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The Effects of Lowering the Statutory Maximum Interest Rate on Non-bank Credit Loans†

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This paper analyzes the effects of the cut in the legal maximum interest rate (from 27.4% to 24%) that occurred in February of 2018 on loan interest rates, the default rates, and the loan approval rate of borrowers in the non-banking sector. We use the difference-in-difference identification strategy to estimate the effect of the cut in the legal maximum interest rate using micro-level data from a major credit-rating company. The legal maximum rate cut significantly lowers the loan interest rate and default rate of low-credit borrowers (i.e., high-credit-risk borrowers) in the non-banking sector. However, this effect is limited to borrowers who have not been excluded from the market despite the legal maximum interest rate cut. The loan approval rate of low-credit borrowers decreased significantly after the legal maximum interest rate cut. Meanwhile, the loan approval rate of high-credit and medium-credit (i.e., low credit risk and medium credit risk) borrowers increased. This implies that financial institutions in the non-banking sector should reduce the loan supply to low-credit borrowers who are no longer profitable while increasing the loan supply to high- and medium-credit borrowers.

Key Word: Statutory Maximum Interest Rate, Household Loan, Market Exclusion, Non-banking Sector
JEL Code: G23, G28, G51

I. Introduction

The statutory maximum interest rate[1] refers to the highest interest rate allowed by law[2] for a loan product. The statutory maximum interest rate system was

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[1]In this paper, I use "statutory maximum interest rate" and "legal maximum interest rate" interchangeably with the same meaning.

introduced to prevent abuse of their market power by financial institutions and to protect low-income households in the loan market. The frequent entries and exits of financial institutions can lead to distrust of financial consumers about the possibility of deposit recovery, which increases the likelihood of a bank run. Accordingly, the government tolerates the market power of financial institutions to some extent based on the authorization of financial institutions' market entrance and promotes system stability. However, based on their market power, financial institutions may impose unreasonable interest rates on households lacking bargaining power. Thus, the government is implementing the statutory maximum interest rate system to prevent this.

Considering these points, many countries, including Korea, have introduced statutory maximum interest rates and have legally restricted the maximum interest rate level. The Loan Business Act, the focus of this study, was enacted in October of 2002, and the legal maximum interest rate was initially set to 66% according to the enforcement decree. Since then, the enforcement decree has been revised seven times, and the current legal maximum interest rate is 20%. In particular, in February of 2018, the legal maximum interest rate was cut by 3.9%p (27.9% → 24%), and in July of 2021, the legal maximum interest rate was cut further by 4%p (24% → 20%). In addition, discussions are continuing in political circles about the possibility of further cuts in the legal maximum interest rate.

Reducing the legal maximum interest rate can have two major effects on the utility of financial consumers. First, the interest rates of some borrowers, especially those who paid interest rates close to the upper limit of the maximum interest rate, may be lowered by the legal maximum interest rate cut. Borrowers who received interest rates close to the upper limit of the maximum interest rate are more likely to be from low-income households. Therefore, a cut in the legal maximum interest rate can increase the disposable income of borrowers who earn relatively low incomes.

On the other hand, financial institutions that provide high-interest loans are likely not to give loans to certain borrowers who are no longer seen at profitable when the legal maximum interest rate is lowered. In particular, such action is highly likely to reduce the availability of loans provided to borrowers with a high probability of default. In general, as borrowers with lower income levels are more likely to default, these borrowers are highly likely to be excluded from the loan market.

Considering these two points, existing studies have estimated whether such a cut in the legal maximum interest rate lowers interest rates and increases market exclusion. However, our paper differs from previous studies in terms of two major aspects.

First, unlike previous studies, this study improved the accuracy of the analysis of the impact of the cut in the legal maximum interest rate by using micro-level data from a credit rating agency. The data used in this study are individual level micro-data provided by the Korea Credit Bureau (hereafter KCB). KCB data include loan data from all financial institutions in the banking and non-banking financial sectors.

[2]Accordingly, for a loan contract that exceeds the statutory maximum interest rate, the interest contract for the excess portion is invalid and cannot be claimed in court. The Loan Business Act and the Interest Restriction Act stipulate the maximum interest rate. The Loan Business Act applies to financial and loan businesses authorized, licensed, and registered under the Act. On the other hand, the Interest Restriction Act stipulates the maximum interest rate for loan transactions between private parties.

The non-banking sector covered in this paper includes card companies, capital companies, savings banks, insurance companies, and cooperatives, but not lending companies. The data also include complete loan histories by borrower, credit evaluation histories, credit and debit card usage, and various individual characteristic variables. KCB also has data on loan attempts. When someone attempts to secure a loan, they conduct a credit check, and this record is kept by KCB. Credit inquiries are divided into simple inquiries and loan evaluation inquiries, and the credit inquiry as used in this study is the latter type. Thus, we can determine whether a particular borrower has attempted to secure a loan through this type of credit check record.

Using micro-level data for each borrower, we can identify the impact of the cut in the legal maximum interest rate in the non-banking sector using a difference-in-difference strategy. In particular, we analyze borrowers by dividing them into high-credit, medium-credit, and low-credit classes. High-credit borrowers who belong to credit grades 1-3 represent low credit-risk borrowers. On the other hand, medium-credit borrowers who belong to credit grades 4-7 represent medium credit-risk borrowers, and low-credit borrowers who belong to credit grades 8-10 represent high credit-risk borrowers.

From various angles through descriptive statistics, the cut in the statutory maximum interest rate mainly affects the loans of low-credit borrowers (i.e., high-credit-risk borrowers) in the non-banking financial sector and some mid-credit borrowers in the non-banking financial sector. However, the cut in the statutory maximum interest rate barely affects the loans of high-credit borrowers (i.e., low-credit-risk borrowers), especially in the banking sector. Accordingly, in this paper, the control group will be credit loans of high-credit borrowers from the banking sector. In contrast, the treatment group will be credit loans of high-credit, medium-credit, and low-credit borrowers from the non-banking financial sector.

Meanwhile, to the best of the author's knowledge, this is the first study to analyze the legal maximum interest rate cut (27.9% → 24%) enacted in February of 2018. Although the statutory maximum interest rate had been cut several times prior, lowering the statutory maximum interest rate from 27.9% to 24% may have a very different effect from the previous cuts, as not only the macroeconomic environment at the time of the statutory maximum rate cut but also the distribution of default rates of financial consumers significantly influence the effectiveness of the statutory maximum rate cut[3].

As a result of the analysis, the statutory maximum interest rate cut in February of 2018 significantly lowered the loan interest rates for households not excluded from the credit loan market even after the statutory maximum rate was cut. In particular, the interest rate on credit loans for the low-credit class in the non-banking financial sector was cut by a significant amount (3.5%p). On the other hand, the interest rate on credit loans for the middle-credit class in the non-banking financial sector was reduced by only 0.16%p, and no statistically significant change was found in the

[3]For example, suppose the probability of default by most borrowers is not that high. In such a case, a 7%p reduction in the statutory maximum interest rate from 34.9% to 27.9% may have a relatively small effect, but continuously cutting the statutory maximum interest rate afterward will gradually increase the impact on the loan interest rate and the market exclusion rate. In the same vein, the reduction of the legal maximum interest rate from 24% to 20% requires further analysis in the future. However, this study focuses on the legal maximum rate cut in February of 2018 (27.9% → 24%) due to data limitations.

credit loan interest rate for the high-credit class in the non-banking financial sector.

In addition, the default rate of borrowers who were approved for a loan despite the cut in the legal maximum interest rate decreased significantly. This effect mainly affected the low-credit class. For the mid-credit class in the non-banking financial sector, the probability of default decreased by about 0.24%p on average due to the reduced maximum interest rate. On the other hand, the low-credit class overall showed a 2.8%p decrease.

As such, the cut in the statutory maximum interest rate reduces the debt repayment burden for borrowers who were not excluded from the market after the cut, hence decreasing the default probability. However, this result is limited to borrowers who were not excluded from the market despite the reduction. Borrowers thus excluded from the market will be pushed to loan businesses or the non-institutional financial market.

According to the analysis, as the legal maximum interest rate was reduced from 27.9% to 24% in February of 2018, the loan approval rate of the low-credit group decreased by 3.6%p. On the other hand, the loan approval rate of the high- and medium-credit groups increased by approximately 1.0%p and 1.4%p, respectively.

The structure of this paper is as follows. Chapter II examines earlier work in this area. Chapter III examines the general changes in the credit loan market before and after the statutory maximum interest rate cut through a descriptive statistical analysis. Chapter IV introduces the empirical method, and Chapter V presents the results of the empirical analysis. Finally, the paper ends with the conclusion in Chapter VI.

II. Literature Review

Previous research related to the legal maximum interest rate is largely divided into studies of the effects of legal maximum interest rate cuts and studies of methods to determine the loan interest rate and the cost of loan businesses.

Kim (2017) estimated the extent to which low-credit borrowers in the bank and non-bank financial sectors were excluded from the market due to a cut in the legal maximum interest rate that occurred between July of 2010 and July of 2017. According to the analysis, as the top interest rate decreases by 1%p in the entire financial sector, the number of new borrowers with low credit will decrease by 3.585%. In non-bank entities, the number of new borrowers with low credit will decrease by 3.398% as the top interest rate decreases by 1%p.

Noh *et al*. (2013) analyzed the impact on financial consumers when the statutory maximum interest rate, which was 39%, was reduced to 30%. According to their analysis, the financial costs associated with loan refusals greatly exceed the interest cost reduction benefit, and financial consumers amounting to more than twice the number of borrowers who receive the benefit are excluded from the low-income financial market. Noh (2014) argues that in order to improve the predictability of statutory maximum interest rates, it is necessary to predict a schedule of changes of the upper limit of the interest rate and/or to consider linking the statutory maximum interest rate to the market interest rate. Lee (2015) also argues that the interest rate cap must be managed in a relative manner that links the interest rate with the market interest rate to reflect both the current low-interest rate trend and market conditions properly.

Ryu (2016) raised the need to expand public microfinance and strengthen follow-up management, as there is a high risk that a cut in the highest interest rate will lead to a reduction in the supply of funds by financial companies to the low-credit class of borrower.

Lee (2011) and Lee and Song (2021) analyzed the effect of legal maximum interest rate cut on loan companies. According to Lee (2011), the loan interest rates of loan businesses are insensitive to changes in market interest rates. He judged that regulations such as lowering the upper limit of interest rates were necessary, as this phenomenon was presumed to be due to chronic excess demand, imperfect competition, and information asymmetry. On the other hand, in Lee and Song (2021), the number of loan users and the number of new loans decreased due to the reductions of legal maximum interest rate. This suggests, unlike in the past, that the recent cut in the legal maximum interest rate has resulted in a level that can seriously damage the loan market.

Lee (2016a) shows that the number of low-credit borrowers excluded from the loan market is expected to range from at least 350,000 to at most 740,000 when the legal maximum interest rate is cut from 34.9% to 27.9%. On the other hand, the size of the loan market increased in terms of the loan amount and number of traders after the previous lowering of the upper limit of the interest rate, meaning that the problem of credit shrinkage due to the lowering of the upper limit of the interest rate did not come to the fore.

According to Jeong (2007a; 2007b), because the loan market has an imperfect competition structure, loan companies can obtain profits by imposing high-interest rates based on their monopoly power, even over high-quality customers. In addition, he argues that if micro-credit loans from low-income financial institutions are activated, the problems caused by high-interest rates by lending companies will be resolved to a large extent.

Lee (2019) suggests the need to change the loan interest rate standard to a more straightforward form to protect financial consumers. In addition, he argues that the government needs to manage the market by focusing on the degree of interest rate fluctuations after lending and the fairness of interest rate application rather than the level of the loan interest rate.

Lee and Han (2013) studied the interest rate determination mechanism in the Korean-Japanese loan market. In their study, they argue that Korea, like Japan, should also make it mandatory to subscribe to a personal credit information DB integrated with lending companies, thereby eliminating the factor of information asymmetry between lenders and users.

Finally, Lee (2016b) showed that the change in the interest rate cap regulation in march 2016 affects the profit and loss of the lending company through a cost rate analysis of loan businesses. In particular, they point out that loan companies that cannot reach the break-even point are eliminated from the market, leading to changes in the market's competitive structure.

III. Data and Descriptive Statistics

This chapter analyzes data for one year (February 2017 - February 2019) before and after the legal maximum interest rate was cut by 3.9%p from 27.9% to 24% on February 8, 2018. The loan interest rate is generally determined by adding a certain amount of margin to the sum of the funding rate, taxes, and the credit risk costs. The credit risk cost depends on the recovery rate. The recovery rate means the ratio of principal that can be recovered by disposing of collateral even in the event of a default. Mortgage loans such as home mortgages have a high recovery rate and thus the credit risk cost is low compared to credit loans. Accordingly, the interest rate level of mortgage loans is significantly different from the legal maximum interest rate level. Therefore, this study mainly focuses on the credit loan market, especially credit loans from the non-banking financial sector. This study uses borrower-level microdata from KCB, one of Korea's representative credit-rating agencies.

A. *Credit Loan Market: Banking and Non-banking Sectors*

This section examines descriptive statistics of the credit loan markets in the banking and non-banking financial sectors. The KCB data used in this study include data from all financial institutions in the banking and non-banking sector, encompassing all loan histories, credit evaluation histories, credit and debit card usage statistics, and a range of characteristic variables for each individual. Also, individuals without a record of receiving a credit loan between January 2013 and March 2021 are also included in the data. Furthermore, KCB also provides data on loan attempts. When someone attempts to secure a loan, a credit check occurs, and this record is kept at KCB. Credit inquiries are divided into simple inquiries and inquiries for a loan evaluation. The type of credit inquiry used in this study is the inquiry for a loan evaluation, not merely a simple inquiry. Thus, we can determine whether a borrower has attempted to secure a loan or not.

Table 1 shows descriptive statistics for newly issued credit loans in the banking sector between February of 2017 and February of 2019 – one year before and after the legal maximum interest rate was cut from 27.9% to 24%. Meanwhile, Table 2 shows descriptive statistics on newly issued non-bank credit loans during the same period.

The distribution of interest rates on credit loans in banks and non-banks is very different. First, the average interest rate for bank credit loans is about 4.2%, whereas, for non-bank credit loans, it is approximately 14.2%, showing a difference of about 10%p. In addition, the interest rate distribution of loans in the banking sector is relatively dense compared to that of loans in the non-banking sector. The 10th percentile of the bank's credit loan interest rate is 2.7%, and the 90th percentile is 6.3%, showing a difference of about 3.6%p. On the other hand, the 10th percentile of non-bank lending rates is 4.9%, and the 90th percentile is 22.7%, showing a considerable difference of about 17.8%p. In other words, the dispersion of interest rates on credit loans in non-banking sectors is much broader than that in banks.

The interest rates on credit loans in banks do not differ significantly between financial institutions and are generally low. On the other hand, in the non-banking sector, some institutions issue credit loans with relatively low interest rates, such as insurance and cooperatives (e.g., credit cooperatives, fisheries cooperatives, NongHyup). However, at the same time, card companies, capital companies, and savings banks included in the non-banking sector supply high-interest credit loans.

Next, the delinquency rates of banks and non-banks also show a considerable difference. In general, when looking at defaults of 90 business days or more, which is the general standard for a default, the average default rate of bank credit loans is about 1.2%. In contrast, the average default rate of non-bank credit loans is about 3.7%. As an index directly related to the difference in delinquency rates, the average credit score of banks is about 869.8, while the average credit score of non-banks is close to 768.7.

The average annual income of borrowers in the banking sector is about 48 million won. In comparison, the average annual income of non-bank borrowers is about 29 million won, indicating that the average income of borrowers in the banking sector is about 65% higher.

Panel D of Table 1 shows the average loan amount per borrower by financial sector and loan type. From February of 2017 to February of 2019, the average sum

TABLE 1—DESCRIPTIVE STATISTIC OF BANK CREDIT LOAN BORROWERS

		N	Mean	S.D.	p10	p50	p90
A. Loan Contract Characteristics (Account)	Interest rate (%)	304,839	4.2	1.8	2.7	3.8	6.3
	Loan Amount(1 million won)	304,839	32.847	35.743	3	20	80
	Loan Term (month)	304,839	17.328	10.775	12	12	36
	Repayment Amount (1 million won)	304,839	2.651	3.886	0.175	1.690	6.222
B. Loan Contract Performances (Account)	Delinquency over 30 Days in the Next Year	304,839	0.006	0.079	0	0	0
	Delinquency over 90 Days in the Next Year	304,839	0.003	0.058	0	0	0
	Delinquency over 30 Days after Loan Contract	304,839	0.017	0.128	0	0	0
	Delinquency over 90 Days after Loan Contract	304,839	0.012	0.110	0	0	0
C. Borrower Characteristics (Account)	Income (1 million won)	304,839	47.579	26.761	21.280	40.170	82.660
	Age	304,839	44.513	10.469	30	40	60
	Job (Employed)	225,879	0.959	0.199	1	1	1
	Credit Score	304,839	869.83	97.69	729	895	974
	Credit Card Usage (1 million won)	304,839	20.519	13.596	3.907	18.537	40.288
	Debit Card Usage (1 million won)	304,839	3.003	3.328	0	1.692	8.42
D. Borrower Characteristics (person, previous month)	Bank Loan Balance (1 million won)	208,290	66.2	95.3	0	28.8	184.2
	Bank Credit Loan Balance (1 million won)	208,290	22.3	34.1	0	7.2	63.4
	Non-bank Credit Loan Balance (1 million won)	208,290	2.6	3.3	0	0	6.7
E. Borrower Characteristics (within 30 days before new contract)	Bank Credit Loan Application	297,401	1.675	0.921	1	2	3
	Non-bank Credit Loan Application	297,401	0.693	1.094	0	0	2

Note: Descriptive statistics for Job (Employed) is calculated, except for borrowers whose job information is ‘other’.

TABLE 2—DESCRIPTIVE STATISTIC OF NON-BANK CREDIT LOAN BORROWERS

		N	Mean	S.D.	p10	p50	p90
A. Loan Contract Characteristics (Account)	Interest rate (%)	894,326	14.2	6.3	4.9	14.6	22.7
	Loan Amount(1 million won)	894,326	8.837	11.396	1	5	20
	Loan Term (month)	894,326	23.945	11.789	11	24	37
	Repayment Amount (1 million won)	894,326	0.475	0.832	0.087	0.282	0.901
B. Loan Contract Performances (Account)	Delinquency over 30 Days in the Next Year	894,312	0.030	0.171	0	0	0
	Delinquency over 90 Days in the Next Year	894,312	0.016	0.124	0	0	0
	Delinquency over 30 Days after Loan Contract	894,326	0.055	0.229	0	0	0
	Delinquency over 90 Days after Loan Contract	894,326	0.037	0.188	0	0	0
C. Borrower Characteristics (Account)	Income (1 million won)	894,326	29.213	13.905	18	26	43
	Age	894,326	49.582	11.442	30	50	65
	Job (Employed)	461,674	0.760	0.427	0	1	1
	Credit Score	894,326	768.7	107.0	638	766	917
	Credit Card Usage (1 million won)	894,326	20.52	13.60	3.91	15.54	40.29
	Debit Card Usage (1 million won)	894,326	19.61	14.37	2.42	17.00	41.23
D. Borrower Characteristics (person, previous month)	Bank Loan Balance (1 million won)	589,153	20.1	52.2	0	0	67.8
	Bank Credit Loan Balance (1 million won)	589,153	4.3	6.4	0	0	10
	Non-bank Credit Loan Balance (1 million won)	589,153	8.4	11.2	0	4.1	19.7
E. Borrower Characteristics (within 30 days before new contract)	Bank Credit Loan Application	461,728	0.215	0.594	0	0	1
	Non-bank Credit Loan Application	461,728	1.970	1.828	1	1	4

Note: Descriptive statistics for Job (Employed) is calculated, except for borrowers whose job information is ‘other’.

of the total credit loan balance of borrowers who received credit from banks approached 24.9 million won in all financial institutions. On the other hand, according to Panel D of Table 2, the average sum of the total credit loan balance of borrowers who took out credit loans from non-banks during the same period was about 12.7 million won in all financial institutions. In other words, borrowers who took out new credit loans from banks during the period have an average of nearly twice the total credit loan balance of those who took out new credit loans from non-banks.

