, respectively. It should also be noted that the optimal coverage ratio is in fact the proportion of the credit guarantee size out of the overall SME loan amounts. The optimal rates of the interest-support spread at recovery rates of 20% and 40% are 25bp and 50bp, respectively. Finally, an increase in the recovery rate of indemnity can alleviate the government's budget line such that it can then expand credit guarantees and interest-support spreads due to the reduced social cost.

Our equilibrium model takes into account the decision-making activities of several economic agents. The coverage ratio of credit guarantees directly influences the decisions of firms, banks, and the government, and the interest-support program directly impacts corporate and governmental decisions, whereas it indirectly affects banks' decisions. However, a clear analysis of the effects of SME financial policies on the real economy is not easy to conduct. The present paper can only foretell the impacts of financial polices by computing social welfare through the distribution of leverage and the default probability via the behavior of economic agents.

In this context, the recovery rate of indemnity critically affects the economy. The recovery rate influences the value of bankrupt companies such that a higher recovery rate increases the debt burdens of firms. Simultaneously, a higher recovery rate of indemnity positively affects social welfare because the social cost decreases. Table 6 shows the effect of the increase of the recovery rate on consumption, the tax rate, and the default rate. If the recovery rate increases by 5%p from the base model, consumption increases by 0.95% and the tax rate and the default rate decrease by 5.29% and 0.38%, respectively. In addition, when consumption increases by 3.50%p, the tax rate and the default rate decrease by 22.7% and 1.66%, respectively, according to an increase in the recovery rate of 100%.

[7]The counterfactual assumptions of recovery rates are 25% and 40%. The rate of 25% is based on the recovery rate of commercial banks, and 40% represents an increase in the recovery rate of the baseline by 100%.

TABLE 5—OPTIMAL COVERAGE RATIO AND INTEREST-SUPPORT SPREAD ACCORDING TO THE RECOVERY RATE

Recovery rate (*rc*)	Coverage ratio (*c*)	Interest-support spread (r_g)
20%	8.75%	0.25%
25%	8.75%	0.25%
40%	16.25%	0.50%

Note: This table shows average values of results according to γ ranging from 4.0 to 5.5.

TABLE 6—CHANGES IN SOCIAL WELFARE ACCORDING TO AN INCREASE IN THE RECOVERY RATE

From 20% (*rc*)	Consumption (*C*)	Tax (τ)	Default rate
5%p increase	+0.95%	-5.29%	-0.38%
20%p increase	+3.50%	-22.7%	-1.66%

Note: This table shows average values of results according to γ ranging from 4.0 to 5.5.

V. Analysis of the Continuous Treatment Effect for Credit Guarantees

In this section, we analyze the effects of a credit guarantee program on performance when SMEs are supported by KODIT. In particular, , we use the ratio of guaranteed debt to the total debt in a continuous treatment rather than a binary treatment effect analysis of the average treatment effect (Hirano and Imbens 2004). This analysis of the continuous treatment effect can determine the marginal effect of policy variables instead of the average effect of exposure to a policy. Moreover, most Korean policy-research papers analyze only the average treatment effect; hence, the present analysis of the continuous treatment effect contributes to the research on SME financial policies in an important way.

A. *Methodology of the Analysis of the Continuous Treatment Effect*

1. *Generalized Propensity Score Method*

Our analysis is based on the GPS (generalized propensity score) method of Hirano and Imbens (2004). The generalized propensity score removes the endogeneity of selection bias by controlling the propensity to be selected through the characteristics of objects as the generalization of the propensity score for a binary treatment effect. Hirano and Imbens (2004) assume unconfoundness when controlling this endogeneity such that for any treatment $t \in [0,1]$, $Y(t) \perp T \mid X$, is satisfied. $Y(t)$ denotes the performance of the treatment variable (t), T represents the continuous treatment, X is the pre-treatment variable in each case, properly defined based on the probability measure. This assumption means that the treatment characteristics are independent of the performance. Another assumption is a balancing property such that

$$X \perp 1\{T = t\} \mid \rho(t,x),$$

in which $\rho(t,x)$ is $f_{T|X}(t \mid x)$, the conditional probability of the treatment derived characteristics.