Figure 1 shows the proportion of loans with an interest rate of 24% or higher (henceforth, exposure) among non-bank credit loans. Therefore, exposure refers to the ratio of borrowers who are borrowing at an interest rate higher than the legal maximum level after the legal maximum interest rate was cut. Panel A shows the exposure in February-July of 2017, and Panel B shows the exposure in August of 2017 to January of 2018.

First, it is notable that more than half of the non-bank credit loans given to the low-credit class (grades 8 to 10) have an interest rate higher than 24%, which would be illegal after the statutory maximum interest rate cut. According to Panel A

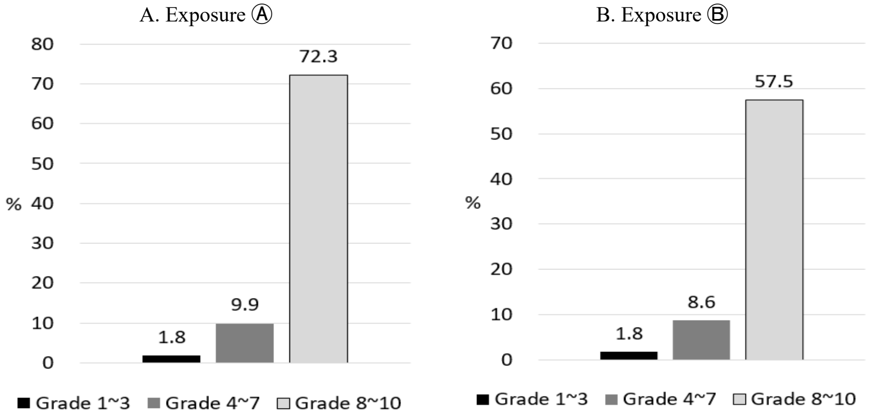

FIGURE 1. EXPOSURE IN THE NON-BANKING SECTOR BY CREDIT RATING GROUP

Note: Exposure Ⓐ is the proportion of new credit loans with interest rates higher than 24% from February of 2017 to July of 2017, and Exposure Ⓑ is the proportion of new credit loans with interest rates higher than 24% from August of 2017 to January of 2018.

(February-July 2017) of Figure 1, approximately 72.3% of non-bank credit loans taken out by borrowers in grades 8-10 have interest rates of 24% or higher. This proportion will decrease to 57.5% after August of 2017 (Panel B). Meanwhile, the exposure of the middle-credit class (grades 4-7) is 4.4~12.9 (%), and the exposure of the high-credit class (grades 1~3) is 0.9~3.6 (%). Thus, the exposure decreases sharply as the credit rating improves.

The statutory maximum interest rate cut, which was implemented in February of 2018, was officially announced in August of 2017. Accordingly, it appears that financial institutions began adjusting interest rates in advance, starting in August of 2017. Accordingly, when conducting the difference-in-difference analysis in this paper, the period before the treatment is set to the time before the official announcement, not the time when the legal maximum interest rate cut was actually executed.

B. *Credit Loan Market before and after the Reduction in the Legal Maximum Interest Rate*

This section examines the changes in the credit loan market for one year before and one year immediately after the legal maximum interest rate was cut from 27.9% to 24% in February of 2018.

The upper part of Figure 2 shows the monthly average interest rate trend of new credit loans in the banking sector for high-credit borrowers (grades 1-3), medium-credit borrowers (grades 4-7), and low-credit borrowers (grades 8-10). On the other hand, the lower part of Figure 2 shows the monthly average interest rate trend of new non-bank credit loans for borrowers with high credit scores (grades 1 to 3), borrowers with medium credit scores (grades 4 to 7), and borrowers with low credit scores (grades 8 to 10). The two vertical lines in the figure indicate when the statutory maximum interest rate cut was announced (August 2017)[4] and when the maximum

[4]Financial Services and Commission·Ministry of Justice (2017. 8. 7).

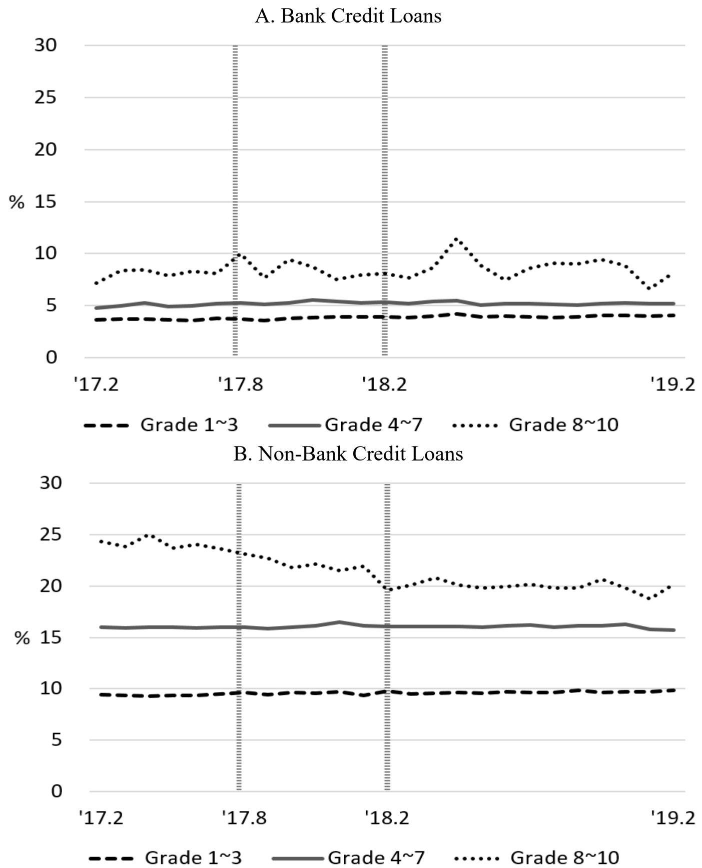

FIGURE 2. AVERAGE MONTHLY INTEREST RATES OF CREDIT LOANS BY THE CREDIT RATING GROUP IN THE BANKING AND NON-BANKING SECTORS

interest rate cut was actually carried out (February 2018).

First, the interest rates on loans to high-credit and mid-credit borrowers in the banking sector do not show significant fluctuations before and after the legal maximum rate cut. On the other hand, for low-credit borrowers, the interest rate on bank credit loans fluctuates in the range of 7-11%. The data used in this analysis excludes policy finance products such as Saitdol loans and Haetsal loans. Besides policy finance, it is rare for those with low credit scores to obtain credit loans from the banking sector; therefore, the monthly average loan interest rates vary greatly.

On the other hand, the interest rate on loans to borrowers with low credit scores from non-banking sector was reduced from 25% to 20%. In particular, looking at the change in the average interest rate of non-banking low-credit borrowers, it can be seen that the average interest rate decreased after August 2017 (when the statutory maximum interest rate cut was announced) rather than February of 2018, when the statutory maximum interest rate cut was actually implemented. This suggests that each financial institution gradually adjusted the loan contract terms starting when the legal maximum interest rate cut was announced. Accordingly, in the empirical analysis of this study, only the period before August of 2017 is regarded as the period before the treatment.

As shown in Figure 2, through the comparison before and after the legal maximum interest rate cut, the effect of the legal maximum interest rate cut on the loan interest rates of low-credit borrowers can be roughly understood. However, this cannot exclude the effect of the difference in the financial market environment over time on loan interest rates. Therefore, in this study, a difference-in-difference analysis is conducted to understand the effect of a cut on the statutory maximum interest rate. To this end, the credit loans for high-credit borrowers (credit grades 1-3) supplied by the bank are set as the control group, and the non-bank credit loans are set as the treated group. The interest rates on credit loans for high-credit borrowers supplied by banks are not likely to be affected by the statutory maximum interest rate cut.

Column A in Table 3 shows the ratio of interest rates and the number of credit loan contracts that occurred before (February to July 2017) the legal maximum interest rate was cut for each credit rating group. Meanwhile, column B in Table 3 shows the ratio of interest rates and the number of credit loan contracts that occurred after (February 2018 to February 2019) the legal maximum interest rate was cut for each credit rating group. First, looking at the change in the interest rate, we find that the average interest rate of the low-credit class declines after the legal maximum interest rate cut, as was evident from the change in the interest rate distribution earlier.

On the other hand, if we look at the change in the proportion of loan contracts by credit score before and after the statutory maximum interest rate cut, we can see that the proportion of loan contracts for the middle- and low-credit classes decreases. Theoretically, a decrease in the proportion of loan contracts is possible either due to a reduction in loan demand or a decrease in the loan supply to borrowers. Therefore, it can only be classified through a rigorous analysis whether the decrease in the ratio

TABLE 3—PROPORTION OF AVERAGE INTEREST RATES AND NUMBER OF LOAN CONTRACTS BEFORE AND AFTER THE LEGAL MAXIMUM INTEREST RATE CUT (BY CREDIT RATING SECTION)

	Credit Rating	A. Before (until 2017. 7)	B. After (from 2018. 2)	B - A
Interest Rate (%)	More than 950	4.87	5.01	0.13
	900 to 950	6.10	6.18	0.08
	850 to 900	8.05	8.21	0.15
	800 to 850	10.85	11.03	0.17
	750 to 800	13.06	13.17	0.12
	700 to 750	14.72	14.87	0.15
	600 to 700	16.96	16.90	**-0.06**
	300 to 600	20.74	18.92	**-1.82**
	Less than 300	16.17	14.46	**-1.71**
Proportion of New Loans (%)	More than 950	7.69	9.00	1.31
	900 to 950	11.94	14.08	2.13
	850 to 900	11.72	13.01	1.29
	800 to 850	13.99	14.28	0.29
	750 to 800	15.36	14.77	**-0.58**
	700 to 750	15.30	14.31	**-1.00**
	600 to 700	18.77	16.81	**-1.96**
	300 to 600	5.20	3.72	**-1.48**
	less than 300	0.04	0.04	**-0.01**

TABLE 4—DESCRIPTIVE STATISTICS OF THE LOAN APPROVAL RATE

		Observations	Mean	S.D.	p10	p50	p90
① Jan 2017 to Dec 2017	Loan Approval	885,128	0.70	0.46	0	1	1
	Bank Loan Approval	885,128	0.17	0.37	0	0	1
	Non-Bank Loan Approval	885,128	0.53	0.50	0	1	1
② Jan 2018 to Dec 2018	Loan Approval	838,091	0.65	0.48	0	1	1
	Bank Loan Approval	838,091	0.17	0.38	0	0	1
	Non-Bank Loan Approval	838,091	0.48	0.50	0	0	1

of loan contracts to those with a specific credit score is due to a demand factor or a supply factor. However, as loan interest rates for borrowers with low credit scores are declining due to the cut in the legal maximum interest rate, the demand for loans from those with low credit scores is highly likely to increase. Therefore, the decrease in the proportion of loan contracts for the low-credit class shown in Table 3 appears to be due to supply-side factors rather than demand-side factors. In other words, the findings here suggests that financial institutions may have reduced the supply of credit loan products for the low-credit class given their reduced profits due to the legal maximum interest rate cut. In this study, we analyze this more strictly through a regression difference-in-difference analysis.

Table 4 shows the proportion of borrowers who actually borrowed from banks or non-banks among potential borrowers for whom credit checks were performed in 2017 and 2018. Among those who underwent a credit check in 2017 in both the banking and the non-banking sectors, the proportion that led to a loan amounted to 70%, whereas in 2018, this proportion decreased by 5%p to 65%. Breaking these outcomes down into bank loans and non-bank loans, the success rate of bank loans after a credit check did not differ significantly between 2017 and 2018, while the success rate of non-bank loans decreased from 53% in 2017 to 48% in 2018, showing a reduction of 5%p. We find that the likelihood of a loan being rejected after a credit check to obtain a credit loan after the cut in the legal maximum interest rate increased. Moreover, we note that this phenomenon was particularly pronounced in the non-banking sector. Financial institutions may reject loans after a credit check, but some borrowers voluntarily give up borrowing after the credit check, from which a limitation of this descriptive analysis comes. Therefore, we will examine this phenomenon more rigorously via an empirical analysis.

IV. Empirical Strategy

The two main effects expected from a cut in the legal maximum interest rate are a reduction in loan interest rates and the exclusion of some borrowers from the loan market. Therefore, this chapter introduces an empirical analysis method to estimate the effect of the cut in the legal maximum interest rate on loan interest rates, loan approval rates, and the default probability.

The identification strategy used in this paper is a difference-in-difference analysis. When viewed from various angles through descriptive statistics, the cut in the statutory maximum interest rate mainly affects credit loans for low-credit borrowers in the non-banking sector. However, it has little effect on bank credit loans, and in particular, it has no effect on the credit loan market of banks for high-credit borrowers. Accordingly, the banks' credit loan market for high-credit borrowers is a good control group for estimating the effect of the legal maximum interest rate cut.

Therefore, this paper utilizes a difference-in-difference analysis, setting the banks' credit loan market for high-credit borrowers as the control group and the non-bank credit loan market as the treatment group. In a general difference-in-difference analysis, the treatment group should be observed both before and after the policy change. However, when the legal maximum interest rate is lowered, some borrowers are excluded from the market and are not observed after policy changes. Therefore, in this paper, the treatment group is limited to borrowers who are not excluded from the market even after the statutory maximum interest rate cut. Meanwhile, we also analyze the market exclusion effect of the legal maximum interest rate cut through a difference-in-difference analysis using loan approval rates in the treatment group and the control group.

The assumption known as the parallel trend assumption is the most crucial identification aspect of the difference-in-difference analysis here. The parallel trend assumption implies that in the absence of a treatment, the difference in the value of the dependent variable between the treatment group and the control group before the treatment and after the treatment would be identical. Although it is impossible directly verify to whether the parallel trend assumption is satisfied, in general, the validity of the assumption is indirectly judged by examining whether the trends of the dependent variables of the control and treated groups are parallel before the treatment.

Figure 3 shows the interest rate trends of bank credit loans for high-credit borrowers and those of non-bank credit loans according to the credit rating group to which the borrowers belong. As mentioned earlier, the bank credit loan market for high-credit borrowers becomes the control group, and the non-bank credit loan market for each credit rating group becomes the treatment group. Until the announcement of the legal maximum interest rate cut, the monthly average interest

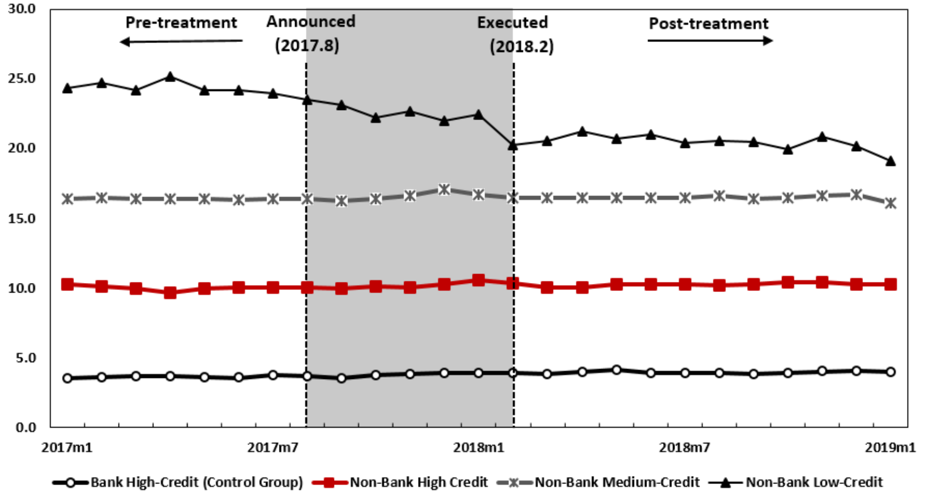

FIGURE 3. AVERAGE MONTHLY INTEREST RATES OF CREDIT LOANS (JANUARY 2017 - JANUARY 2019)

rates for each group moved in parallel. However, after August of 2017, when the statutory maximum interest rate cut was announced, the interest rates of non-banking low-credit loans started gradually to decrease. Accordingly, only the period before the announcement is used as the pre-treatment period.

Based on the parallel trend assumption introduced earlier, a regression difference-in-difference is performed in this study. The regression equation of the analysis is as follows:

$$(1)\qquad \begin{aligned} y_{ijt} &= \beta_0 + \beta_1 D_{Treated} + \beta_2 D_{After\ the\ Cut} \\ &\quad + \beta_3 D_{Treated \times After\ the\ Cut} + X_{ijt} + f_j + \varepsilon_{ijt} \end{aligned}$$

$$(2)\qquad \begin{aligned} y_{ijt} &= \beta_0 + \beta_1 D_{Non-bank\ Credit\ Check} + \beta_2 D_{After\ the\ Cut} \\ &\quad + \beta_3 D_{Non-bank\ Credit\ Check \times After\ the\ Cut} + X_{ijt} + f_j + \varepsilon_{ijt} \end{aligned}$$

$$D_{Treated} = \begin{cases} 1, & Non-bank\ Credit\ Loan \\ 0, & Bank\ Credit\ Loan \end{cases}$$

$$D_{After\ the\ Cut} = \begin{cases} 1, & After\ the\ Legal\ Maximum\ Interest\ Rate\ Cut \\ 0, & Before\ the\ Legal\ Maximum\ Interest\ Rate\ Cut \end{cases}$$

$$D_{Non-bank\ Credit\ Check} = \begin{cases} 1, & Non-bank\ Credit\ Check \\ 0, & Bank\ Credit\ Check \end{cases}$$

First, to analyze the effect of the cut in the legal maximum interest rate on non-bank credit loan interest rates and the probability of default, we use the regression difference-in-difference equation (1). The dependent variable is the interest rate when analyzing the effect of the cut in the statutory maximum interest rate on the interest rates given by non-bank credit loans. On the other hand, when analyzing the effect of the cut in the legal maximum interest rate on the default probability, the dependent variable is whether the loan is overdue for more than 90 business days (1 if a delinquency occurs for more than 90 business days after the loan is issued or 0 otherwise). As explanatory variables, we use different variables that can affect credit loan interest rates or the default probability. Specifically, these include credit scores, income, job status, age, bank credit loan balances, non-bank credit loan balances, total loan balances, total credit card usage in the preceding year, total debit card usage in the preceding year, and the Bank of Korea’s base rate. Also, we include individual financial institution fixed effects in the regression model.

Meanwhile, we use equation (2) to analyze the effect of the legal maximum interest rate cut on the loan approval rate. In this case, the dependent variable is a new loan occurrence dummy after a loan application. We can know whether a potential borrower applies for a new loan through the credit check history. KCB distinguishes between simple credit inquiries and credit checks for opening new loans. The credit check history used in this study is the latter type, sourced from the credit check record for the opening of new loans, not the simple credit inquiry type.

Of course, there may be cases where the loan is voluntarily abandoned after a credit check for a loan application. Therefore, in this difference-in-difference regression model, another identification assumption is added in addition to the parallel trend assumption. The additional assumption is that there may be a difference in the rate of the voluntary giving up of loans between high-credit borrowers and low-credit borrowers. However, we assume that this difference does not vary before and after the legal maximum interest rate cut; i.e., we utilize a parallel trend assumption for the rate of voluntarily giving up.

V. Empirical Results

This chapter introduces the empirical results of the effect of the cut in the legal maximum interest rate on credit loan interest rates, default rates, and loan approval rates for the different credit rating groups.

Table 5 is the regression difference-in-difference estimates of the effect of the reduction in the legal maximum interest rate on the loan interest rate. The control group for the regression difference-in-difference is the group of credit loans given to high-credit borrowers (credit grades 1 to 3) in the banking sector that are not affected by the cut in the statutory maximum interest rate. In column 1 of Table 5, credit loans to non-bank high-credit borrowers (grades 1 to 3) are the treatment group. In column 2, the treatment group is credit loans from non-bank medium-credit borrowers (grades 4-7). Finally, in column 3 of Table 5, the treatment group is credit loans to non-banking low-credit borrowers (grades 8-10).

As a result of the analysis, the cut in the legal maximum interest rate mainly lowered the interest rate of credit loans for low-credit borrowers in the non-banking sector. According to the third column of Table 5, the average interest rate of low-credit credit loans in the non-banking sector decreased by about 3.5%p due to the cut in the legal maximum interest rate. On the other hand, for medium-credit borrowers, the interest rate on new loans only decreased by 0.20%p during the same period, and no significant change in the loan interest rates was found for high-credit borrowers. This phenomenon is also consistent with the observation of the distribution of interest rates on non-bank credit loans by credit rating before and after the legal maximum rate cut.

The effects of the credit score, income, and base interest rate on loan interest rates also coincide with common sense. The higher the credit score and income, the lower the interest rate, and the higher the Bank of Korea base rate, the higher the loan interest rate. On the other hand, borrowers who hold higher bank loan balances have lower interest rates on new credit loans. This appears to be a phenomenon in which soft information that accumulates in financial institutions through existing transactions lowers the interest rate of new credit loans. Moreover, access to a bank loan in the past signals a borrower with low credit risk when non-banks evaluate borrower credit risk. On the other hand, the higher the non-bank credit loan balance is, the higher the interest rate also is. This may stem from the fact that the default probability is higher for borrowers who have multiple non-bank credit loans.

Due to the nature of the difference-in-difference analysis, the constant term

TABLE 5—EFFECTS OF THE LEGAL MAXIMUM INTEREST RATE CUT ON LOAN INTEREST RATES (DIFFERENCE-IN-DIFFERENCE)

Variables	Dependent Variable: Interest Rate (%)		
	(1) Grade 1~3	(2) Grade 4~7	(3) Grade 8~10
DID Effect	-0.032	-0.20***	-3.47***
	(0.027)	(0.024)	(0.040)
After Cutting the Interest Rate Cap (After the Treatment Dummy)	0.092***	0.11***	0.17***
	(0.024)	(0.024)	(0.013)
Non-Banking Credit Loan Dummy (Treatment Dummy)	5.54***	7.00***	18.5***
	(0.021)	(0.026)	(0.051)
Credit Score/100	-1.94***	-2.71***	-3.00***
	(0.016)	(0.0082)	(0.010)
Annual Income (1 Million Won)	-0.056***	-0.080***	-0.084***
	(0.0034)	(0.0036)	(0.0019)
Credit Card Usage (1 Million Won)	-0.027***	-0.059***	-0.032***
	(0.0053)	(0.0039)	(0.0035)
Debit Card Usage (1 Million Won)	-0.019***	-0.035***	-0.091***
	(0.0023)	(0.0018)	(0.0014)
BOK Base Rate (%)	0.69***	0.57***	0.38***
	(0.060)	(0.048)	(0.039)
Bank Loan Balance (1 Million Won)	-0.014***	-0.029***	-0.027
	(0.0084)	(0.0094)	(0.05)
Bank Credit Loan Balance (1 Million Won)	-0.033***	-0.044***	-0.030***
	(0.0026)	(0.0029)	(0.0014)
Non-bank Credit Loan Balance (1 Million Won)	0.010***	0.014***	0.089*
	(0.00066)	(0.00048)	(0.049)
Constants	21.8***	29.0***	6.51***
	(0.16)	(0.100)	(0.10)
Job Dummy	O	O	O
Age Group Dummy	O	O	O
Financial Institute Fixed Effect	O	O	O
Observations	265,128	610,652	155,051
R^2	0.483	0.670	0.855

Note: 1) Statistical Significance levels: ***p<0.01, **p<0.05, *p<0.1, 2) Standard errors in parentheses.

estimate refers to the average interest rate of the control group when the values of explanatory variables other than double-difference-related variables are 0. In this difference-in-difference regression, the control group is the group of high-credit borrowers in the banking sector. Looking at the constant term estimates in Table 5, the values of the high- and medium-credit classes are high, at 21.8 and 29.0, respectively. On the other hand, in the case of the low-credit class, the constant term is estimated to be 6.51, which is relatively small. This is mainly explained by the effect of credit ratings on loan interest rates. The average credit score of high-credit borrowers in the banking sector is approximately 919.9 points (9.199 if divided by 100). Therefore, multiplying the credit score coefficient estimates (-1.94, -2.71, -0.30) of each credit class by 9.199 generates corresponding values of 17.8, -24.9, and -2.8 for the respective groups. That is, for the high-credit group and the middle-credit group, a large value is subtracted from the constant term when credit scores are taken into account. However, a relatively small value is subtracted for the low-credit group. In the same way, by substituting the average value of high-credit borrowers in the banking sector into each explanatory variable and adding this value to the constant term estimate, we find similar values of 3.4, 3.4, and 3.5 for the corresponding groups. As mentioned earlier, this is the average interest rate on loans for high-credit borrowers in the banking sector in the first half of 2017.

Table 6 shows the diff-in-diff estimates of the effect of the cut in the legal maximum interest rate on the loan approval rate. According to the empirical results, when the legal maximum interest rate was reduced, the loan approval rate for low-credit borrowers decreased by about 3.6%p. On the other hand, the loan approval rate for high-credit borrowers increased by approximately 1.0%p, and the loan

TABLE 6—EFFECTS OF THE LEGAL MAXIMUM INTEREST RATE CUT ON LOAN APPROVAL RATES (DIFFERENCE-IN-DIFFERENCE)

Variables	Dependent Variable: Loan Approved Dummy		
	(1) Grade 1~3	(2) Grade 4~7	(3) Grade 8~10
DID Effect	0.010***	0.014***	-0.036***
	(0.0030)	(0.0018)	(0.0069)
After the Treatment Dummy	-0.017***	-0.017***	-0.016***
	(0.0015)	(0.0013)	(0.0015)
Non-Bank Credit Check Dummy (Treated Dummy)	-0.043	-0.024	-0.065
	(0.031)	(0.015)	(0.050)
Credit Score/100	0.006**	0.02***	0.01***
	(0.0018)	(0.0015)	(0.0019)
Annual Income (1 Million Won)	0.00022***	0.00023***	0.00024***
	(0.00003)	(0.00003)	(0.00003)
Credit Card Usage (1 Million Won)	0.00026***	0.00020***	0.00029***
	(0.00005)	(0.00003)	(0.00006)
Debit Card Usage (1 Million Won)	0.00050***	0.00030***	0.00055***
	(0.00002)	(0.00002)	(0.00003)
Bank Loan Balance (1 Million Won)	0.00006***	0.00007***	0.00007***
	(0.00001)	(0.00001)	(0.00001)
Bank Credit Loan Balance (1 Million Won)	0.0022***	0.0022***	0.0023***
	(0.0002)	(0.0002)	(0.0003)
Non-bank Credit Loan Balance (1 Million Won)	-0.0021***	-0.0013***	-0.0029***
	(0.0007)	(0.0004)	(0.0009)
Constants	0.696***	0.691***	0.739***
	(0.0160)	(0.0137)	(0.0167)
Job Dummy	O	O	O
Age Group Dummy	O	O	O
Financial Institute Fixed Effect	O	O	O
Observations	243,482	358,821	192,590
R^2	0.2099	0.2122	0.2482

Note: 1) Statistical Significance level: ***p<0.01, **p<0.05, *p<0.1, 2) Standard errors in parentheses.

approval rate for medium-credit borrowers increased by about 1.4%p. These outcomes stem from the fact that non-bank financial institutions reduce the supply of credit loans for low-credit borrowers, who are no longer generating profits. On the other hand, non-bank financial institutions increase the supply of credit loans for high-credit and medium-credit borrowers, as these loans can still generate profits after the cut in the legal maximum interest rate.

On the other hand, the effects of the credit score, income, credit card usage, and debit card usage of borrowers on their loan approval rates is also consistent with a priori outcomes. The higher the credit score, the higher the loan approval rate, and the higher the income, the higher the loan approval rate. Lastly, at the time of the loan review, the greater the bank loan balance, the higher the loan approval rate, whereas the greater the non-bank credit loan balance, the lower the loan approval rate.

Table 7 shows the diff-in-diff estimates of the effect of a cut in the legal maximum

TABLE 7—EFFECTS OF THE LEGAL MAXIMUM INTEREST RATE CUT ON LOAN APPROVAL RATES (ONLY CONSIDERING CARD AND CAPITAL COMPANIES AND SAVINGS BANKS)

Variables	Dependent Variable: Loan Approved Dummy		
	(1) Grade 1~3	(2) Grade 4~7	(3) Grade 8~10
DID Effect	0.007**	0.014***	-0.048***
	(0.0024)	(0.0019)	(0.0069)
After the Treatment Dummy	-0.017***	-0.017***	-0.016***
	(0.0018)	(0.0014)	(0.0015)
Non-Bank Credit Check Dummy (Treated Dummy)	-0.048***	-0.042***	-0.086***
	(0.0025)	(0.0017)	(0.028)
Credit Score/100	0.01***	0.02***	0.01***
	(0.002)	(0.002)	(0.002)
Annual Income (1 Million Won)	0.0025***	0.0018***	0.0024***
	(0.0003)	(0.0003)	(0.0003)
Credit Card Usage (1 Million Won)	0.00043***	0.00040***	0.00030***
	(0.00005)	(0.00004)	(0.00006)
Debit Card Usage (1 Million Won)	0.00066***	0.00030***	0.00058***
	(0.000021)	(0.000016)	(0.000025)
Bank Loan Balance (1 Million Won)	0.00007***	0.00008***	0.00007***
	(0.00001)	(0.00001)	(0.00001)
Bank Credit Loan Balance (1 Million Won)	0.0026***	0.0023***	0.0023***
	(0.0002)	(0.0002)	(0.0003)
Non-bank Credit Loan Balance (1 Million Won)	-0.0040***	-0.0030***	-0.0031***
	(0.0006)	(0.0004)	(0.0009)
Constants	0.683***	0.653***	0.736***
	(0.0186)	(0.0155)	(0.0168)
Job Dummy	O	O	O
Age Group Dummy	O	O	O
Financial Institute Fixed Effect	O	O	O
Observations	189,436	329,394	191,888
R^2	0.2293	0.1894	0.2527

Note: 1) Statistical Significance level: ***p<0.01, **p<0.05, *p<0.1, 2) Standard errors in parentheses.

interest rate on the loan approval rate, but the treatment group is limited to card companies, capital companies, and savings banks. According to the analysis, with the 3.9%p cut the legal maximum interest rate from 27.9% to 24%, the approval rate of non-bank credit loans for low-credit borrowers decreased by about 4.8%p. On the other hand, the approval rate for non-bank credit loans for high-credit borrowers increased by about 0.7%p, while that for medium-credit borrowers increased by approximately 1.4%p. This result is similar to the result of the previous analysis (Table 6) in which all non-bank financial institutions were included as the treatment group.

In particular, as the legal maximum interest rate is lowered, the supply of credit loans for low-credit borrowers, who no longer generate profits, is reduced, and the supply of credit loans for high-credit borrowers and medium-credit borrowers is increased.

Table 8 shows the diff-in-diff estimates of the effect of the cut in the legal maximum interest rate on the default probability. The analysis results indicate that the default rate decreased after the legal maximum interest rate was cut. However, this only affected borrowers who successfully obtained a loan despite the cut in the legal maximum interest rate.

The effect was particularly apparent in the low-credit class. As the statutory maximum interest rate decreased from 27.9% to 24% in February of 2018, no significant change was observed in the default probability for high-credit borrowers. On the other hand, the default probability of medium-credit borrowers decreased by about 0.24%p on average. Considering that the average default probability of non-bank medium-credit borrowers is close to 4.44%, the statutory maximum rate cut reduced the default probability of medium-credit borrowers by about 5.4%. On the other hand, the default probability of low-credit borrowers decreased by a whopping 2.8%p. This means that the default probability decreased by nearly 21% when considering the average default probability (13.6%) of non-bank low-credit borrowers.

The effects of various explanatory variables, such as the credit score, income, amount of credit card use in the previous year, and debit card use amount in the previous year on the default probability also coincides with common sense. The higher the credit score, the lower the default probability, and the higher the income, the lower the default probability. On the other hand, when the Bank of Korea base rate is high, the probability of a default decreases because the base rate decreases and the probability of a default increase during an economic downturn.

According to Table 5, the cut in the legal maximum interest rate mainly led to loan rate cuts for low-credit borrowers in the non-banking sectors. In addition, according to Table 8, the default probability of low-credit borrowers significantly decreased due to the reduction of the legal maximum interest rate. Taken together, for low-credit borrowers who successfully took out loans despite the legal maximum rate cut, the loan interest rate was reduced. As a result, the monthly repayment burden decreased, thereby reducing the probability of a default.

TABLE 8—EFFECTS OF CUTTING THE INTEREST RATE CAP ON THE DEFAULT RATE (DIFFERENCE-IN-DIFFERENCE)

Variables	Dependent Variable: Default Dummy		
	(1) Grade 1~3	(2) Grade 4~7	(3) Grade 8~10
DID Effect	0.00030	-0.0024**	-0.028***
	(0.00074)	(0.0010)	(0.0023)
After Cutting Down Interest Rate Cap (After the Treatment Dummy)	-0.00062	0.0011	-0.00023
	(0.00065)	(0.0011)	(0.00077)
Non-Banking Credit Loan Dummy (Treatment Dummy)	0.0078***	0.047***	0.10***
	(0.00058)	(0.011)	(0.0029)
Credit Score/100	-0.0061***	-0.019***	-0.0084***
	(0.00044)	(0.00035)	(0.00058)
Income (1 Million Won)	-0.00014	-0.00061***	-0.00022**
	(0.000092)	(0.00015)	(0.00011)
Credit Card Usage (1 Million Won)	-0.000043***	-0.00017***	-0.00068***
	(0.000015)	(0.000017)	(0.00020)
Debit Card Usage (1 Million Won)	0.000011	0.000076***	0.000054
	(0.00063)	(0.0000078)	(0.00082)
BOK Base Rate (%)	-0.0023	-0.0090***	-0.0034
	(0.0017)	(0.0021)	(0.0023)
Bank Loan Balance (1 Million Won)	-0.00016***	-0.00065***	-0.00010***
	(0.000023)	(0.000041)	(0.000029)
Bank Credit Loan Balance (1 Million Won)	0.00037***	0.00046***	6.0e-06
	(0.000071)	(0.00012)	(8.3e-06)
Non-bank Credit Loan Balance (1 Million Won)	0.0021***	0.0040***	0.0022
	(0.00018)	(0.0002)	(0.0028)
Constants	0.069***	0.18***	0.088***
	(0.0045)	(0.0043)	(0.0060)
Job Dummy	O	O	O
Age Group Dummy	O	O	O
Financial Institute Fixed Effect	O	O	O
Observations	265,128	610,652	155,051
R^2	0.006	0.015	0.079

Note: 1) Statistical Significance level: ***p<0.01, **p<0.05, *p<0.1, 2) Standard errors in parentheses.

VI. Conclusion

This study showed that the loan interest rates of low-credit borrowers who use the non-banking sector fell significantly due to the legal maximum interest rate cut in February of 2018. On the other hand, no significant decline was found in the loan interest rates of high-credit and low-credit borrowers. In addition, the default rates of low-credit borrowers using the non-banking sector decreased significantly due to a reduction in the monthly repayment burden caused by the reduced loan interest rates.

However, this phenomenon is limited to borrowers who could still obtain a loan despite the cut in the legal maximum interest rate. In fact, according to the results of the analyses here, the loan approval rates of low-credit borrowers using the non-banking sector decreased significantly due to the cut in the legal maximum interest rate.

Many borrowers receiving loans at a level similar to the legal maximum interest rate are likely to be from vulnerable classes with low incomes or low credit ratings. Considering this, the findings here suggest that policy supplements are necessary for borrowers excluded from the market due to the reduced legal maximum interest rate. In particular, as the statutory maximum interest rate is lowered continuously, the number of borrowers excluded from the market due to further cuts in the statutory maximum interest rate is highly likely to increase.

In fact, when policy authorities recently lowered the legal maximum interest rate from 24% to 20%, taking into account the possibility of excluding vulnerable borrowers from the market, they implemented follow-up measures, such as providing policy loans for low-income borrowers. Therefore, in future research, examining whether these follow-up measures following the cut in the legal maximum interest rate sufficiently provided a safety net is necessary. These points represent limitations of this study and are left as future research tasks.

On the other hand, lowering the legal maximum interest rate reduces financial institutions' profits. Therefore, a cut in the statutory maximum interest rate creates a new market environment for financial institutions, providing incentives to develop new markets. In particular, compared to the high-interest-rate loan market, where obtaining a loan has become relatively difficult due to the cut in the legal maximum interest rate, the medium-rate market can be a new avenue for card companies, capital companies, and savings banks.

Two major problems have been pointed out as factors that prevented the middle-interest rate credit loan market from being activated. The first problem is that there is serious information asymmetry between financial providers and consumers, and the second point is the lack of incentives to supply medium-rate credit loans due to limited competition from financial institutions. In particular, according to a previous study by Kim (2019), it is highly likely that the lack of incentives to supply medium-rate credit loans has been the greatest obstacle to the vitalization of medium-rate loans. Therefore, by introducing an appropriate incentive system when the statutory maximum interest rate is reduced, it would be possible to absorb some of the borrowers who could potentially be excluded from the market due to the statutory maximum interest rate cut. This is a case in which the government could intervene more actively through policy compared to the situations in other countries, and it is a necessary measure in the current situation where the legal maximum interest rate is relatively low compared to those in other countries.

APPENDIX

Table A1 presents the analysis result without including any explanatory variables other than the variables for the difference-in-difference for the robustness check of Table 5. Similar to the results in Table 5, the cut in the legal maximum interest rate mainly leads to a reduction in the loan rates for low-credit borrowers in the non-banking sector. On the other hand, due to the nature of the difference-in-difference analysis, the estimate of the constant term (β_0) refers to the average interest rate of credit loans for high-credit borrowers in the banking sector.

Table A2 is the result of an additional robustness check for Table 5 and confirms once again that the cut in the legal maximum interest rate mainly lowers the loan rates for low-credit borrowers in the non-banking sector.

Table A3 is the robustness check result for Table 8, confirming once again that lowering the legal maximum interest rate mainly reduces the default rate for low-credit borrowers in the non-banking sector. As shown in Table 5, the cut in the legal maximum interest rate leads to a lower loan interest rate for low-credit borrowers. As a result, the monthly repayment burden is reduced. A reduction in the monthly repayment burden can lead to a reduction in the default rate. Therefore, the results in Table 8 are consistent with our expectations.

TABLE A1—EFFECTS OF THE CUTTING THE INTEREST RATE CAP ON INTEREST RATES (DID)

Variables	Dependent Variable: Interest Rate (%)		
	(1) Grade 1~3	(2) Grade 4~7	(3) Grade 8~10
DID effect	-0.064***	-0.28***	-3.54***
	(0.024)	(0.023)	(0.037)
After cutting the Interest Rate Cap (After the Treatment Dummy)	0.24***	0.24***	0.24***
	(0.017)	(0.020)	(0.0074)
Non-Banking Credit Loan Dummy (Treated Dummy)	6.36***	12.7***	20.1***
	(0.017)	(0.016)	(0.024)
Constants	3.72***	3.72***	3.72***
	(0.012)	(0.014)	(0.0052)
Observations	431,978	805,364	228,396
R^2	0.390	0.601	0.821

Note: 1) Statistical Significance level: ***p<0.01, **p<0.05, *p<0.1, 2) Standard errors in parentheses.