Hirano and Imbens (2004) prove using these assumptions that

$$f_T(t \mid \rho(t,X), Y(t)) = f_T(t \mid \rho(t,X) \text{ and}$$
$$\Xi(t,r) = E\left[Y(t) \middle| \rho(t,X) = \rho\right] = E\left[Y \middle| T = t, R = \rho\right].$$

The last equation is $\psi(t) = E\left[\Xi\left(t, \rho(t,X)\right)\right]$, representing the expected performance according to continuous treatment and the generalized propensity score.

In more detail, regarding the given characteristics X_i, we assume that $g(T_i) \mid X_i \tilde{} N\left\{h(\xi, X_i), \sigma^2\right\}$. Subsequently, from

$$\hat{R}_i = \frac{1}{\sqrt{2\pi\widehat{\sigma^2}}} exp\left[-\frac{1}{2\widehat{\sigma^2}}\left\{g(T_i) - h\left(\hat{\xi}, X_i\right)\right\}\right],$$

we can estimate the GPS by means of maximum likelihood estimation.

Next, we estimate the expected performance using quadratic regression with estimated GPS and continuous treatment variable such that

$$E\left(Y_i | T_i, R_i\right) = \psi\left(T_i, R_i; \alpha\right) = \alpha_0 + \alpha_1 T_i + \alpha_2 T_i^2 + \alpha_3 R_i + \alpha_4 R_i^2 + \alpha_5 T_i R_i.$$

Finally, we calculate the average performance according to each treatment level with the expected performance such that

$$E\left\{\hat{Y}(t)\right\} = \frac{1}{N}\sum_{i=1}^{N}\hat{\psi}\left\{T_i, \widehat{R_i}; \hat{\alpha}\right\}.$$

2. *Data for the Empirical Analysis*

The data used here are the loan guarantees of KODIT from FY 2008 to FY 2011. We merged the KODIT data and the KIS database for the SME performance measures. The average coverage ratio for a loan is 92.9%, and most loan guarantees have coverage ratios of 85%, 90%, 95%, and 100%. It should also be noted that most coverage ratios are high and inflexible considering the SMEs' characteristics.

We compute the ratio of guaranteed debt to total debt for each SME in the KODIT data as the continuous treatment effect. If the maturity of the guarantee is longer than one year, we cover that loan for longer than one year, and if one SME

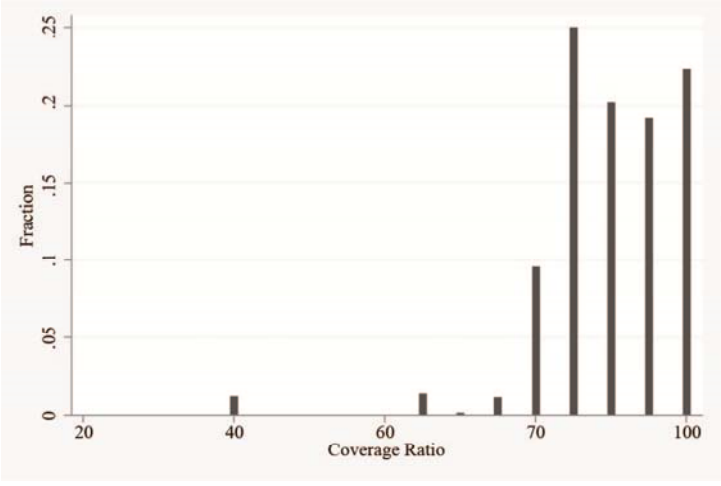

FIGURE 5. DISTRIBUTION OF COVERAGE RATIOS FOR SME LOANS

Data: Loan cases of KODIT from FY 2008 to FY 2011.

has several credit guarantees for loans, we compute the ratio of guaranteed debt to total debt by adjusting the data according to the maturities of the guarantees.

The log values of total assets, debt to assets, ROA, trade payables to sales, and the financial cost to the total cost are used as the SMEs' pre-treatment variables (or characteristics), from Nam (2013b). Additionally, the interest expense with regard to the total debt and the log values of sales are used to control the interest cost according to the change in the credit amount and the operating performance values, respectively.

When dealing with raw data, we need to adjust the time lag between the pre-treatment variables (firm characteristics) and the performance measures. First, every variable is computed as of the end of the fiscal year. If a credit guarantee was approved before the end of June in year t, then both the pre-treatment variables at year $t-1$ and the performance measures at year t are considered in the same observation. If the time of approval of the credit guarantee was after the end of June in year t, then the pre-treatment variables at year t, and performance measures at year $t+1$ are considered in the same observation. Thus, guaranteed debt which started in March of 2008 is grouped with the pre-treatment variables of 2007, but not with the pre-treatment variables of 2008, implying that KODIT's decisions on credit guarantees made before July are assumed to include only the information up to the previous year, whereas decisions after June are assumed partially to use the information of the same year. In addition, if one firm has several guarantees during the same year, we compute the observation as of the latest time because the first decision among the guarantees is assumed to have used the most crucial information.