TABLE A2—EFFECTS OF CUTTING THE INTEREST RATE CAP ON INTEREST RATES (DID)

Variables	Dependent Variable: Interest Rate (%)	
	(1)	(2)
DID Effect	-0.19*** (0.025)	
DID Effect (Credit grade 1-3)		-0.0067 (0.033)
DID Effect (Credit grade 4-7)		-0.19*** (0.025)
DID Effect (Credit grade 8-10)		-3.39*** (0.100)
After cutting down Interest Rate Cap (After Treatment Dummy)	0.10*** (0.025)	0.10*** (0.025)
Non-bank Credit Loan Dummy (Treated Dummy)	5.86*** (0.021)	
Non-bank Credit Loan Dummy (Credit grade 1-3)		5.29*** (0.025)
Non-bank Credit Loan Dummy (Credit grade 4-7)		6.81*** (0.026)
Non-bank Credit Loan Dummy (Credit grade 8-10)		7.17*** (0.074)
Credit Score/100	-3.10*** (0.0057)	-2.76*** (0.0082)
Annual Income (1 Million Won)	-0.072*** (0.0034)	-0.072*** (0.0034)
Credit Card Usage (1 Million Won)	-0.011*** (0.0037)	-0.013*** (0.0037)
Debit Card Usage (1 Million Won)	-0.021*** (0.0017)	-0.025*** (0.0017)
BOK Base Rate (%)	0.59*** (0.046)	0.60*** (0.045)
Bank Loan Balance (1 Million Won)	-0.034*** (0.0087)	-0.030*** (0.0087)
Bank Credit Loan Balance (1 Million Won)	-0.013*** (0.0028)	-0.012*** (0.0028)
Non-bank Credit Loan Balance (1 Million Won)	0.055*** (0.00045)	0.075*** (0.00045)
Constants	33.0*** (0.080)	29.7*** (0.097)
Job Dummy	O	O
Age Group Dummy	O	O
Financial Institute Fixed Effect	O	O
Observations	736,379	736,379
R^2	0.629	0.633

Note: 1) Statistical Significance level: ***p<0.01, **p<0.05, *p<0.1, 2) Standard errors in parentheses.

TABLE A3—EFFECTS OF CUTTING THE INTEREST RATE CAP ON THE DEFAULT RATE (DID)

Variables	Dependent Variable: Default Dummy	
	(1)	(2)
DID Effect	-0.0022**	
	(0.00099)	
DID Effect (Credit grade 1-3)		0.00043
		(0.0013)
DID Effect (Credit grade 4-7)		-0.0024**
		(0.0010)
DID Effect (Credit grade 8-10)		-0.027***
		0.0011
After cutting down Interest Rate Cap (After Treatment Dummy)	0.0012	0.0011
	(0.0010)	(0.0010)
Non-bank Credit Loan Dummy (Treated Dummy)	0.0012	
	(0.00085)	
Non-bank Credit Loan Dummy (Credit grade 1-3)		0.0040***
		(0.00099)
Non-bank Credit Loan Dummy (Credit grade 4-7)		0.0024**
		(0.0010)
Non-bank Credit Loan Dummy (Credit grade 8-10)		0.054***
		(0.0030)
Credit Score/100	-0.018***	-0.018***
	(0.00023)	(0.00033)
Annual Income (1 Million Won)	-0.00078***	-0.00059***
	(0.00014)	(0.00014)
Credit Card Usage (1 Million Won)	-0.00018***	-0.00015***
	(0.000015)	(0.000015)
Debit Card Usage (1 Million Won)	0.000067***	0.000063***
	(0.0000070)	(0.0000070)
BOK Base Rate (%)	-0.0091***	-0.0089***
	(0.0018)	(0.0018)
Bank Loan Balance (1 Million Won)	-0.00055***	-0.00057***
	(0.000035)	(0.000035)
Bank Credit Loan Balance (1 Million Won)	0.00052***	0.00055***
	(0.00011)	(0.00011)
Non-bank Credit Loan Balance (1 Million Won)	0.0035***	0.0037***
	(0.00018)	(0.00018)
Constants	0.17***	0.18***
	(0.0032)	(0.0039)
Job Dummy	O	O
Age Group Dummy	O	O
Financial Institute Fixed Effect	O	O
Observations	736,379	736,379
R^2	0.017	0.018

Note: 1) Statistical Significance level: ***p<0.01, **p<0.05, *p<0.1, 2) Standard errors in parentheses.

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The Impact of Tax Treaties on Foreign Direct Investment: The Evidence Reconsidered

By SIWOOK LEE AND DAEYONG KIM*

This paper reconsiders the empirical evidence of the relationship between tax treaties and FDI using U.S. outbound FDI to 78 countries over the period of 2007–2018. Unlike previous studies, we explicitly consider differences in the tax environments of recipient economies, including their tax-haven status, transfer pricing rules, CFC rules and anti-avoidance regulations, in our estimations. Our results confirm the importance of controlling for country-specific tax environments, especially the tax-haven status and transfer pricing rules. We find that tax treaties positively contribute to FDI inflows in developing countries, while they have no statistically significant impacts on OECD countries. Recently signed tax treaties still foster FDI but less than older ones do. Finally, our results indicate, all other things being equal, that the weaker the transfer pricing regulations, the greater the amount of U.S. direct investment into a non-OECD economy.

Key Word: Tax Treaties, Foreign Direct Investment, Tax Havens, Transfer Pricing, Tax Regulations
JEL Code: O11, O47, C21, F4

I. Introduction

Foreign direct investment (FDI hereafter) is generally regarded as an important driver of economic growth, a composite package of investment resources, technological know-how and managerial expertise (de Mello, 1997). Recognizing this, many countries compete to attract FDI by providing favorable incentives to foreign investors. In addition, countries enter into bilateral and/or multilateral economic agreements, such as tax treaties, investment treaties, and preferential trade agreements, to assure foreign investors that they adhere to global norms in trade and investment practices.

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Among these agreements, tax treaties are aimed at ameliorating tax-related impediments to cross-border trade and investment. While the primary objective of tax treaties is to avoid the double taxation of income by more than one jurisdiction, they also cover other issues, including the prevention of tax evasion, excessive taxation, and tax discrimination. Since the League of Nations initially proposed the modern tax treaty model in 1928, such agreements have proliferated worldwide, and currently more than 3,000 bilateral tax treaties are in effect. The pace at which tax treaties are being established has even accelerated since the mid-1990s (Leduc and Michielse, 2021).

Despite the proliferation of tax treaties, there has been a growing sense of skepticism regarding their effectiveness, especially in recent years (Kysar, 2020; Brooks and Krever, 2015). The contemporary architecture of bilateral tax treaties largely preserves the principles and structure of the League of Nations model (Kobetsky, 2011). That model was developed when international transactions were usually carried out in a tangible form, but the world economy has changed considerably since then. In the face of accelerated globalization and digitalization, the roles of multinational companies have grown and international transactions increasingly take place in intangible forms.

At the present time, the digital transformation of cross–border transactions has contributed to the emergence of various techniques of tax avoidance or tax evasion across countries. Multinational enterprises can abuse tax treaties by ‘treaty shopping’ to avoid taxation, causing what are known as “double non-taxation” problems. Consequently, cross-border taxation issues are becoming more complex and the current tax treaty system has not sufficiently responded to these changes.

Tax treaties are designed to handle double taxation mainly by limiting a source country’s taxation on income not derived via a permanent establishment within the country (Brooks and Krever, 2015; Petkova *et al*., 2020).[1] In other words, they shift the taxing rights from the source country to the investor’s country of residence at the expense of tax revenue to the former. Hence, if capital inflows are greater than capital outflows for an economy, as is true for most developing countries, the cost of lost tax revenue may outweigh the potential benefits of forgoing taxing rights, unless tax treaties induce a sufficient level of FDI inflow and other externalities that create jobs and sustained economic growth.[2]

Several researchers have empirically investigated the tax treaty–FDI nexus but a consensus on whether tax treaties increase FDI flows remains elusive. Such mixed findings have contributed to the controversy over the validity of tax treaties. Against this backdrop, this paper aims to reconsider the empirical evidence of the impact of tax treaties on FDI. While the tax environments of recipient countries, such as their local tax systems and regulations, are inarguably a decisive factor in investment decisions, previous studies of the tax treaty–FDI nexus failed to consider country-specific tax environments as a determinant of FDI, leading to omitted variable bias

[1]The source country refers to the country that hosts the inward investment, while the residence country is the investor’s country of residence.

[2]Given the heightened suspicion of the unfavorable revenue impacts of tax treaties, a number of countries have recently canceled or restructured their existing tax treaties, especially those with tax-haven countries. For instance, Mongolia canceled its tax treaties with Luxembourg, the Netherlands and the UAE in 2012. In 2014, Uganda also suspended new treaty negotiations. Cyprus, Malta and Luxembourg recently agreed to amend the terms of their tax treaties with Russia. The key change is an increase in the withholding tax rates for dividends and interest.

in their estimations.[3]

To explain this, it is helpful to consider tax treaties with tax-haven countries. Tax havens do not tax foreign-sourced income; thus, tax treaties do not in fact affect their tax systems. This implies that a country that enters into a tax treaty with a tax haven could potentially give up a significant amount of tax revenue. Consequently, countries may be reluctant to sign tax treaties with tax havens. Even tax havens may hesitate to conclude tax treaties due to the built-in obligation to provide tax information. For similar reasons, the likelihood of a tax agreement can vary depending on the tax regulations of partner countries. In this context, our analysis explicitly controls for the specific tax environments of recipient countries, including their tax-haven status, the quality of their tax avoidance regulations, controlled foreign corporation rules (CFC rules hereafter) and transfer price rules. We find that the mixed evidence from previous studies may stem to some extent from omitted variable bias in the estimations.

In addition, while most of the existing literature analyzes the effectiveness of tax treaties during the 1980s and 1990s, this study deliberately targets a more recent period, 2007–2018. One of the striking findings in the existing literature is that tax treaties signed more recently tend to have a more negative impact on FDI flows than older treaties. Blonigen and Davis (2004) and Egger *et al*. (2006) interpret this as evidence that new tax treaties or the revision of old ones may reduce FDI flows, as they contain more sophisticated incentive schemes to limit FDI for tax avoidance purposes. If this claim is correct, our estimates for the more recent period should show an even stronger negative impact of tax treaties on FDI. It is also plausible that the accelerated pace of globalization and digitalization may further undermine the effectiveness of tax treaties.

The paper is organized as follows. Section II provides a brief literature review on the relationship between tax treaties and FDI. Section III discusses our estimation strategy and describes the data used in this study. In section IV, we present the estimation results based on our model. Some concluding remarks are provided in section V.

II. Literature Review

As mentioned above, the empirical evidence on the effectiveness of tax treaties is largely mixed. Blonigen and Davis (2004) analyze the effects of tax treaties on both U.S. inward and outward FDI for the period of 1980–1999, separating tax treaties signed before the sample period from those signed during the sample period. They find little evidence of an impact of tax treaties on FDI. In addition, their analysis indicates that new tax treaties may even have a negative impact on U.S. direct investment activities abroad. Blonigen and Davis (2005) find similar results for OECD countries over the period of 1982–1992, suggesting that tax treaties serve as a mechanism for reducing tax evasion rather than boosting foreign investment. Egger *et al*. (2006) employ the propensity score matching (PSM) method to analyze the

[3]Please see Feld and Heckemeyer (2011) for an excellent meta-study of the relationship between taxation and FDI.

impact of bilateral direct investments on OECD countries during the period of 1982–1992. They also show that tax treaties have a negative effect on direct investment abroad. Likewise, Davis (2003) examines 20 cases of U.S. tax treaty revisions during the period of 1966–2000 and reports that treaty renegotiations do not increase FDI.

On the other hand, Stein and Daude (2007) report that with regard to OECD direct investment abroad over the period of 1997–1999, tax treaties affected which recipient countries were chosen. Neumayer (2007) analyzes a sample of developing countries and finds that FDI stocks were on average approximately 20 percent higher if a treaty was concluded during the sample period. However, these effects were confined to medium-income countries. Neumayer's results also suggest that countries that have more tax treaties with major developed countries have greater FDI inflows. Recently, Lejour (2014) applied a propensity score matching estimation to a sample of 34 OECD countries over the period of 1985–2011. In the estimation, tax treaties are instrumentalized using exogenous geographic variables to control for endogeneity issues. Contrary to Blonigen and Davis (2005) and Egger *et al*. (2006), Lejour (2014) shows that tax treaties significantly contribute to FDI, and new treaties have an especially large effect. Employing a quantile treatment model with U.S. FDI data over the period of 1988–1999, Kumas and Millimet (2018) suggest that the impacts of tax treaties on FDI differ depending on the extent of FDI activity at the time of treaty conclusion. Specifically, tax treaties increase FDI at lower quantiles of the FDI distribution but decrease FDI at upper quantiles.

Potential reasons for the mixed findings on the effectiveness of tax treaties are as follows. First, tax treaties aim to prevent both double taxation and tax evasion. Consequently, tax treaties have conflicting effects in that they promote direct investment by preventing double taxation but reduce FDI through their anti–tax-avoidance provisions.[4] Hence, empirical studies may observe negative impacts of tax treaties if they reduce the inflow of new direct investments for tax avoidance purposes more than they promote investment through the prevention of double taxation (Blonigen and Davis, 2004; Egger *et al*., 2006).

Second, as Baker (2014) argues, developed countries are equipped with organized legal frameworks and policies to prevent double taxation and tax avoidance. This mitigates the major benefits of signing tax treaties with partner countries, meaning that the effect of tax treaties on developed countries could be minimal. However, this finding does not explain why the effect of tax treaties could be negative.

Third, it cannot be ruled out that the ambiguous evidence stems from estimation problems that are inherent to the existing studies. This paper pays special attention to potential omitted variable bias in previous studies. As described above, the exclusion of variables related to countries' specific tax environments from the regression analyses may mean that tax treaties are correlated with the error terms, resulting in biased and inconsistent estimates.

Fourth, most of the aforementioned empirical studies treat tax treaties as a binary variable – regardless of whether a tax treaty exists or not – without considering differential attributes among these treaties. It is highly plausible that the effectiveness of individual tax treaties may not be the same, especially considering the possibility

[4] An increase in cross-border FDI flows is also expected when prevention of excessive taxation and tax discrimination are stipulated in tax treaties.

of treaty shopping.

In this context, there has been recently a growing strand of research that employs network analysis techniques and/or micro-level data to identify the specific conduits through which tax treaties contribute to FDI (van't Riet and Lejour, 2018; Hong, 2018; Petkova *et al*., 2020). For instance, Petkova *et al*. (2020) find that tax treaties that offer investors a financial advantage both over domestic law and the entire treaty network would increase FDI, while others do not. Similarly, Hong (2018) demonstrates the existence of tax-minimizing direct routes that contribute to FDI. Such a differential impact of tax treaties may not be well captured in the existing literature, which relies on a binary treatment of tax treaties.

III. Empirical Strategy and Data Description

A. *Empirical Strategy*

This paper examines the impact of tax treaties on U.S. outbound FDI destined to 78 countries over the period of 2007–2018. In our empirical analysis, we consider the gravity model as a benchmark and augment it with other key predictors. Our conceptual framework can be summarized by the following equation:

$$(1) \qquad fdi_{it} = f(GRAV_{it}, KNOW_{it}, T\_COST_{it}, Tax\_treaty_{it}, Z_{it}, TAX_{it})$$

where fdi_{it} represents the volume of U.S. FDI destined to a country i in year t, $GRAV_{it}$ is the vector of gravity variables, such as GDP (GDP_{it}) and the physical distance between the U.S. and country i ($dist_i$). $KNOW_{it}$ represents the extent of knowledge capital of recipient country i, T_COST_{it} is the bilateral trade cost between the U.S. and country i. Tax_treaty_{it} is a dummy variable indicating whether a bilateral tax treaty is in effect between i and the U.S. in year t, and Z_{it} is the vector of other bilateral and multilateral economic agreements in effect at time t. TAX_{it} represents the vector of tax environment variables for recipient country i.

Among the gravity variables, GDP is expected to be a robust determinant of FDI, as horizontal FDI is often destined to countries that boast large markets and great purchasing power.[5] While physical distance is inarguably a crucial factor that determines trade flows as it is indicative of trade costs, its impact on FDI flows is ambiguous. On one hand, physical distance could be an indicator of costs related to FDI activities, such as transport, communication, market search, among others, implying that distance could affect FDI flows. On the other hand, firms often locate production in direct proximity to a foreign market to avoid distance-related costs. Therefore, all other factors being equal, it is plausible that the greater the distance, the greater the horizontal FDI incentives for firms.

As shown in Equation (1), we consider the degree of knowledge capital of a

[5]Horizontal FDI represents the overseas production of products and services similar to those a firm produces in its home market. It occurs when a firm directly serves a foreign market to avoid distance-related costs associated with exports.

recipient country as a determinant of FDI flows. Markusen (2007) presents a theoretical model showing that FDI provides knowledge-intensive services to recipient countries for whom developing their own knowledge-intensive inputs would be cost prohibitive. At the same time, the absorptive capacity of recipient countries matters when multinationals decide on a location for FDI. In this paper, we use the relative level of human capital (HC_{it}) and total factor productivity (TFP_{it}) for country i compared to the U.S. as a proxy for its absorptive capacity.

Our trade cost variable, T_COST_{it}, is the geometric average of bilateral tariff rates between the U.S. and country i. High tariff rates induce foreign firms to avoid tariff barriers by locating their production within the destination market. This implies that, other things being equal, we can expect a positive relationship between tariff rates and FDI. However, the likelihood of such tariff-jumping FDI is influenced by other factors, including differential production costs, relocation costs and local demand conditions. Furthermore, high tariff rates may also be related to the inclination to protect domestic producers. Therefore, the estimated coefficient of this trade cost variable could be ambiguous.

The vector of Z_{it} includes dummy variables for tax information exchange agreements (TIEA hereafter), free trade agreements (FTA hereafter) and WTO membership (WTO hereafter). TIEA allows for the exchange of tax information to address harmful tax practices. It can complement tax treaties or can be used by countries for whom taxes on income or profits are low or even zero, making tax treaties inappropriate.[6] FTAs often contain investment provisions to foster FDI flows between member countries. More importantly, FTAs and WTO membership can assure foreign investors that recipient countries adhere to global norms in trade and investment practices.

Finally, TAX_{it} consists of several variables to capture country-specific tax environments. These include tax haven status (tax_hvn_i), transfer pricing rules (trn_prc_{it}), controlled foreign corporation (cfc_{it}) rules, anti-avoidance regulations ($anti\_avd_{it}$) and corporate income tax rates (tax_cp_{it}). Tax havens tend to attract a large amount of FDI relative to their market size, especially by enabling multinationals to attain tax rates that are effectively close to zero. These impacts are not confined to tax havens but indeed apply to all countries with which an investing country has a tax treaty. Therefore, tax haven status should definitely be included in the estimation of a tax treaty–FDI nexus. Transfer pricing rules require firms to establish prices based on similar transactions between unrelated parties, and CFC rules prevent the artificial diversion of profits to a related company to minimize tax liabilities. Anti-avoidance regulations are designed to discourage or prevent tax avoidance in advance rather than addressing it after the fact.

Taking the abovementioned discussions into account, our estimation specification is as follows:

[6]Kysar (2020) and Brooks and Krever (2015) claim that TIEA could be a good substitute for tax treaties that avoid changing the taxing jurisdiction. On the other hand, Sheppard (2009) casts doubt on the effectiveness of the current TIEA architecture.

$$\begin{aligned}(2)\qquad ln\_fdi_{it} \equiv \beta_0 + \beta_1 \ln\_GDP_{it} + \beta_2 \ln\_dist_i + \beta_3 TFP_{it} + \beta_4 \ln\_HC_{it} \\ + \beta_5 T\_Cost_{it} + \beta_6 Tax\_treaty_{it} + Z_{it}^{'}\psi + TAX_{it}^{'}\xi + \mathrm{T}_t^{'}\varpi + \varepsilon_{it}\end{aligned}$$

where T_t represents the vector of year dummies; ψ, ξ and ϖ are the vectors of the coefficients; and ε_{it} is the error term. While we employ several different estimators, including ordinary least squares (OLS hereafter), a fixed effects estimator, and a random effects estimator, we consider the potential impact of time-invariant unobserved heterogeneity across countries using a fixed effects estimator. In addition, as a robustness check, we also employ Arellano and Bond (1991)'s GMM estimation method to control for potential endogeneity bias.