B. *Results of the Continuous Treatment Analysis*

1. *Basic Statistics*

The number of firm-year observations computed through the method explained in the previous section is 38,370, but most SMEs in the data are not externally audited. Therefore, the data are less reliable given this information and the great number of outliers. Thus, we randomly sample 3,000 observations.[8]

Table 7 shows the statistics of the randomly sampled data used here. The mean, median, and standard deviation of the continuous treated guaranteed debt over the total debt are 39.3%, 28.3%, 32.5%, respectively. However, the proportion of examples with a 100% guarantee for all debt amounts to 12%, indicating that all of the debt of some SMEs is completely covered by credit guarantees. Thus, these SMEs may be considered to have excessive financial support from the government.

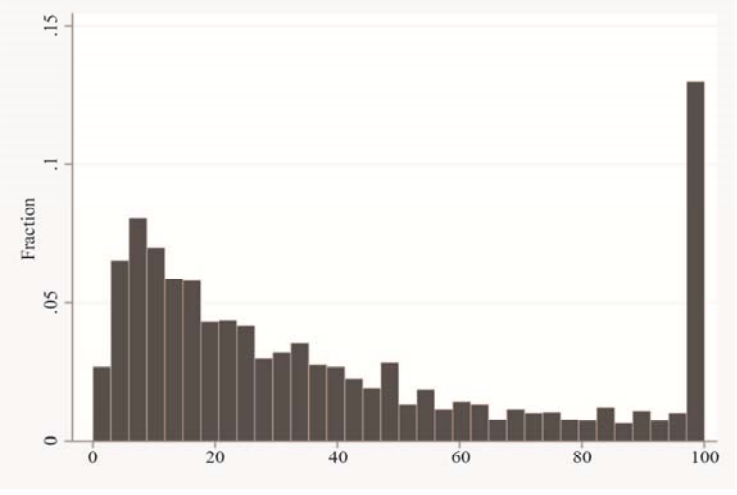

FIGURE 6. DISTRIBUTION OF THE RATIO OF GUARANTEED DEBT TO TOTAL DEBT

Data: KODIT's loan data (FY 2008~FY 2011), KIS Database.

TABLE 7— BASIC STATISTICS OF PRE-TREATMENTS AND PERFORMANCES

Variables	Mean	Median	S.D.
Guaranteed debt/total debt	39.3%	28.3%	32.5%
Log (assets)	14.3	14.4	1.39
Log (sales)	15.0	15.0	1.36
Total debt/assets	34.8%	35.8%	26.1%
ROA	6.62%	9.96%	80.4%
Trade payables/sales	4.09%	6.57%	9.04%
Financial cost/total cost	0.98%	1.91%	3.62%
Interest cost/total debt	5.96%	7.04%	7.58%

Data: KODIT's loan data (FY 2008~FY 2011), KIS database.

[8]As a pre-analysis, the main features of the empirical results with all observations are consistent wit h those of the randomly sampled observations.

TABLE 8—COEFFICIENTS OF GPS FOR GUARANTEED DEBT TO TOTAL DEBT

Variables	Coefficients	t-values
Log (sales)	0.03	1.04
Log (assets)	-0.41	-12.9
Total debt/assets	-0.02	-15.8
ROA	-0.01	-5.67
Trade payables/sales	0.000	0.01
Financial cost/total cost	0.001	0.15
Interest cost/total debt	0.002	0.89

According to Table 8, the log (assets), total debt/assets, and ROA significantly influence the GPS function.

2. *Analysis Results*

Table 8 shows the GPS coefficients for the continuous treatment log-normalized, ratio of guaranteed debt to total debt using the method of Hirano and Imbens (2004). The log values of sales, the total debt to assets, and the ROA are negatively and significantly correlated with the continuous treatment, indicating that as a firm is smaller, has less debt, and is less profitable, it can have relatively more guarantees.