B. *Data Description*

Our country panel data come from various data sources, including the U.S. Bureau of Economic Analysis (BEA hereafter), the Tax Treaty database, Penn World Tables, the ESCAP-WB trade cost database and CEPII, as presented in Table 1. The set of tax haven countries is created based on Garcia-Bernardo *et al*. (2017), and the tax environment variables come from Schanz *et al*. (2017). Our dependent variable is the U.S. outbound stock for each recipient country obtained from the U.S. BEA.[7] We use the CEPII dataset as the source of the gravity variables. TFP_{it} is measured as the PPP-adjusted TFP level relative to that of the U.S. This measure, along with the human capital index (HC_{it}), comes from Penn World Tables 10.0. Moreover, the trade cost variable, proxied by the geometric average of bilateral tariff rates between the U.S. and country i, comes from the ESCAP-WB trade cost database.[8]

For tax-related regulation variables such as anti-avoidance regulations, transfer pricing rules and CFC rules, the higher the value of these variables, the weaker the regulations of recipient country i. For example, if transfer pricing rules are not well established or are not applied appropriately for country i, then the related dummy variable has a value of 1, implying that the likelihood of tax evasion or avoidance increases. Therefore, if the estimated coefficients for these variables are positive for country i, it means that the weaker the relevant regulations are to prevent tax evasion, the greater the amount of U.S. direct investment is into the country. In addition, as countries with low tax rates are more likely to attract FDI, we also include the corporate tax rate (tax_cp_{it}) of recipient country i in the estimation. Schanz *et al*. (2017) construct this variable in the following way: first, they note the maximum observed tax rate among all the countries in their sample data, after which

[7]Missing data of FDI stocks for recipient countries account for less than 3% of the total observations in our sample, and these are left out mostly to protect the confidentiality of individual companies.

[8]Specifically, the bilateral tariff cost is measured by $\sqrt{(1+tariff_{ijt})(1+tariff_{jit})}$. The ESCAP-WB trade cost database contains another bilateral trade cost index proposed by Anderson and van Wincoop (2003). This index captures not only tariff-related costs but also other indirect and direct costs associated with bilateral trade. We do not adopt this index in our study for the following reasons. First, the magnitude of the trade costs for this measure is sensitive to underlying assumptions on the elasticity of substitution (Novy, 2013). More importantly, by construction, this index is highly correlated with other explanatory variables in Equation (1), including GDP, physical distance and FTA, which may change the statistical property of our estimations. For this reason, we use the bilateral tariff cost as a proxy for the exogenous trade cost.

TABLE 1—DATA SOURCES AND DESCRIPTION

Variable	Description	Data Source
FDI	Logged values of US outbound FDI flow at a historical-cost basis	US-Bureau of Economic Analysis
GDI	Logged values of GDP (US$) at current prices	CEPII Database
Distance	Logged values of population-weighted distances	CEPII Database
TFP	TFP level at current PPP (U.S. =1)	Penn World Tables 10.0
Human Capital	Human capital index based on years of schooling and returns to education (U.S. =1)	Penn World Tables 10.0
Trade Cost	The geometric average of bilateral tariff rates between the U.S. and country *i*	ESCAP-WB trade cost database
Tax Treaty	= 1 for tax treaty in effect, otherwise 0	Tax Treaty Database
TIEA	= 1 for TIEA in effect, otherwise 0	Tax Treaty Database
FTA	= 1 for FTA in effect, otherwise 0	CEPII Database
WTO Membership	= 1 if GATT/WTO member, otherwise 0	CEPII Database
Tax Haven	= 1 for tax haven countries, otherwise 0	Garcia-Bernardo *et al*. (2017)
Anti-Avoidance Rules	= 1 for no anti-avoidance rules, 0 if general/special rules	Schanz *et al*. (2017)
CFC Rules	= 1 for no CFC rules, 0 for CFC rules	Schanz *et al*. (2017)
Transfer Pricing Rules	= 1 for no transfer pricing rules, 0 for Transfer pricing rules	Schanz *et al*. (2017)
Corporate Tax Rate	Normalized to a range between 0 and 1, with a higher value indicating a more attractive statutory tax rate	Schanz *et al*. (2017)

they subtract each country's tax rate from this value and divide that by the maximum rate. Thus, the corporate tax rate is normalized to a range between zero and one, with a higher value indicating a more attractive statutory tax rate. We use this variable in our regression and expect its estimated coefficient to be negative.

Defining which countries are tax havens is a complicated challenge. Traditional methods for identifying tax havens are based on differences in the tax and legal structures for base erosion and profit-shifting (BEPS) practices.[9] Using this method, the EU includes 12 countries in its 2021 tax haven blacklist, mostly Caribbean and Channel Islands economies.[10] Likewise, the OECD defined a total of 35 locations as tax havens in 2000, but by 2017, only Trinidad & Tobago was still listed as a tax haven. Recently, Garcia-Bernardo *et al*. (2017) identified a larger set of tax havens using an analysis of big data on the ownership networks of 98 million global companies across countries. They identify a total of 55 tax haven countries, including some advanced countries that function as offshore financial centers (OFCs hereafter), such as the Netherlands, the U.K., Switzerland, Ireland and Singapore.[11]

[9]BEPS refers to the tax strategies used by multinationals to shift profits from higher-tax countries to lower-tax countries.

[10]They are American Samoa, Anguilla, Dominica, Fiji, Guam, Palau, Panama, Samoa, the Seychelles, Trinidad and Tobago, the US Virgin Islands and Vanuatu.

[11]According to Garcia-Bernardo *et al*. (2017), these five advanced countries channel about 47% of offshore

TABLE 2—COUNTRY LIST BY STATUS OF TAX TREATIES AND TIEA WITH U.S.

	In the Sample	Out of the Sample
Bilateral tax treaty with the U.S.	Australia, Austria, Belgium, Bulgaria, Canada, China, Cyprus*, the Czech Republic, Denmark, Egypt, Estonia, Finland, France, Germany, Greece, Hungary, India, Indonesia, Ireland*, Israel, Italy, Japan, Kazakhstan, Korea, Latvia, Lithuania, Luxembourg, Malta*, Mexico, Morocco, the Netherlands*, New Zealand, Norway, the Philippines, Poland, Portugal, Romania, Russia, Slovakia, Slovenia, South Africa, Spain, Sweden, Switzerland*, Thailand, Tunisia, Turkey, U.K.*, Ukraine, Venezuela	Bangladesh, Barbados*, Bermuda*, Iceland, Jamaica, Sri Lanka,. Pakistan, Trinidad and Tobago
TIEA with the U.S.	Brazil, Columbia, Costa Rica, the Dominican Republic, Mauritius, Mexico, the Netherlands*, Panama*, Peru	Antigua and Barbuda*, Aruba*, the Bahamas, Barbados*, Bermuda*, Curaçao, Dominica*, Gibraltar*, Grenada*, Guernsey*, Guyana, Honduras, the Isle of Man*, Jamaica, Jersey*, Liechtenstein*, the Marshall Islands*, Monaco*, Saint Lucia*, Saint Maarten*, Trinidad and Tobago, the British Virgin Islands*
No tax treaty nor TIEA with the U.S.	Angola, Argentina, Bahrain*, Botswana, Chile, Ecuador, Guatemala, Hong Kong*, Croatia, Kenya, Malaysia, Namibia, Nigeria, Nicaragua, Saudi Arabia, Singapore*, Uruguay, Zimbabwe	-

Note: * indicates a tax haven country as identified by Garcia-Bernardo *et al*. (2017).

We adopt the approach of Garcia-Bernardo *et al.* (2017) to define tax havens.

As our panel data are collected from multiple sources produced by various institutions, there is underlying variation in the data coverage across data sources. As a result, our final panel data contain a total of 78 countries over the period of 2007–2018. The country list is presented in Table 2.[12]

As of 2018, the U.S. had a total of 60 tax treaties and 32 TIEA in effect. As shown in Table 2, our dataset comprises the majority of the countries that have a tax treaty with the U.S. Meanwhile, many TIEA treaty signatories are excluded from our sample due to the unavailability of information on taxation-related regulations and rules. Another observation is that the lion's share of these countries consists of tax havens that have neither a tax treaty nor a TIEA with the U.S.

Tables 3 and 4 contain summary statistics and present the correlations among the variables, respectively. Tax treaties are positively correlated with FDI flows, as are TIEA but with a much lower correlation coefficient. The correlations of FDI flows with tax environment variables except for tax haven status are negative, meaning that the better and stronger the tax systems and rules, the larger the amount of FDI inflows. However, because these are simple correlation coefficients, if the regression

investments from tax havens.

[12]The tax treaty database (https://eoi-tax.com) contains information on only the most recent revisions of tax treaties. The original tax treaties for many countries, especially OECD countries, date from the 1930s to the 1970s. We revise the data on the dates that tax treaties went into effect based on Blonigen and Davies (2004).

TABLE 3—SUMMARY STATISTICS

Variable	Obs	Mean	Std. Dev.	Min	Max
FDI (logged value)	1,023	8.394	2.656	0	13.75
GDP (logged value)	1,120	25.711	1.744	21.835	30.242
Distance (logged value)	1,086	8.987	.475	7.64	9.709
TFP (logged value)	940	-.394	.301	-1.834	.372
Human Capital (logged value)	1,012	1.032	.207	.322	1.424
Tax Treaty	1,120	.56	.497	0	1
TIEA	1,120	.095	.293	0	1
TFA	1,086	.171	.377	0	1
WTO	1,086	.908	.289	0	1
Tax Havens	1,120	.182	.386	0	1
Anti-Avoidance Rules	1,112	.417	.388	0	1
CFC Rules	1,112	.661	.474	0	1
Transfer Pricing Rules	1,112	.248	.432	0	1
Corporate Tax Rate	1,112	.413	.219	0	1

analysis controls for other determinants of FDI, the relationship between these variables could change. Tax haven status is positively correlated with FDI, while physical distance appears to have less of an impact on investment than it does on international trade. Tax treaties are negatively correlated with TIEA and FTA.

TABLE 4—CORRELATION MATRIX

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)
(1) FDI	1.000														
(2) GDP	0.668	1.000													
(3) Distance	-0.115	0.195	1.000												
(4) TFP	0.380	0.133	-0.092	1.000											
(5) Human Capital	0.419	0.283	-0.049	0.363	1.000										
(6) Trade Cost	-0.126	-0.082	0.108	-0.229	-0.243	1.000									
(7) Tax Treaty	0.332	0.450	0.100	0.202	0.428	0.007	1.000								
(8) TIEA	0.121	-0.162	-0.491	-0.025	-0.084	-0.022	-0.255	1.000							
(9) FTA	0.056	-0.021	-0.382	-0.002	-0.111	-0.594	-0.204	0.329	1.000						
(10) WTO Member	0.043	0.283	0.243	0.097	0.034	-0.341	0.207	-0.221	0.139	1.000					
(11) Tax Haven	0.250	-0.284	-0.164	0.388	0.148	-0.065	-0.156	0.285	0.037	-0.284	1.000				
(12) Anti-avoidance	-0.363	-0.479	-0.115	-0.150	-0.398	0.201	-0.318	0.080	0.018	-0.279	0.084	1.000			
(13) CFC	-0.371	-0.534	0.011	-0.230	-0.346	0.102	-0.448	0.053	0.002	-0.154	0.297	0.331	1.000		
(14) Transfer Pricing	-0.221	-0.492	-0.079	-0.026	-0.361	0.113	-0.349	0.077	0.022	-0.237	0.317	0.349	0.406	1.000	
(15) Corporate Tax	-0.102	-0.390	-0.074	0.208	0.227	0.085	-0.153	0.050	-0.097	-0.363	0.386	0.292	0.257	0.288	1.000

IV. Empirical Results

A. *Main Results*

In this section, we report our estimation results from several different estimators. Table 5 contains the empirical results of the OLS, fixed effects, and random effects models. As presented in Column (1), a negative coefficient for tax treaties emerges when we run the OLS regression only using the gravity and absorptive capacity variables, which is often the case in existing studies (Blonigen and Davis, 2004; 2005).

TABLE 5—ESTIMATION RESULTS I: OLS, FE AND RE ESTIMATIONS

	(1) OLS	(2) OLS	(3) OLS	(4) Fixed Effects	(5) Random Effects
GDP	1.151***	1.152***	1.250***	0.534***	1.106***
	(0.040)	(0.039)	(0.046)	(0.138)	(0.071)
Distance	-0.320***	-0.040	-0.163*		-0.665**
	(0.092)	(0.108)	(0.090)		(0.280)
TFP	3.299***	3.120***	1.760***	-0.450	-0.510
	(0.293)	(0.281)	(0.232)	(0.402)	(0.346)
Human Capital	1.869***	1.551***	1.712***	1.520	1.523**
	(0.382)	(0.387)	(0.375)	(1.063)	(0.716)
Trade Cost		-11.514***	-3.786*	-1.793	-2.587
		(2.366)	(2.286)	(1.943)	(1.955)
Tax Treaty	-0.472***	-0.275**	-0.047	0.596***	0.566***
	(0.124)	(0.117)	(0.096)	(0.182)	(0.156)
TIEA		0.636***	-0.027	0.030	0.021
		(0.145)	(0.160)	(0.142)	(0.142)
FTA		-0.028	0.291**	-0.255	-0.259
		(0.168)	(0.148)	(0.173)	(0.164)
WTO Membership		-0.239	-0.429	-0.401	-0.363
		(0.326)	(0.401)	(0.253)	(0.253)
Tax Haven			2.261***		2.705***
			(0.154)		(0.337)
Anti-avoidance Rules			0.025	0.150	0.011
			(0.159)	(0.144)	(0.138)
CFC Rules			0.221**	0.007	-0.015
			(0.098)	(0.114)	(0.109)
Transfer Pricing Rules			0.169	0.288***	0.336***
			(0.142)	(0.085)	(0.086)
Corporate Tax Rate			-1.629***	-0.363	-0.584*
			(0.426)	(0.379)	(0.348)
Observations	851	851	851	851	851
R-squared	0.689	0.703	0.801		
(within)				0.171	0.151
(between)				0.559	0.794
(overall)				0.521	0.754
Hausman statistic				75.35***	

Note: 1) All regressions include year dummies. 2) Figures in parentheses are heteroscedasticity-robust standard errors. 3) *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

The size of the estimated coefficient becomes smaller if additional economic agreements are included in the estimation, but the negative sign remains, with high statistical significance (Column 2). In this regression, TIEA seemingly increases U.S. FDI into the partner countries. The effect of distance also becomes statistically insignificant. As we expect, the absorptive capacity variables all increase FDI.

On the other hand, as shown in Column (3), once we control for country-specific tax environments, the negative impact of tax treaties on FDI disappears. This implies that the omission of these variables in the estimation could lead to biased estimated coefficients and create some spurious inference regarding the impact of tax treaties on FDI. Despite the fact that the tax environments of recipient countries, such as their local tax systems and regulations, are inarguably decisive factors in investment decisions, previous studies have failed to consider country-specific tax environments as a determinant of FDI, leading to omitted variable bias in their estimations. Among the tax environment variables, tax haven status appears to be the most decisive factor affecting FDI flows. Other factors being equal, tax havens' FDI stock from the U.S. tends to be approximately 2.3% higher than that of non–tax havens. The regression results also indicate that the weaker a country's CFC rules, the higher their FDI inflow from the U.S. One puzzling finding in this regression is the negative coefficient of the corporate tax rate. As mentioned above, a higher value of this variable indicates a more attractive statutory tax rate. Therefore, we can expect a positive coefficient for this variable if a lower tax burden increases FDI inflows from the U.S. However, our estimate finds the exact opposite, with strong statistical significance. We suspect that the OLS regression does not sufficiently control for cross-country characteristics to produce a sensible marginal effect of the independent variables.

One way to control for country-specific heterogeneity in the estimation is to use a fixed effects or random effects estimator. Our panel data allow for the use of these estimators in order to control for time-invariant unobserved characteristics across countries. We report the estimation results using these estimators in Columns (4) and (5). As depicted in Table 5, the impact of tax treaties is statistically significant, with a positive sign for both the fixed effects and random effects estimations. Both models suggest that tax treaties increase FDI stock into a recipient country by about 0.6%. The estimated coefficients of the absorptive capacity variables are now insignificant. Similarly, the effect of FTA is statistically insignificant. Transfer pricing rules, instead of CFC rules, emerge as one of the key variables that determine the magnitude of FDI stock.

While our Hausman test suggests that the fixed effects model is the more appropriate model, the sizes of the estimated coefficients for tax treaties are similar across the two models. A pitfall of the fixed effects estimator, however, is that one cannot examine time-invariant causes of the dependent variables separately, as time-invariant predictors, such as tax haven status and physical distance, are perfectly collinear with the individual fixed effects.

As discussed above, Blonigen and Davis (2004) and Egger *et al.* (2006) found that recently signed tax treaties tend to decrease FDI flows. They claim that new tax treaties or the revision of old ones may reduce the incentive for FDI, as they contain a more sophisticated incentive scheme for limiting FDI for tax avoidance purposes. Given that our sample contains tax agreements concluded in more recent years, we

TABLE 6—CORRELATIONS BETWEEN TAX TREATIES AND OTHER VARIABLES

	Tax treaties before 1990s†	Tax treaties during1990s†	Tax treaties after 2000†	Tax treaties with OECD	Tax treaties with non-OECD
GDP	0.400*	0.341*	-0.085*	0.051	0.393*
Distance	0.030	0.142*	0.012	0.149*	0.251*
TFP	0.372*	0.012	-0.032	0.098	-0.094
Human Capital	0.289*	0.029	0.209*	0.340*	0.031
TIEA	-0.130*	-0.096*	-0.125*	-0.113	-0.198*-
FTA	-0.062	0.032	-0.179*	-0.335*	0.227*
WTO Membership	0.156*	0.086*	0.116*	-	0.120*
Tax Havens	-0.093*	-0.182*	0.026	0.110	-0.178*
Anti-Avoidance Rules	-0.247*	-0.058	-0.181*	0.028	-0.108*
CFC Rules	-0.365*	-0.333*	-0.014	-0.202*	-0.297*
Transfer Pricing Rules	-0.157*	-0.197*	-0.075	-0.016	-0.179*
Corporate tax Rate	-0.166*	-0.080*	0.036	0.069	-0.154*

Note: 1) † The period classification is based on the year when the first tax agreement came into force, 2) * indicates the significance at the 1% level after Bonferroni adjustment.

may find even more negative impacts if this claim is correct.

In Table 6, we report the correlation coefficients between tax treaties and other explanatory variables after classifying tax treaties into three groups based on the time of entry into force of the agreement. As shown in the table, for tax agreements that took effect before the 1990s, the target countries are mainly countries with a large GDP and high levels of productivity and human capital during our sample period. The strong negative correlation with tax environment variables suggests that countries with well-organized domestic regulations to prevent tax avoidance are more likely to have signed tax agreements with the United States.

On the other hand, tax treaties that came into force in the 1990s have a slightly lower correlation coefficient with GDP compared to the previous period, and the correlation with productivity and human capital level is not statistically significant. In addition, the correlation with tax haven status appears to have a negative relationship, indicating that there had been a strong tendency to enter into agreements with countries other than tax havens. For tax treaties that took effect after the 2000s, we find a negative correlation with GDP but a positive correlation with human capital. In addition, these treaties have no statistically significant relationships with tax haven status, CFC or transfer pricing rules. Hence, it appears likely that the United States had signed tax treaties with countries with large economies, high productivity and human capital levels, and well-organized tax regulations until the 1980s but that the target countries of tax treaties became more diversified after that decade.

Table 6 also includes correlation coefficients among these variables when the sample is divided into OECD countries and non-OECD countries. The results suggest that the size of GDP shows a positive correlation only in cases of tax treaties with the non-OECD countries. The negative correlation between tax treaties and TIEA in the non-OECD sample implies that the TIEA is used complementarily in countries that do not enter into tax treaties. We also find that the quality of domestic regulations to prevent tax avoidance matters for non-OECD countries. Countries with well-organized tax regulations are more likely to have signed tax treaties with the United States.