Table 9 shows the statistics of firms' financial ratios as continuous treatment performance measures. Recently, the means of the growth rates of net income and capital have been negative due to poor operating conditions and low profitability, but the medians stand at 14.9% and 19.4%, respectively. Although most studies commonly use the growth rate of assets or sales as performance measures, these variables are not completely free from endogeneity because the growth of assets must be correlated with an increase in debt by credit guarantees and the approval of a credit guarantee may be due to a new contract promising massive sales. Therefore, we only use the value-added to capital and the value-added to sales as performance measures of SME productivity and use the growth rates of net income and capital to represent profitability and growth potential.

Figures 7-10 show the continuous treatment effects and the percentage ratio of guaranteed debt to all debt on firms' performances. Specifically, the upper panel in the figure shows the marginal effect of a 1%p increment in treatment, and the lower panel shows the level effect of the treatment.

Figure 7 shows the effect of the treatment on the percentage ratio of value-added to capital. The upper panel shows that the marginal effect of the policy treatment is highest from 15% to 20%, and the marginal effect decreases with doses of the

TABLE 9—BASIC STATISTICS OF PERFORMANCE MEASURES

Variables	Mean	Median	S.D.
Value-added/capital	68.6%	43.9%	85.7%
Value-added/sales	28.1%	24.3%	25.6%
Growth of net income	-36.2%	14.9%	3,566%
Growth of capital	-8.53%	19.4%	3,994%

Data: KIS Database.

treatment such that the response function peaks at a guaranteed debt to total debt ratio of 65%.

Figure 8 shows that the marginal effect of the treatment on value-added to assets is highest at around 15% and that the performance of the treatment peaks at 65%. Figure 13 shows the similarity of the effect of net income growth with the results presented in Figure 12. Particularly, the marginal effect of this treatment becomes more significance as the treatment increases.

Figure 9 shows a similar pattern in that the marginal effect is higher at lower treatment values, and the response has its greatest effect in the middle of the treatment. Figure 10 shows the effect of the treatment on capital growth. The marginal effect decreases with the treatment and negatively affects capital growth when the ratio of guaranteed debt to all debt exceeds 65%.

According to the empirical results, the public guarantee program has a positive impact on a firm's performance when the firm's total debt is less covered by such a guarantee program. However, as coverage by the guarantee program for a firm's

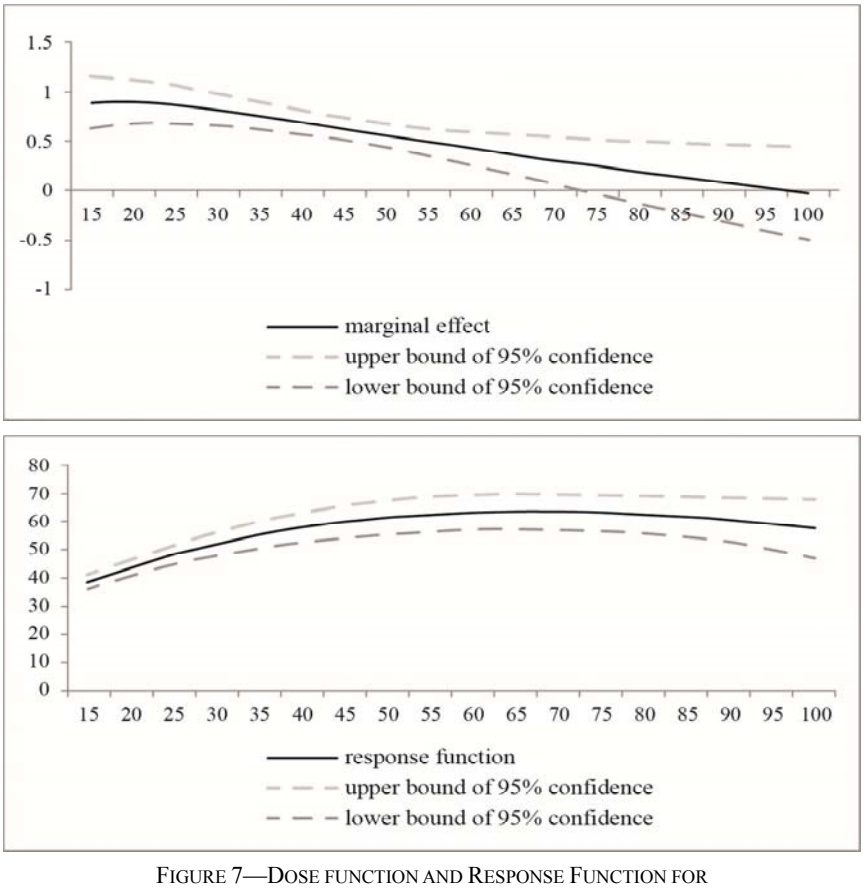

FIGURE 7—DOSE FUNCTION AND RESPONSE FUNCTION FOR VALUE-ADDED TO CAPITAL

Note: The horizon axis denotes guaranteed debt to total debt (%), and the vertical axis is performance measure (%). The upper panel shows the dose effect for a 1%p increment of the treatment, and the lower panel shows the response of the performance measure to the treatment.

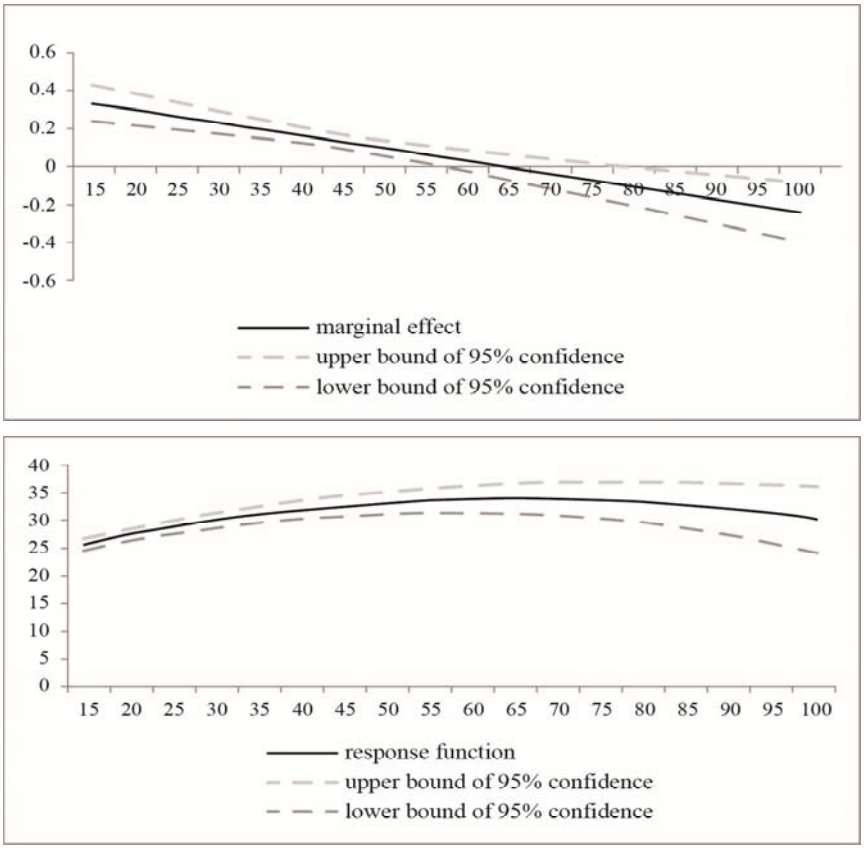

FIGURE 8. DOSE FUNCTION AND RESPONSE FUNCTION FOR VALUE-ADDED TO SALES

Note: The horizon axis denotes guaranteed debt to total debt (%), and the vertical axis is performance measure (%). The upper panel shows the dose effect for a 1%p increment of the treatment, and the lower panel shows the response of the performance measure to the treatment.