Table 7 contains the analysis results estimated by comparing the impacts of recently signed tax treaties with those of older ones. In Columns (1) and (2), we

TABLE 7—ESTIMATION RESULTS II: OLD VERSUS NEW TAX TREATIES

	(1) Fixed Effects	(2) Random Effects	(3) Fixed Effects	(4) Random Effects
GDP	0.498***	1.057***	0.498***	1.118***
	(0.136)	(0.073)	(0.136)	(0.073)
Distance	-	-0.669**	-	-0.719**
		(0.280)		(0.284)
TFP	-0.455	-0.600*	-0.455	-0.504
	(0.397)	(0.347)	(0.397)	(0.346)
Human Capital	1.575	1.156	1.575	1.657**
	(1.050)	(0.729)	(1.050)	(0.713)
Trade Cost	-1.904	-2.598	-1.904	-2.652
	(1.921)	(1.943)	(1.921)	(1.952)
Tax Treaties (~1980s)	-	1.127***		
		(0.303)		
Tax Treaties (1990s~)	1.260***	0.818***		
	(0.235)	(0.193)		
Tax Treaties (~1990s)				0.711***
				(0.266)
Tax Treaties (2000s~)			1.260***	0.703***
			(0.235)	(0.207)
TIEA	0.028	0.031	0.028	0.019
	(0.140)	(0.141)	(0.140)	(0.142)
FTA	-0.247	-0.240	-0.247	-0.287*
	(0.171)	(0.163)	(0.171)	(0.164)
WTO Membership	-0.408	-0.374	-0.408	-0.355
	(0.250)	(0.252)	(0.250)	(0.253)
Tax Haven	-	2.773***	-	2.744***
		(0.339)		(0.342)
Anti-avoidance Rules	0.127	0.002	0.127	-0.004
	(0.142)	(0.138)	(0.142)	(0.138)
CFC Rules	-0.000	-0.000	-0.000	-0.006
	(0.112)	(0.109)	(0.112)	(0.110)
Transfer Pricing Rules	0.274***	0.327***	0.274***	0.342***
	(0.084)	(0.085)	(0.084)	(0.085)
Corporate Tax Rate	-0.296	-0.573*	-0.296	-0.612*
	(0.374)	(0.346)	(0.374)	(0.349)
Observations	851	851	851	851
R-squared				
(within)	0.190	0.168	0.190	0.162
(between)	0.362	0.779	0.421	0.784
(overall)	0.344	0.738	0.413	0.746
Hausman statistic	53.39***		60.49***	

Note: 1) All regressions include year dummies. 2) Figures in parentheses are heteroscedasticity-robust standard errors. 3) *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

compare the effects of tax treaties that took effect before the 1980s and those that took effect thereafter. In the fixed effects model, for the tax agreements that took effect before the 1980s, the coefficient cannot be estimated because the agreements are perfectly collinear with the fixed effects. On the other hand, according to the random effects estimation results, both new and old tax treaties appear to have a positive effect on FDI in our sample. However, the size of the effect is smaller for the new treaties compared to older ones. Columns (3) and (4) show the results of a re-estimation when the tax agreements are divided into those that went into effect before the 1990s and those that went into effect afterward. These results are qualitatively similar to those shown in columns (1) and (2). Therefore, our results suggest that tax treaties have a positive effect on direct investment regardless of the time they went into effect, but the investment promotion effect is somewhat reduced in the case of newer agreements.

Table 8 includes the estimates of the effects of tax treaties when the sample is divided into OECD countries and non-OECD countries. Columns (1) and (2) include the results estimated for the entire sample after creating interaction terms between the tax treaty dummy and the OECD country dummy. In this case, the estimated coefficient represents the effect of tax agreements in other groups compared to non-OECD countries without tax treaties.[13] According to the random effect analysis, the group of non-OECD countries with tax treaties has, on average, 0.96% more FDI stock invested from the U.S. than the group of non-OECD countries that do not have tax treaties with the U.S. While OECD countries have positive estimated coefficients regardless of whether they have a tax treaty, it should be noted that the size of the estimated coefficient for the OECD group with tax treaties is smaller than that of the OECD group without tax treaties. This implies that among OECD countries, tax agreements may not have a positive effect on FDI inflows from the U.S.

To verify this, after dividing the entire sample into OECD countries and non-OECD countries, we separately estimate each subsample.[14] These results are presented in Columns (3) through (6). As shown in Table 8, we find that tax treaties appear to increase FDI among non-OECD countries. The estimated coefficient of tax treaties for the non-OECD sample is statistically significant, and tax treaties appear to increase the FDI stock invested from the U.S. by about 1%. On the other hand, there is no statistically significant effect of tax treaties on FDI in the OECD sample. Our results can be interpreted as evidence, as Baker (2014) suggests, that developed countries have various institutional mechanisms to prevent double taxation other than tax treaties, implying that the net effect of tax agreements may not appear. Brooks and Krever (2015) also argue that tax treaties could be redundant in developed countries, taking into account that most advanced economies have domestic tax laws stipulating either an exemption for tax income derived from other countries or a tax credit for taxes paid in the source country.

[13] All OECD countries except Chile, Columbia, Costa Rica and Croatia have tax treaties in effect with the United States. In addition, these countries entered into tax treaties before the mid-2000s, when our analysis begins. Consequently, in the fixed effects model, the effects of tax treaties on these countries are included in the fixed effects; thus, separate coefficients cannot be estimated.

[14] When we run the regressions for the OECD and non-OECD subsamples separately, we find that the standard Hausman test cannot be used, as its asymptotic assumptions are not met. An alternative is to adopt the correlated random effects approach proposed by Mundlak (1978). We report the F-test statistics based on this approach in Table 7 and confirm that the fixed effects estimator is more appropriate.

TABLE 8—ESTIMATION RESULTS (III): OECD VERSUS THE NON-OECD SAMPLE

	Whole Sample		OECD Sample		Non-OECD Sample	
	(1) Fixed Effects	(2) Random Effects	(3) Fixed Effects	(4) Random Effects	(5) Fixed Effects	(6) Random Effects
GDP	0.485***	1.057***	1.230***	1.221***	0.307	1.019***
	(0.137)	(0.073)	(0.241)	(0.098)	(0.205)	(0.101)
Distance	-	-0.616**	-	-0.761*	-	-0.502
		(0.294)		(0.402)		(0.363)
TFP	-0.472	-0.587*	0.073	0.715	-0.931*	-0.851*
	(0.397)	(0.348)	(0.563)	(0.490)	(0.558)	(0.488)
Human Capital	1.549	1.234	2.331	1.728	1.028	1.287
	(1.050)	(0.784)	(2.464)	(1.284)	(1.328)	(1.004)
Trade Cost	-1.879	-2.579	3.758	2.472	-1.922	-3.990
	(1.921)	(1.941)	(3.326)	(3.396)	(2.548)	(2.546)
OECD (Tax Treaty=0)	0.282	1.206***				
	(0.271)	(0.397)				
Non-OECD (Tax Treaty=1)	1.254***	0.958***				
	(0.235)	(0.203)				
OECD (Tax Treaty=1)		0.991***	-0.112	-0.195	1.179***	0.891***
		(0.310)	(0.224)	(0.221)	(0.269)	(0.221)
TIEA	0.026	0.035	0.334*	0.253	0.043	0.132
	(0.140)	(0.141)	(0.190)	(0.195)	(0.205)	(0.204)
FTA	-0.246	-0.261	-0.166	0.037	-0.429	-0.554**
	(0.171)	(0.163)	(0.222)	(0.214)	(0.285)	(0.265)
WTO Membership	-0.413*	-0.367	-	-21.046***	-0.461	-0.317
	(0.250)	(0.251)		(5.785)	(0.290)	(0.290)
Tax Haven	-	2.799***	-	3.226***	-	2.473***
		(0.343)		(0.427)		(0.463)
Anti-avoidance Rules	0.127	0.011	0.260	0.251	0.017	-0.086
	(0.142)	(0.138)	(0.176)	(0.174)	(0.218)	(0.208)
CFC Rules	0.002	-0.009	-0.030	0.052	-0.064	-0.127
	(0.112)	(0.109)	(0.138)	(0.133)	(0.175)	(0.169)
Transfer Pricing Rules	0.275***	0.333***	-0.205	-0.129	0.330***	0.349***
	(0.084)	(0.085)	(0.153)	(0.157)	(0.110)	(0.110)
Corporate Tax Rate	-0.294	-0.571*	0.274	-0.244	-1.172*	-0.823
	(0.374)	(0.347)	(0.410)	(0.402)	(0.673)	(0.562)
Observations	851	851	400	400	451	451
R-squared						
(within)	0.191	0.171	0.259	0.247	0.236	0.202
(between)	0.400	0.785	0.559	0.812	0.260	0.718
(overall)	0.350	0.745	0.487	0.773	0.237	0.658
Hausman statistic	48.19***		11.79***		31.62***	

Note: 1) All regressions include year dummies. 2) Figures in parentheses are heteroscedasticity-robust standard errors. 3) *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively, 4) F-test statistic based on the correlated random effects approach.

B. *Robustness Check*

We perform robustness tests of the regression results in the following way.[15] First, it is possible that our analytic results may be affected by endogeneity problems, especially reverse causality bias. While our testing hypothesis is the causal impact of tax treaties on FDI, it may also be likely that tax treaties tend to be signed with countries with large FDI flows. To mitigate this potential bias, we employ Arellano and Bond (1991)'s GMM estimation method. In the meantime, we should recall that our tax treaty variable is a dummy variable and thus the first-differencing of this variable implies that the tax treaty effect can be estimated only for the exact year following the year the tax treaty entered into force. Following Barthel *et al*. (2010), we overcome this problem by applying an alternative tax treaty variable that measures the total number of years that pass after a tax treaty enters into effect. In this case, the first-differenced variable has a value of one for all years after a tax treaty becomes effective and zero otherwise, which is exactly what we want to test for.

Our GMM results are presented in Table A1 in the Appendix. We confirm that the estimated effect of tax treaties on FDI remains statistically significant and positive, although its statistical significance becomes weaker than in the fixed effect estimation. The Hansen statistics for over-identifying restrictions indicate that the validity of our instruments is not rejected.

Second, we use the approach of Garcia-Bernardo *et al*. (2017) to define tax havens, which includes several OECD countries. Given that some OECD countries, notably the U.K. and the Netherlands, have more FDI stock invested from the U.S. than others, it is possible that they are outliers in the analysis, affecting the estimation results. Thus, we re-run the regression excluding the Netherlands, the U.K., Switzerland, Ireland and Singapore from the list of tax havens. We confirm that such a change does not greatly affect our results quantitatively or qualitatively.

Third, considering the possibility that the tax prevention regulations have been strengthened in recently revised tax treaties, we conduct a regression analysis based on the time of entry into force of the most recently revised tax treaties. We find that despite these changes, the robustness of our analytic results is preserved. This result is consistent with findings by Davies (2003) showing that treaty renegotiations have no robust impact on FDI.

Fourth, we add more tax-related variables, such as country-specific tax withholding rates for dividends, interest, and royalties, as explanatory variables in the estimation. The estimated coefficients of these variables are largely insignificant. At the same time, our results suggest that a lower tax withholding rate levied on dividends attracts more FDI among the non-OECD sample, while a lower tax withholding rate levied on royalties increases FDI in the OECD sample.

Finally, Blonigen and Davies (2004) employ a slightly different specification compared to ours. In particular, basing on the empirical specification of Carr, Markusen and Maskus (2001), they include the following GDP-related variables in their estimations to distinguish between horizontal and vertical motivations for FDI: the sum of the two countries' GDPs and the squared difference between the GDPs.

[15]These results are available upon request.

Again we confirm that our results remain robust when these variables are taken into account.

V. Conclusion

This paper empirically examines the relationship between tax treaties and foreign direct investment using U.S. outbound FDI to 78 countries over the period of 2007–2018. Our results suggest the importance of controlling for country-specific tax environments in the estimation. Once these, along with other unobserved country-specific characteristics, are controlled, we find a positive impact of tax treaties among the non-OECD sample but no statistically significant impact of tax treaties among the OECD sample. Our results indicate that recently signed tax treaties increase FDI but with a smaller impact than the older treaties.

As discussed above, the mixed empirical evidence pertaining to the effect of tax treaties on FDI has contributed to controversy over the validity of such treaties. For instance, Kysar (2020) suggests that the United States should cancel or scale down its tax treaties, given the lack of evidence of their overall positive effect. Brooks and Krever (2015) claim that tax treaties could be a 'poisoned chalice' for developing countries, encouraging such countries to give up their tax rights without receiving sufficient benefits, such as increased FDI. Thuronyi (1999) even propose the establishment of a World Tax Organization to create a fairer global tax system. Taking into consideration the accelerated pace of globalization and digitalization, reform of the existing architecture of bilateral tax treaties may inevitably be needed. However, prior to any institutional reform, more extensive research on bilateral tax treaties is needed.

Based on the empirical results in this paper, we suggest the following agenda for future research. First, although this paper confirms the benign effect of tax treaties on FDI flows, it does not guarantee that the benefits are sufficiently large to outweigh the costs incurred from forfeiting taxation rights. Hence, a more detailed cost-benefit analysis is imperative. Second, estimations using either a wider set of data or more micro-level data would definitely be helpful. Third, taking into consideration that many countries are parties to multiple tax treaties, further analysis of the tax treaty network across countries is needed. Finally, it is desirable explicitly to consider differential attributes of tax treaties in the estimation as opposed to using a simple binary treatment. This is particularly the case because we observe that there is not much year-to-year variation in tax treaty status as a binary treatment. While there exists a growing number of studies dealing with treaty attributes, most of them focus only on dividend withholding tax rates. A more comprehensive analysis of treaty attributes and their linkage to FDI would provide a better understanding of the tax treaty-FDI linkage.

APPENDIX

TABLE A1—ESTIMATION RESULTS: ARELLANO-BOND GMM ESTIMATION

	(1) Whole Sample	(2) OECD Sample	(3) Non-OECD Sample
FDI (t-1 lagged)	0.636***	0.593***	0.562***
	(0.069)	(0.128)	(0.086)
GDP	0.066	0.770*	0.260
	(0.443)	(0.400)	(0.472)
TFP	-0.526	-1.750	0.906
	(0.659)	(2.197)	(0.932)
Human Capital	3.001	1.575	4.006
	(2.480)	(2.702)	(2.430)
Trade Cost	3.495	-0.001	5.370
	(3.770)	(1.507)	(4.819)
Tax Treaties	0.054*	0.002	0.035*
	(0.032)	(0.007)	(0.020)
TIEA	0.042*	-0.012	0.056
	(0.022)	(0.026)	(0.038)
FTA	0.479	-0.027	-0.098
	(0.320)	(0.113)	(0.467)
WTO Membership	-0.103		0.479
	(0.607)		(0.875)
Anti-avoidance Rules	0.021	-0.314	-0.118
	(0.235)	(0.299)	(0.389)
CFC Rules	0.209	0.001	0.202
	(0.223)	(0.123)	(0.339)
Transfer Pricing Rules	0.067	-0.215	0.332*
	(0.144)	(0.194)	(0.188)
Corporate Tax Rate	-0.721	0.614*	-2.204
	(0.973)	(0.344)	(1.672)
Observations	674	319	355
Number of id	75	35	40
Arellano-Bond test for AR(2)	-1.43	0.36	-14.1
	(0.152)	(0.721)	(0.158)
Hansen Test Statistic	51.45	16.24	19.36

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Effects of Fiscal Instability on Financial Instability†

By SUNJOO HWANG*

This paper empirically examines how fiscal instability affects financial instability. According to an IMF forecast (2021a), the fiscal space in Korea will be steadily reduced in the future. The theoretical literature predicts that if fiscal stability is undermined, financial stability will also be in danger given that government guarantees on banks are weakened and/or sovereign bonds held in banks become riskier. This paper empirically finds the existence of this negative impact of fiscal instability on financial instability. I also find that the intensity of this fiscal-financial relationship is greater in a country where (i) its currency is not a reserve currency such as the US dollar or euro, (ii) its banking sector is large relative to government sector, and/or (iii) its private credit to GDP is high. Korea has all of these three characteristics and hence needs to put more effort into maintaining fiscal stability.

Key Word: Fiscal Instability, Financial Instability, Sovereign Bond, Implicit Government Guarantees, Noncore Currency
JEL Code: G01, G21, H60

I. Introduction

Since the 2020 Covid-19 recession, the fiscal space in Korea has been significantly reduced. According to the long-term forecast by the IMF (2021a), government debt and the budget deficit will continue to increase. This is in stark contrast to the forecasts for other advanced countries without reserve currencies in which fiscal stability will be steadily improved over the same period.

Although fiscal instability itself has attracted a considerable amount of attention, its effects on financial instability do not attract much attention despite its importance. The 2010-2014 European Sovereign Debt Crisis is an important example that shows that fiscal instability leads to financial instability. During that time, the government debt in many European countries expanded and the risk of sovereign default spiked.

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Therefore, sovereign bond prices fell, which triggered a deterioration of bank balance sheets as European banks held significant amounts of sovereign bonds. Also, the implicit government guarantees on banks became unreliable and, therefore, banks' funding conditions were also severely damaged given that capital market investors usually lend money to banks believing in the implicit government guarantees. In the end, the financial system melted down. According to Laeven and Valencia (2020), if a financial crisis occurs, it takes more than seven years for the economic growth path to recover its original trend. Therefore, Obstfeld (2013) asserts that (i) financial stability should be an objective of fiscal policy, (ii) the fiscal space should be large enough to deal with financial crises, and (iii) government debt should be reduced.

The objective of this paper is empirically to examine whether fiscal instability leads to financial instability in Korea and other advanced countries. The literature suggests two channels of the fiscal-financial relationship. The first channel of implicit government guarantees can be understood in the following way (Leonello, 2018). Banks are highly likely to obtain a government bailout when they fail because they are systemically important. That is, there exists implicit government guarantees on banks. If the fiscal space of a government is weakened for some reason, the creditability of the implicit guarantees will be undermined simply because the government does not have sufficient reserves. Therefore, capital market investors withdraw their trust in banks and hence the bank default risk rises. The second channel is related to sovereign debt. Banks are in most countries the largest investors in sovereign bonds. For instance, Korean banks hold 40% of outstanding Korean treasury bonds as of 2020. The share of sovereign bonds on bank balance sheets is also usually high. In this regard, if the sovereign default risk rises and hence the sovereign bond price falls, the asset quality of banks will be decreased and financial turmoil could therefore arise (Acharya *et al.*, 2014).

Another aim of this paper is empirically to explore the determinants of the effects of fiscal instability on financial instability in the context of Korea. The Korean economy can be distinguished from other advanced economies in the following three ways. First, the Korean currency is not a reserve currency or core currency such as the US dollar or euro[1]. Second, the banking sector is huge relative to the government sector. Third, bank loans made to households and the corporate sector relative to GDP are also very large. All three of these characteristics potentially affect the fiscal-financial relationship. First, countries with a noncore currency may be in a disadvantageous position compared to countries with a core currency in overcoming financial crises. Second, if the banking sector is large but government revenue amounts are small, the banking sector is too large to be saved by the government. Third, if bank loans given to households and nonfinancial corporations are huge and hence the debt of these economic agents is overridden, it becomes more difficult for the government to rescue the banking sector when it faces a crisis.

This paper uses the CDS premium measure as a measure of default risk or the degree of instability because it reflects the forward-looking information of market participants pertaining to the likelihood of a default. The dataset contains data on all advanced countries, including Korea, and many deposit-taking commercial banks

[1]Throughout this paper, I use the two terms core currency and reserve currency interchangeably.

headquartered in these countries for the sample period of 2003-2021.

The first main result of this paper is that an increase in the sovereign CDS premium is associated with a rise in the bank CDS premium, consistent with the two aforementioned theoretic channels. The magnitude of this influence of the sovereign default risk on the bank default risk is large and is pronounced in countries with relatively higher levels of default risk. The second main result is that this magnitude is greater if the currency of the country of interest is not a core currency, if total assets in the banking sector relative to government revenue are relatively large, or the ratio of private credit to GDP is relatively high.

These empirical findings provide important implications about the Korean economy. If fiscal stability is undermined in Korea, the financial system may experience crises more likely at a greater magnitude than in other advanced countries. To resolve this adverse link between fiscal instability and financial instability, policies should reduce banks' reliance on the government and improve their stand-alone prudence and competitiveness.