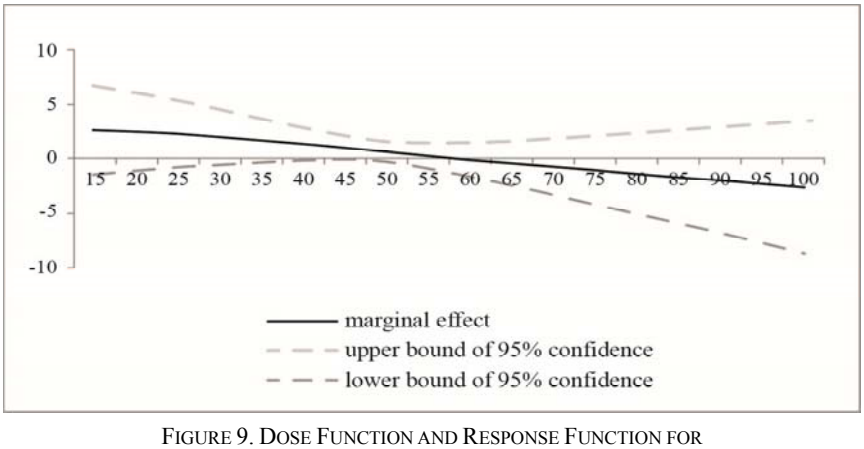

FIGURE 9. DOSE FUNCTION AND RESPONSE FUNCTION FOR NET INCOME GROWTH

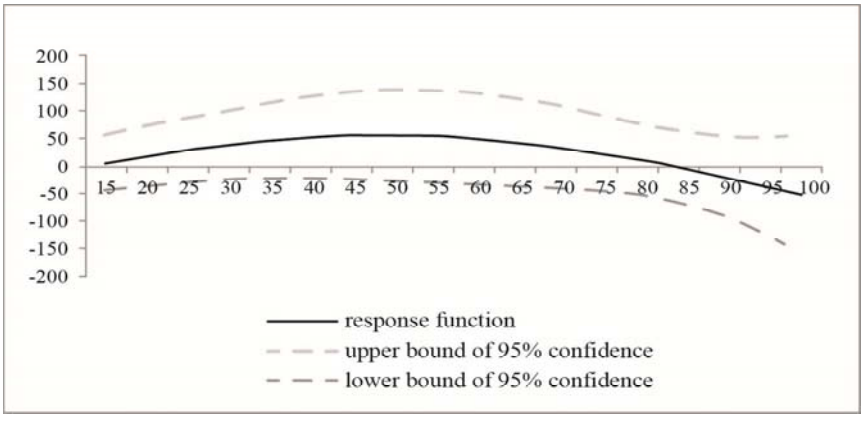

FIGURE 9. DOSE FUNCTION AND RESPONSE FUNCTION FOR NET INCOME GROWTH (*CONTINUED*)

Note: The horizon axis represents the guaranteed debt to total debt (%), and the vertical axis is the performance measure (%). The upper panel shows the dose effect for a 1%p increment of the treatment, and the lower panel shows the response of the performance measure to the treatment.

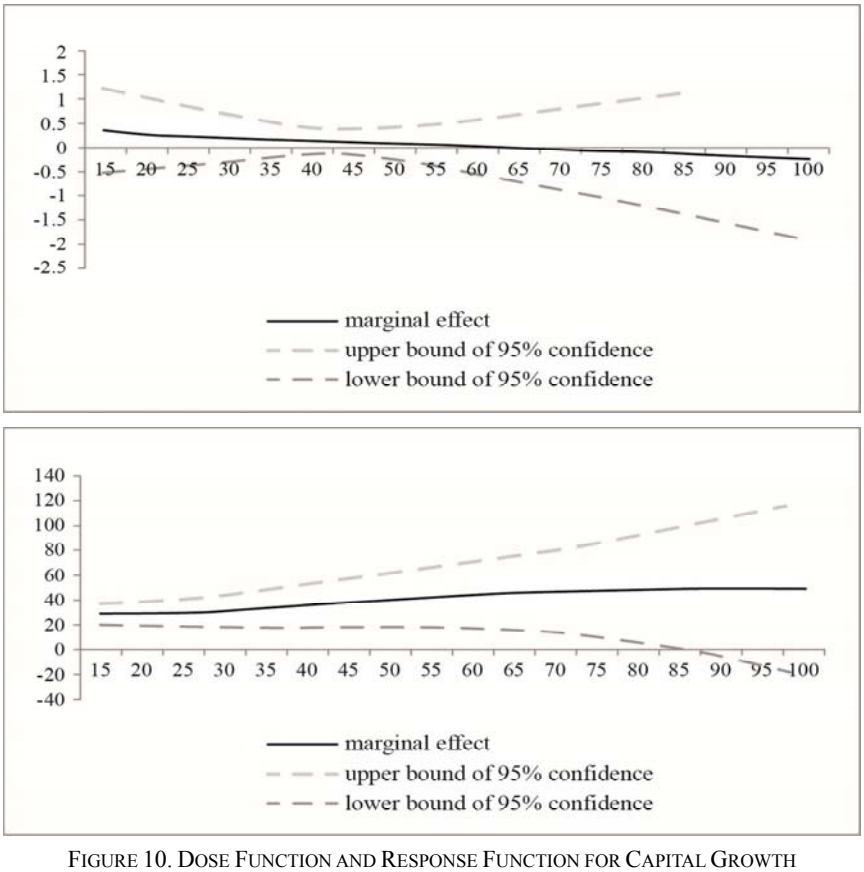

FIGURE 10. DOSE FUNCTION AND RESPONSE FUNCTION FOR CAPITAL GROWTH

Note: The horizon axis denotes the guaranteed debt to total debt (%), and the vertical axis is the performance measure (%). The upper panel shows the dose effect for a 1%p increment of the treatment, and the lower panel shows the response of the performance measure to the treatment.

total debt increases, the firm faces a moral hazard because the burden of all debt can be avoided by the credit guarantee. Therefore, the excessive supply of public guarantees for SMEs while not considering the efficient allocation of public resources may negatively influence economic performance overall given the prevalence of moral hazard on the part of the borrower.

VI. Conclusion

This paper investigates how financial support schemes such as credit guarantee and interest support programs for SMEs dynamically affect corporate decisions, including default decisions, in an equilibrium model with information asymmetry. Particularly, we calculate the optimal sizes of the credit guarantee and interest-support spread in a perfectly competitive loan market with imperfect information for SMEs. The simulation results show that the current levels of the credit guarantee size and interest-support spread may be excessive, thus above the optimal levels.

Second, our study empirically analyzes the effects of credit guarantee programs using the ratio of guaranteed debt to total debt as a continuous treatment variable with the GPS method. According to the results, the marginal effects of credit guarantees of productivity, profitability and growth potential decrease with the ratio of guaranteed loans to total debt. In addition, the response functions for a credit guarantee peaks between 50% and 60%.

Finally, this paper proposes several policy improvements for the credit guarantee programs. First, the government needs to lower the amount of total credit guarantees with reference to all SME loans because the proportion of the public credit guarantee to all SME loans at Korea is higher than those in other OECD countries, and theoretical simulations show that consumers pay more tax than the optimal level of social welfare. Moreover, it is possible that the excessive financial support for SMEs hampers prudential firm dynamics, including productivity and investment, by delaying the exit of poor SMEs from the market.

Second, an increase in the SME guarantee fee or bank contribution rate is more likely to boost the default rate or loan interest rate such that any positive effects of the SME financial support may be weakened. Thus, the government needs to strengthen regulations pertaining to recovery for indemnity from bankrupt companies, which would distort the default rate and loan interest rate less. Despite the limitations of our model, it was found that an increase in the recovery rate of 100% can reduce the consumer tax burden by 22.7%. Additionally, strong policies for collecting debt from bankrupt SMEs can prevent moral hazard.

Lastly, current regulations which control the amount of credit guarantees per firm do not limit the proportion of credit guarantees to a firm's total debt. Therefore, the government must regulate the limit on the ratio of guaranteed debt to overall debt and must flexibly manage the coverage ratio for each loan application in order to prevent the negative effects of the credit guarantee program.

APPENDIX

1. *Computation of transition probability of* z

In order to compute z, we define the return on assets such that

$$ROA = z(1+b)^{\alpha-1},$$

and if the leverage is defined as b, then

$$\frac{ROA}{(1+b)^{\alpha-1}} = z$$

can be computed. First, we compute z with data from FY 2000 to FY 2011 from the KIS database, categorizing 25 ranges of z. Finally, the transition probability of z is computed with a matrix of z 25×25 in size. This method is analogous to the approach taken by Adda and Cooper (2003).

2. *Computation of the dynamic model*

We compute the upper limit of b, $\bar{b}$, such that

$$\frac{\bar{z}(1+\bar{b})^{\alpha}}{\bar{q}-1} \geq \bar{b}.$$

1. Compute $V_T(b,z)$ by the backward induction of dynamic programming given q_0.
2. Compute $V_{T-1}(b,z)$ with $V_{0,T-1}(b,z)$ and $V_{1,T-1}(b,z)$ after computing $E[V_T(b)]$ using the transition matrix of z.
3. Repeat step 2 until $V(b,z)$ converges with some extent to precision.
4. Through steps 2 and 3, find the firm's optimal policy given (b,z) and then compute $\mu(b,z)$.
5. Compute q_1 with μ.
6. Repeat from step 1 to step 5 until q is converged with some extent to precision.
7. Compute B^*, τ^*, and C^* with μ^* and q^* in step 6.

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