This paper is related to the literature on fiscal-financial relationships. Because the European Sovereign Debt Crisis is a historic event that highlighted the importance of the fiscal-financial link, the literature focuses on European countries. Acharya *et al.* (2014) consider Eurozone countries, north European countries, Switzerland, and the UK, finding a two-way feedback loop between fiscal instability and financial instability. De Bruyckere *et al.* (2013) also empirically examine only European countries. Demirgüç-Kunt and Huizinga (2013) consider both European and non-European countries, including Morocco, Romania, Mexico, Turkey, Hungary, Poland and other developing countries. Therefore, the outcome of their research may be less relevant to Korea. An additional difference between earlier findings and this paper is that several important characteristics of the Korean economy, such as the absence of a core currency, a large banking sector relative to the government sector, and very high private credit are addressed only in this paper.

Among the relevant papers, Acharya *et al.* (2014) is the most closely related to the current paper. They use daily CDS data to analyze very short-term *two-way* interactions between fiscal and financial instability in stressed times such as the Global Financial Crisis. Interestingly, they do not find any significant relationship between fiscal and financial instability in *normal times*. In contrast, the objective of this paper is to determine the *one-way* influences of fiscal instability on financial instability in *normal times*, as fiscal instability is currently not severe in Korea, but in the longer run obviously the fiscal space will be greatly reduced. Therefore, this paper seeks to determine whether the effect of fiscal instability on financial instability exists in normal periods.

II. Background

A. *Channels through which fiscal instability affects financial instability*

1. Implicit guarantee channel

It is well known that the failure of large-sized banks can trigger an economic crisis.

Banks offer a number of socially vital financial services, such as payments, money transfers, deposits, and loans. If a bank fails, these crucial services may no longer be properly supplied, causing significant trouble in the overall economy. Also, banks are related to nearly all economic agents, including households, corporations, and nonbank financial companies through deposits, loans, and other financial products. Therefore, bank failures can trigger chain-reaction failures of these and many other economic agents.

Due to this systemic importance, governments usually bail out troubled banks. Whenever there were system-wide financial crises, such as the 1997 Asian Currency Crisis, the 2008 Global Financial Crisis, and the 2010 European Sovereign Debt Crisis, governments rescued failed banks. That is, it is obvious that governments provide banks certain implicit guarantees. Although these guarantees are implicit and often not formalized in national laws in many countries, one cannot deny the fact that they exist.

The source of these government bailouts is taxpayer's money. Only if there is enough fiscal space can governments save banks (Komárek and Komárková, 2015). In previous financial crises, it was rare to find only one or two banks failing while other many banks remained sound. Because banks are interconnected and their business models are almost identical, if one bank fails for some reason, it is highly likely that other many banks will fail for a similar reason. Therefore, governments need a large enough fiscal space to save all of such banks at the same time during crisis periods.

Thus, fiscal instability could cause financial instability. If the fiscal space is not large enough to save all banks at the same time, many banks cannot survive crisis periods by themselves and hence could fail and be forced into a disorderly liquidation (Leonello, 2018; Caruana and Avdjiev, 2012). If capital market participants realize that governments are unable to support banks, banks' credit ratings will be downgraded and their cost of funding will thus increase. In some stressed situations, banks find it impossible to refinance, greatly increasing the likelihood of a bank default (Bobetko *et al.*, 2013; Das *et al.*, 2010).

2. Sovereign debt channel

If fiscal stability is undermined, a sovereign entity's ability to repay its debt will be questionable, increasing the risk of a sovereign default and in turn causing a decrease in sovereign bond prices. In many counties, banks are primary investors in sovereign bonds. For instance, Korean banks held 40% of Korean treasury bonds as of 2020. Similarly, in other countries, domestic banks are predominant investors in domestic sovereign bonds.

Because sovereign bonds are one of the major asset classes on bank balance sheets, if the sovereign bond price falls due to a surge in sovereign credit risk, the asset quality of most banks will be severely undermined, which could trigger financial turmoil (Acharya *et al.*, 2014).

B. *Current status in Korea*

Below, I shall observe the current status of fiscal stability and financial stability in Korea. Although fiscal stability and financial stability could depend on a number of different factors, I consider the debt levels of the government and the private sector as indebtedness is easy to understand and frequently considered as important in the literature.

Until recently, South Korea has maintained a low sovereign debt level relative to other advanced countries. The ratio of government debt to GDP was only 40% in 2019. However, the fiscal space in Korea has been reduced quickly due to the Covid-19 pandemic. Even worse, the debt-to-GDP ratio of Korea is forecasted to increase continuously in the medium to longer term, whereas other advanced countries are forecasted to deleverage their sovereign debt. As a benchmark country group, I shall consider the group of advanced countries (i.e., *noncore currency country group*) that do not have core currencies such as the US dollar, euro, UK pound sterling, Japanese yen, Swiss franc, and Canadian dollar.[2] These advanced countries with noncore currencies are South Korea, Australia, the Czech Republic, Denmark, Israel, New Zealand, Norway, and Sweden.

According to forecasting by the IMF (2021a), the ratio of government debt to GDP in Korea will increase steadily to 70% by 2026 (see Figure 1). In contrast, the average debt-to-GDP ratio of other advanced countries with noncore currencies will increase only to 55% by 2023 and then will decrease thereafter.

Other forecast indicators also show that the fiscal stability of Korea will weaken over time. For instance, *net* sovereign debt, which is *gross* sovereign debt minus certain financial assets, will also increase to 40% by 2026 in Korea, while in other advanced countries with noncore currencies it will be only 14% by 2026 (IMF, 2021a). Similarly, the government budget deficit of Korea will exceed 2% of GDP by 2026, whereas for the group of advanced countries with noncore currencies, it will converge to 0% of GDP by 2026.

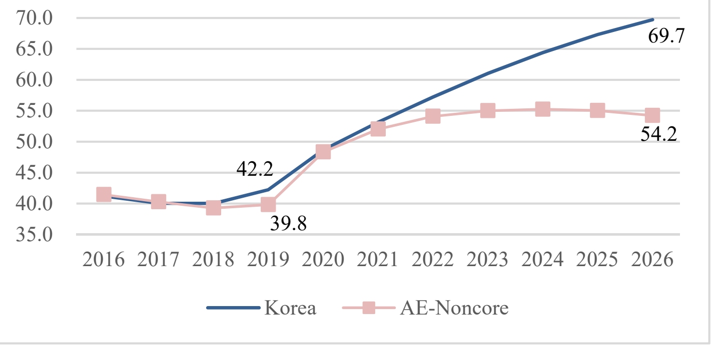

FIGURE 1. FORECAST OF THE RATIO OF GOVERNMENT DEBT TO GDP

Note: 1) The numbers for 2020 and thereafter are forecasts by the IMF (2021a), 2) AE-Noncore is the average of Australia, the Czech Republic, Denmark, Israel, New Zealand, Norway, and Sweden.

Source: IMF (2021a).

[2]Although there may be disagreement on which currency is a core currency (or reserve currency), I adopt the opinion of the Bank of Korea that US Dollar, Euro, UK Pound Sterling, Japanese Yen, Swiss Franc, and Canada Dollar are core currencies (Bank of Korea, 2018). According to the Bank of Korea, the central banks of the United States, the United Kingdom, Japan, Switzerland, and Canada, and the European Central Bank form a permanent currency swap network and, hence, their currencies can be regarded as core currencies.

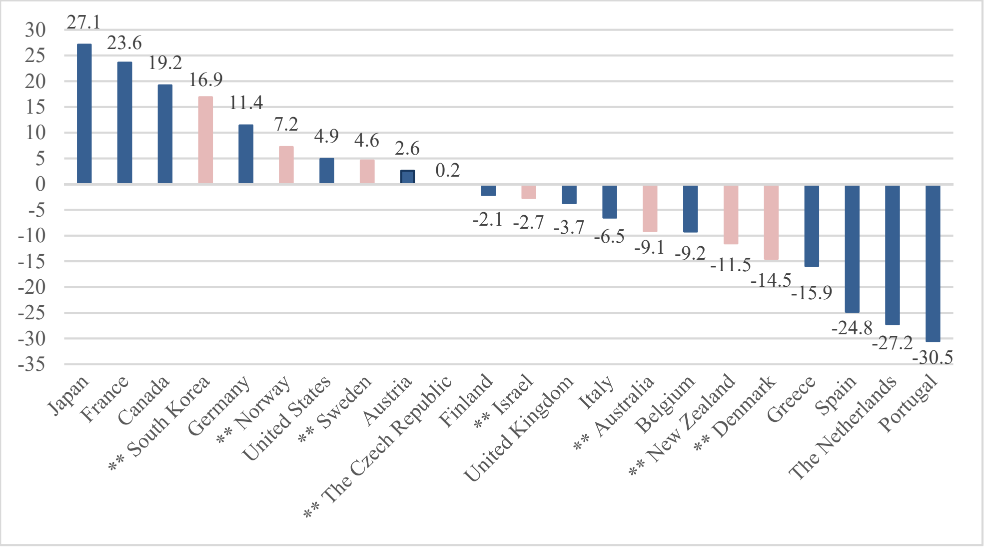

FIGURE 2. PRIVATE CREDIT TO GDP GAP (2020. 3Q)

Note: The countries with double asterisks belong to the noncore currency group.

Source: BIS (2021).

The ratio of private credit to GDP measures the volume of bank loans made to households and nonfinancial companies. The empirical literature finds that higher ratios of private credit to GDP are associated with a greater likelihood of a credit crisis (Hahm *et al.*, 2013). In 2020, private credit to GDP in Korea is 212%, which is 25%p higher than the average of all advanced countries. Recently, this ratio increased rapidly in Korea (see Figure 2).

The difference between the *level* and the *trend* of the ratio of private credit to GDP is called the private credit to GDP *gap*. This credit gap measures how quickly a country's private credit increases relative to its long-term trend. According to the Bank of International Settlements, the likelihood of a credit crisis is alarming if this credit gap exceeds 10%p (BCBS, 2010). As of 2020, the credit gap in Korea was 17%p, the highest among all advanced countries with noncore currencies and 24%p higher than the average of all advanced countries.

In sum, the indebtedness of the government and of households and nonfinancial companies in Korea is not just high but also increasing rapidly. Therefore, maintaining fiscal stability and financial stability should be important policy objectives.

C. *Testable hypotheses*

The objective of this paper is to examine whether fiscal instability leads to financial instability. Fiscal instability depends on many factors, such as the credit risk of sovereign debt, tax revenues, and government spending. Therefore, fiscal instability can be defined in a number of different ways. To be more specific, this paper focuses on the default risk of sovereign debt. Similarly, with regard to financial instability, I confine my attention to the default risk of bank debt. Thus, this paper will empirically examine the following hypothesis:

Hypothesis 1. *If the default risk of sovereign bond increases, the default risk of bank bond also increases.*

If this hypothesis is correct, it will be useful to ascertain what determines the intensity of this fiscal-financial relationship. To this end, this paper considers three factors: noncore currencies, banking sector total assets to government revenue, and the private credit to GDP. In the following sections, I discuss how these factors potentially affect the magnitude of the influence of fiscal instability on financial instability. These three factors are also relevant to the Korean economy given that Korea is a country with a noncore currency and that its ratio of private credit to GDP is large and increasing rapidly. In addition, Korea is special among advanced countries in that the banking sector is very large relative to the government sector. The ratio of total assets in the banking sector over government revenue was 620% in 2017, the highest among all advanced countries (World Bank, 2017; IMF, 2021b) and substantially higher than the average (260%) in advanced countries. During the 2008 Global Financial Crisis, Iceland was unique among European countries because it decided not to bail out certain bank creditors. In that country, the banking sector was notoriously massive relative to the government sector. Even in Iceland, the ratio of total assets in the banking sector to government revenues was only 415% in 2008, meaning that the banking sector in Korea is too large for the government to bail out.

First, the intensity of this fiscal-financial relationship may depend on whether the currency of a country of interest is a core currency. If the sovereign default risk increases for a country, it is more difficult for the government to procure funds to save banks when the banks are distressed. As an alternative to sovereign debt issuance, the central bank of the country may consider issuing more money to save distressed banks, though doing so could increase the inflation rate.

The magnitude of the inflation risk may be lower if the country has a core currency because a core currency is used not just in the given country but globally. For this reason, the burden of the decreased value of money will be shared not only by citizens of the country but also by all nations around the world. In contrast, if the country does not have a core currency, only the citizens of the country must shoulder the burden of inflation. In this regard, one can expect that the value of a core currency will not decrease much even if the central bank responsible for that currency increases the rate of money issuance, whereas the value of a noncore currency will decrease by a greater magnitude when the central bank of such a currency increases money issuance by the same amount.

Therefore, if the sovereign default risk is high and banks fail, countries with a core currency can easily find funds to save distressed banks even if they cannot rely on sovereign bond issuance, whereas countries with a noncore currency are in a more difficult position with regard to saving their banks. In this sense, this paper considers the following hypothesis.

Hypothesis 2. *If the default risk of sovereign bonds increases, the default risk of bank bonds increases at a greater magnitude in countries with a noncore currency as opposed to countries with a core currency.*

Secondly, the intensity of the fiscal-financial relationship may also depend on the size of the banking sector relative to government revenues. If the sovereign default risk rises and hence government borrowing becomes more difficult, governments may consider increasing tax rates and taxable income levels. Such an effort to increase tax revenues may be futile with regard to overcoming financial crises if the tax revenue is not enough to rescue distressed banks. In sum, if the risk of sovereign default increases and hence alternative financing arrangements such as increases in tax revenues are required when banks fail, the government's capacity to save the banking sector is low in countries where the banking sector is too large to save. In this regard, this paper considers the following hypothesis.

Hypothesis 3. *If the default risk of sovereign bonds increases, the default risk of bank bonds increases at a greater magnitude in countries where the size of the banking sector relative to government revenues is relatively large.*

Last but not least, the level of the private credit to GDP may also determine the intensity of the fiscal-financial relationship. The greater the ratio of private credit to GDP, the more banks are interconnected with the private nonfinancial sector via loans. The more bank loans there are, the more banks are exposed to outside shocks. If the sovereign default risk rises, banks are more likely to be required to overcome outside shocks by themselves without relying on government support. If the size of bank loans is relatively large, it will be more difficult for banks to deal with outside shocks by themselves. In this regard, this paper considers the following hypothesis.

Hypothesis 4. *If the default risk of sovereign bonds increases, the default risk of bank bonds increases at a greater magnitude at higher ratios of private credit to GDP.*

III. Empirical Analysis

A. *Influences of Sovereign Default Risk on Bank Default Risk*

I utilize a sovereign-bank panel dataset. I construct this dataset by combining individual datasets from Eikon, the World Bank, the BIS, Bloomberg, and Moody's. Sovereign and bank default measures are obtained from Eikon. Fiscal and financial variables are obtained from the World Bank. Information on private credit is obtained from the BIS. Bank-level financial ratios are obtained from Bloomberg, and the credit rating information is sourced from Moody's investor service.

The panel dataset covers the period from 2003 1Q to 2021 1Q and contains quarterly variables. The dataset contains information on all advanced countries except Hong Kong, Singapore, Luxembourg, and Iceland, as they are very small countries in terms of territory and population. Consequently, there are 29 advanced countries in the dataset used here.

The objective of this paper is to explore the effects of the risk of sovereign default on the risk of bank default, but not vice versa. However, bank default risk can in reverse also affect sovereign default risk. This is particularly the case during periods of financial crises (BIS, 2016). If financial crises occur, banks are highly likely to become insolvent and hence governments typically provide massive bailout packages, causing a meltdown of fiscal stability.

However, financial instability does not have material impacts on fiscal instability during normal times. In normal times, the financial system does not face serious stability issues. Of course, it is not impossible for one or a few banks to be troubled in normal times with temporary liquidity problems. However, central banks can easily resolve such problems as the lender of last resort by providing emergency liquidity. Even if central banks do not intervene, banks are usually able to overcome these idiosyncratic problems given that the overall capital market works well, allowing them to borrow money from the capital market.

The BIS (2016) also observes that financial stability affects fiscal stability only in financial crisis periods. By looking at the movement of government debt during the period of 1970 to 2015, it can be observed that government debt was generally stable before and after financial crises, whereas debt increased dramatically only during crisis periods.

In this sense, I confine my attention to normal times. There are two major financial crises during the period of 2003 to 2021. During the 2007-2009 Global Financial Crisis, all advanced countries suffered to some extent. In contrast, only European countries suffered from the 2010-2014 European Sovereign Debt Crisis. Therefore, in the following empirical analysis, I rule out every country from the analysis for the period of 2007 to 2009. For the period of 2010 to 2014, I rule out only European countries from the analysis. For other periods, all countries are considered.

The dataset contains 84 deposit-taking commercial banks. There are two types of banking services: commercial banking and investment banking. Commercial banks take deposits and make loans and hence are conventional banks. However, typical investment banks do not take deposits but engage in intermediate securities trading or invest by themselves. My primary focus in this paper is on deposit-taking commercial banks because they are core members of the financial system. However, I shall also consider investment banks when comparisons across types of banking services are useful.

As a default risk measure, I use the credit default swap (CDS) premium. Because CDS contracts are similar to credit insurance against default risk, the corresponding premium represents the premium for default risk, making this measure an ideal measure of default risk. The CDS premium is determined in the market where CDS protection sellers and protection buyers participate and make bids and offer premia. One may also consider other measures such as the spread between the bond yield and the benchmark rate. However, it is widely accepted in the literature that the CDS premium is a better option than the yield spread for a number of reasons. First, the CDS premium is a direct measure of the default risk, while the yield spread is an indirectly constructed measure. Second, the yield spread may be flawed because it is often difficult to find an ideal benchmark rate that perfectly matches the duration. Blanco *et al.* (2005) presents a discussion of why the CDS premium is a more appropriate measure of the default risk.

Table 1 shows descriptive statistics of the variables of interest. Bank CDS is the premium on the CDS contract that protects the CDS holder from the default risk of the bank bond. Sovereign CDS is the premium on a CDS contract that hedges the default risk of a sovereign bond. In the following empirical analysis, noncore

TABLE 1— DESCRIPTIVE STATISTICS

Variables	Unit	Sample size	Mean	S.D.	Min	Max
Bank CDS	bp	1,938	91.0	94.7	9.8	1,421.5
Sovereign CDS	bp	2,372	59.4	139.5	8.4	2,826.7
Noncore currency	Dummy	2,437	0.24	0.43	0	1
Size to revenue	%	1,212	353.4	131.3	152.1	619.9
Private credit	% of GDP	2,269	176.5	45.3	107.5	401.6
NA CDS index	Bp	2,329	71.4	18.3	45.2	145.2
EU CDS index	Bp	2,305	72.5	23.4	44.2	173.0
Total assets	Million $	1,230	557,220	545,763	34,823	3,125,813
BIS	%	1,053	13.4	2.9	1.5	25.1
Leverage	%	1,230	28.2	14.9	2.1	85.5
ROA	%	1,161	0.34	0.66	-4.49	2.23

currency, size to revenue, and private credit are considered as important with regard to hypotheses 2 to 4. Noncore currency is a dummy variable whose value equals 1 if the currency of a given country is a noncore currency and 0 otherwise. Among all observations, advanced countries with a noncore currency comprise 24%, while the remaining observations are of advanced countries with a core currency. Size to revenue refers to the ratio of total assets in the banking sector to government revenue. The NA CDS index is an index that consists of premiums on major CDS contracts traded in the North American CDS market. These major CDS contracts include highly liquid sovereign CDS contracts, bank CDS contracts, and CDS contracts on nonbank companies. The EU CDS index is an index that consists of premiums on major CDS contracts traded in the European CDS market.

Sovereign CDS may depend on a number of factors, such as geopolitical shocks, macroeconomic shocks, and fiscal stability. Presumably, sovereign CDSs respond quickly and sensitively to unexpected geopolitical shocks, such as North Korean missile risks. In contrast, fiscal stability may have a relatively mild but long-standing impact on sovereign CDSs. In this regard, I shall conduct a preliminary analysis to examine whether the ratio of government debt to GDP has any effect on sovereign CDSs for the entire sample period, including crisis periods. In most cases, CDS premia are non-stationary variables. Thus, a regression based on CDSs could be highly likely to be spurious. I conduct unit-root tests of all sovereign CDSs, with the result suggesting that almost all CDSs have unit roots. However, the test results also suggest that almost no sovereign CDSs have unit roots when I consider the first difference in the CDSs. These test results imply that the sovereign CDS is an I(1)-variable.

Model 1 in Table 2 shows the result of a preliminary regression analysis in which the dependent variable is the log difference in the sovereign CDS and the independent variable of interest is the difference in government debt relative to GDP. I also include the intersection term of the *difference* in the government debt to GDP and the *level* of the government debt to GDP to examine whether there is a quadratic relationship between sovereign CDS and government debt.[3] I consider a quadratic relationship because the literature finds that debt and GDP growth or other variables

[3]If a variable y has a quadratic relationship with x, i.e., $y = a + bx + cx^2$, the total differentiation yields $\Delta y = b\Delta x + 2cx\Delta x$. In this regard, I include the intersection term of the *level* and the *difference* of the government debt to GDP in the quadratic regression analysis.

TABLE 2—GOVERNMENT DEBT AND SOVEREIGN CDS

	Model 1 (First difference)	Model 2 (Cointegration)
Independent var. \ Dependent var.	Δ log Sovereign CDS	Sovereign CDS
Δ Government debt to GDP	0.01902*** (0.00570)	-
Δ Government debt to GDP * Government debt to GDP	-0.00014** (0.00005)	-
Government debt to GDP	-	16.9526*** (3.7517)
Government debt to GDP^2	-	-0.0968*** (0.0179)
Year-Quarter Fixed Effect	Yes	-
Country Fixed Effect	Yes	-
Observation (Total/Country)	1,026/26	1,185/26
R-squared	0.5326	0.6935
Threshold	69.7%	87.5%
Panel Cointegration Test	-	Passed

Note: 1) The first model examines a quadratic relationship between the log difference in sovereign CDS and the difference in the ratio of government debt to GDP. If there is a quadratic relationship between y and x, i.e., if $y = a + bx + cx^2$, the total differentiation yields $\Delta y = b\Delta x + 2cx\Delta x$. Based on this observation, I use the intersection term of the difference and the level of government debt to GDP as an independent variable, 2) The second model examines a panel cointegration relationship between the level of sovereign CDS and the level of government debt to GDP, 3) *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 4) The threshold is defined as $-b/2c$, which is the peak of the quadratic relationship, where b is the coefficient estimate of the difference in government debt to GDP in model 1 and the coefficient estimate of the level of government debt to GDP in model 2. c denotes the coefficient estimates of the corresponding interaction terms, 5) The panel cointegration test is 'passed' if the Kao test rejects the null hypothesis that there is no cointegration.

of interest often have nonlinear inverse U-shaped relationships.[4] The estimation result suggests that the higher government debt becomes, the higher the sovereign CDS as long as the debt level is less than the threshold of 69.7%.

Some readers may wonder if the ratio of government debt to GDP affects the *level* of sovereign CDS. I conduct a panel cointegration analysis to determine whether there is any long-term relationship between the ratio of government debt to GDP and the level of sovereign CDS. The Cao Panel cointegration test suggests that there exists a quadratic cointegrating relationship between these two variables. Model 2 in Table 2 demonstrates that the level of sovereign CDS is positively associated with the ratio of government debt to GDP unless the debt level exceeds the threshold of 87.5%. This finding is qualitatively consistent with the regression result of model 1.

Below, I conduct the main analysis which examines the influence of sovereign CDS on bank CDS. To this end, I estimate the following fixed-effect regression model:

$$(1) \qquad \Delta \log BankCDS_{it} = \alpha + \beta \Delta \log SovereignCDS_{it} + \gamma' x_{it} + \theta_i + \delta_t + \varepsilon_{it}$$

Here, θ_i and δ_t denote the bank fixed effect and the year-quarter fixed effect.

[4]See Arcand *et al.* (2012), Cecchetti and Kharroubi (2012), and Cournède and Denk (2015).

x_{it} is a set of control variables. In this paper, the NA CDS index and the EU CDS index are important control variables. As noted earlier, this paper aims to study the effects of the sovereign default risk on the bank default risk. If there are confounding factors that affect both risks, this influence cannot easily be captured by a regression analysis. For instance, the overall business cycle could affect both sovereign CDS and bank CDS. That is, both sovereign CDS and bank CDS are low when the economy is booming, whereas both are high when the business cycle enters a trough. I use the two CDS indices to control for these confounding factors. Because the NA and EU CDS indices consist of nearly all major CDS contracts in the world, they are expected to capture the co-movement of sovereign CDS and bank CDS as driven by the business cycle or other confounding factors. However, the co-movement of CDS indices with banks may be heterogeneous for each bank. Thus, I shall estimate the bank-specific coefficients of CDS indices using the intersection vector of the bank fixed effect and CDS indices.

At times, I also consider bank-level information on balance sheets or income statements, such as total assets, the BIS capital ratio, the ratio of total debt to total assets (i.e. the leverage ratio), and the ROA ratio. However, this information is available only for publicly listed banks in my dataset, while there are many non-listed banks in the dataset.

Table 3 shows the estimation results for a number of model specifications. In all models, it is found that sovereign CDS and bank CDS are positively associated. I consider model 1 as the benchmark because it controls for the two important control variables, the NA CDS index and the EU CDS index, and it utilizes a large sample. The result based on model 1 suggests that a 1% increase in sovereign CDS leads to a 0.33% increase in bank CDS and that this association is statistically significant at the 1% level. This empirical result suggests that hypothesis 1 is acceptable.

Models 2 and 3 contain bank-level financial statement information, which is available only for listed banks and hence utilize a smaller sample set. In these models, the magnitude of the fiscal-financial relationship is even larger at 0.43%. This result is also statistically significant at the 1% level irrespective of whether the standard error is robust or clustered at the bank level.

Occasionally, sovereign CDS is affected by geopolitical shocks. For instance, South Korean sovereign CDS tends to increase sharply when North Korea launches a missile. To control for such time-varying country-specific factors, I use an intersection vector of the year-quarter fixed effect and the country fixed effect. Model 4 shows that sovereign CDS is still positively associated with bank CDS.

Although I consider important control variables such as CDS indices and time-varying country-specific factors, one concern is possible reverse causality or the influence of uncontrolled confounding factors. However, it is found that the *current sovereign* CDS is positively associated with the *forwarded* (i.e. future) *bank* CDS (in model 5), whereas the *current bank* CDS does not have a statistically significant relationship with the *forwarded sovereign* CDS. These findings suggest that at least during non-crisis periods, fiscal stability has a meaningful impact on financial stability, but not vice versa.

As an additional robustness check against reverse causality or confounding factors, I use the system GMM approach, which utilizes lagged variables as instrumental variables. In this case, I do not rule out crisis periods because instrumental variables

TABLE 3—BASELINE RESULTS

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
	Baseline	Financial information	Financial information	Time-varying country shock	Forwarded dependent variable	Reverse effect
Dependent var. / Independent var.	Δ log Bank CDS	Δ log Bank CDS	Δ log Bank CDS	Δ log Bank CDS	Forwarded Δ log Bank CDS	Forwarded Δ log Sovereign CDS
Δ log Sovereign CDS	0.3281*** (0.0341)	0.4329*** (0.0603)	0.4328*** (0.0790)	0.2577*** (0.0403)	0.1952*** (0.0271)	-
Δ log Bank CDS	-	-	-	-	-	0.0396 (0.0256)
Δ log NA CDS Index × Bank Fixed Effect	Yes	Yes	Yes	-	Yes	Yes
Δ log EU CDS Index × Bank Fixed Effect	Yes	Yes	Yes	-	Yes	Yes
Year-Quarter Fixed Effect × Country Fixed Effect	-	-	-	Yes	-	-
Δ log assets	-	-0.0789 (0.1609)	-0.0789 (0.1731)	-	-	-
Δ log Leverage	-	0.0492 (0.0253)	0.0492 (0.0214)	-	-	-
Δ log BIS	-	-0.1364 (0.1717)	-0.1634 (0.1823)	-	-	-
Δ log ROA	-	0.0025 (0.0188)	0.0025 (0.0183)	-	-	-
Year-Quarter Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes
Bank Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes
Observation (Total/Bank/Country)	1,762/78/19	607/40/14	607/40/14	1,865/78/19	1,750/78/19	1,765/78/19
R-squared	0.5957	0.6682	0.6682	0.7974	0.5685	0.6744
Standard error	Robust	Robust	Bank-cluster	Bank-cluster	Bank-cluster	Bank-cluster

Note: 1) The dependent variable is the log difference in bank CDS except for models 5 and 6, 2) The numbers in parentheses are the Huber-White-Sandwich robust standard error or the bank-level clustered error. *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 3) In model 5, the dependent variable is the one-period *forwarded* log difference in *bank* CDS, 4) In model 6, the dependent variable is the one-period *forwarded* log difference in *sovereign* CDS and the independent variable of interest is the log difference in bank CDS.

are used. Table 4 shows the estimation result based on the system GMM approach. I consider three model specifications. In the first model, I use the second and third lags of sovereign CDS, with the lagged dependent variable as an instrumental variable. In the second model, I use the second and third lags of sovereign CDS, the lagged dependent variable, and the NA and EU CDS indices as instrumental variables. In the third model, I use the second, third, and fourth lags of sovereign CDS, the lagged dependent variable, and the NA and EU CDS indices as instrumental variables. AR(1), AR(2), and Hansen test results suggest that all three models are properly specified. The estimation result suggests that an increase in sovereign CDS leads to an increase in bank CDS and that this positive association is statistically significant at the 1% level. This result is robust to model specification.

Thus far, I have used the first difference in CDS to deal with the non-stationarity problem. Another way to address this problem is to use the level of CDS and to conduct a cointegration analysis. Although this approach is less conventional in the corporate finance literature, I examine a cointegrating relationship in an attempt to check the robustness of the main result. As noted earlier, the unit-root tests suggest that almost all sovereign and bank CDSs are I(1) variables. The Kao panel

TABLE 4—SYSTEM GMM RESULTS

Variables	Model 1	Model 2	Model 3
Δ log Sovereign CDS	0.5925*** (0.0170)	0.5891*** (0.0371)	0.5778*** (0.0368)
Lagged dependent variable	Yes	Yes	Yes
CDS Indices	Yes	Yes	Yes
IV: Lags	2, 3	2, 3,	2, 3, 4
IV: CDS Indices	-	Yes	Yes
Observation (Total/Bank/Country)	2,277/78/19	2,277/78/19	2,277/78/19
AR(1) test	Passed	Passed	Passed
AR(2) test	Passed	Passed	Passed
Hansen test of over-identification	Passed	Passed	Passed

Note: 1) The dependent variable is the log difference of bank CDS, 2) *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 3) Model 1 uses the second and third lags of sovereign CDS and the lagged dependent variable as instrumental variables, 4) Model 2 uses the second and third lags of sovereign CDS, the lagged dependent variable, and the NA and EU CDS indices as instrumental variables, 5) Model 3 uses the second, third, and fourth lags of sovereign CDS, the lagged dependent variable, and the NA and EU CDS indices as instrumental variables, 6) Each model passes the AR(1) test if the hypothesis that the first difference in the error term does not have a first-order serial correlation is rejected, 7) Each model passes the AR(2) test if the hypothesis that the first difference in the error term does not have a second-order serial correlation is not rejected.

TABLE 5—PANEL COINTEGRATION ANALYSIS

Variables	Model 1	Model 2	Model 3
Sovereign CDS	0.4467*** (0.0200)	0.3825*** (0.0374)	0.3347*** (0.0369)
Constant	Yes	Yes	Yes
Linear Trend	-	-	Yes
NA CDS Index	-	Yes	Yes
EU CDS Index	-	Yes	Yes
AIC automatic leads and lags selection	Yes	Yes	Yes
Observation (Total/Bank/Country)	2,256/78/19	2,231/78/19	2,223/78/19
R-squared	0.7202	0.7929	0.8356
Panel Cointegration Test	Passed	Passed	Passed

Note: 1) The dependent variable is the level of bank CDS, 2) In model 1, the independent variable is the level of sovereign CDS. In models 2 and 3, the independent variables are the levels of sovereign CDS, the North American CDS index, and the European CDS index, 3) *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 4) The numbers of leads and lags are automatically selected by the Akaike Information Criterion, 5) The panel cointegration test is 'passed' if the Kao test (or Phillips-Perron test when the linear trend is included) rejects the null hypothesis that there is no cointegration.

cointegration test suggests that there is a cointegrating relationship between sovereign and bank CDSs. This test result is robust to the inclusion/exclusion of a linear time trend and/or CDS indices. Table 5 shows the estimation result. All models show that sovereign and bank CDSs are positively cointegrated and that the estimated coefficient is similar to that in the baseline model (i.e., Model 1 in Table 3).

B. *Determinants of the Intensity of the Fiscal-financial Relationship*

1. High-risk vs. Low-risk Countries

One can argue that the baseline result is mainly driven by certain outlier countries,

TABLE 6—COMPARISON: HIGH-RISK VS. LOW-RISK

Variables	Model 1 (Low-risk country)	Model 2 (High-risk country)	Model 3 (Low-risk bank)	Model 4 (High-risk bank)
Δ log Sovereign CDS	0.2490*** (0.0388)	0.6115*** (0.0805)	0.3665*** (0.0401)	0.3597*** (0.0704)
NA and EU CDS Indices × Bank Fixed Effect	Yes	Yes	Yes	Yes
Year-Quarter Fixed Effect	Yes	Yes	Yes	Yes
Bank Fixed Effect	Yes	Yes	Yes	Yes
Observation (Total/Bank/Country)	1,386/71/17	376/31/8	1,253/70/18	509/52/14
R-squared	0.5843	0.7819	0.6427	0.6882

Note: 1) The dependent variable is the log difference of bank CDS, 2) The standard error is the Huber-White-Sandwich robust standard error. *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 3) Model 1 considers low-risk countries whose sovereign CDS premia are below the average of 59.4bp, while model 2 considers high-risk countries whose CDS premia are above the average, 4) Model 3 considers low-risk banks whose bank CDS premia are below the average of 91.0bp, while model 2 considers high-risk banks whose CDS premia are above the average.

such as Italy, Portugal, Spain, and Greece, which underwent from severe fiscal instability. To address this concern, I estimate the baseline empirical model separately for a high-risk group and a low-risk group, where the high-risk group consists of countries whose sovereign CDS premia are higher than the average CDS premium (59.4bp) and the remaining countries form the low-risk group.

Table 6 shows the estimation result. The result suggests that the intensity of the fiscal-financial relationship is greater in the high-risk group as opposed to the low-risk group. However, a positive and statistically significant association can be found in both groups. One can draw a policy implication for Korea from this result. Currently, South Korea is in the low-risk group because its sovereign CDS premium is 26.7bp on average as of 2020. However, if the expansionary fiscal policy due to the Covid-19 pandemic and the low fertility and mortality rates raise the sovereign default risk significantly in the future and, hence, Korea becomes a high-risk country, financial stability will be not merely undermined but will take on a larger magnitude.

2. Banks vs. Other Financial Companies

Another criticism is that the positive association between sovereign CDS and corporate CDS is not a special characteristic of the banking sector but a general property that can be observed in other financial sectors as well. This criticism is related to the argument that the corporate default risk generally increases with the sovereign default risk, as sovereign bonds are used as a benchmark when assessing the credit rating of corporate bonds. This criticism is also related to the notion that governments may bail out not just commercial banks but also other financial companies if they are too big to fail in terms of size and inter-connectedness.

To address this concern, I estimate the same empirical model for other financial companies, such as real estate firms, insurance companies, and investment banks. Table 7 shows that a 1% increase in sovereign CDS premia is associated with at most a 0.19% increase in the CDS premia of other financial corporate bonds. In comparison, the same 1% increase in sovereign CDS leads to a 0.33% increase in bank CDS. Although a positive association between sovereign CDS and corporate

TABLE 7—BANK VS. OTHER FINANCIAL COMPANIES

Industry	Coefficient	Observation (Total/Firm/Country)	R-squared
Commercial Banking	0.3281***	1,762/78/19	0.5957
Real Estate	0.1859***	1,271/44/6	0.4178
Insurance	0.1715***	1,727/58/9	0.3464
Investment Banking	0.0886***	2,829/117/16	0.3516

Note: 1) The dependent variable is the log difference of corporate CDS, 2) The standard error is the Huber-White-Sandwich robust standard error. *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 3) Coefficient denotes the estimated coefficient of the log-differenced sovereign CDS, 4) For each sector regression, the product of the NA CDS index and firm fixed effect, the product of the EU CDS index and firm fixed effect, the firm fixed effect, and the year-quarter fixed effect are controlled.

CDS is observed in both the commercial banking sector and other financial sectors, the magnitude is substantially different, presumably because commercial banks play much more important roles than other financial companies in the overall economic system.

3. Moral Hazard Issue: Commercial Banks vs. Investment Banks

An additional criticism relates to the issue of moral hazard. A well-known and long-standing problem in the banking industry is that bank managers can fall prey to moral hazard under the presence of implicit government guarantees. Because banks believe that they will be rescued by the government when they fail, they have an incentive to take excessive risks ex-ante. However, if the government's promise to bail them out is less trustworthy, bank managers may be less tempted to take excessive risks. That is, if the sovereign default risk rises, the bank default risk increases due to the weakened government support on the one hand (i.e., *direct effect*), while the bank default risk may also decrease because the moral hazard problem of bank managers is relaxed on the other hand (i.e., *indirect effect*). Therefore, the relationship between the sovereign and bank default risks is a matter to be tested empirically.

A comparison of commercial banks and investment banks in terms of the intensity of the fiscal-financial relationship could be meaningful in light of the moral hazard issue. Commercial banks have standard business models that include deposit-taking and loan-making activities and hence there is little room to take excessive risks. In contrast, investment banks typically do not take deposits but rely on riskier wholesale funding and invest in stocks, bonds, derivatives, and other riskier products. Therefore, all other factors being equal, one can expect that investment banks are more exposed to the moral hazard problem than commercial banks. Given this expectation, I estimate the empirical model in (1) separately for commercial banks and investment banks.

Table 7 shows the estimation result. The result shows that the intensity of fiscal-financial relationship is as high as 0.3281 for commercial banks, while the intensity is only 0.0886 for investment banks. The intensity outcome for investment banks implies that the direct effect of an increase in the sovereign default risk on the bank default risk dominates the indirect effect. For investment banks, one can expect both direct and indirect effects, as investment bankers have more room to take risks. However, the empirical result shows that the direct effect outweighs the indirect effect and, hence, a positive association between the dependent and independent variables is observed. With regard to commercial banks, in contrast, one can expect only a direct effect, as commercial bank managers have little room to take risks. The empirical result is consistent with this expectation, as the intensity of the fiscal-financial relationships for commercial banks is substantially greater than the intensity for investment banks.

This empirical finding does not fully address the moral hazard issue because commercial banks and investment banks may differ not simply in terms of the extent of risk-taking but also in terms of many other factors. Although this paper considers a number of important control variables, such as the North American and European CDS indexes, bank fixed effects, and time fixed effects, I hope to see more accurate comparisons in future research.

Thus far, I have shown that the greater the sovereign default risk is, the greater the bank default risk becomes. Because this result holds generally for all advanced countries, it may not provide Korea-specific implications. The Korean economy can be differentiated from those of other advanced countries in three senses. First, its currency is not a core currency. Secondly, the size of its banking sector relative to government revenues is very high. Last but not least, the ratio of private credit to GDP in Korea is relatively large compared to other countries. In the following analysis, I shall test hypotheses 2 to 4 to determine the important determinants of the fiscal-financial relationship, through which I shall draw Korea-specific implications.

4. Noncore Currency

According to hypothesis 2, an increase in the risk of a sovereign default leads to a greater increase in the risk of a bank default if the given country does not have a core currency. In order to test this hypothesis, I estimate the empirical model in (1) separately for counties with a noncore currency and those with a core currency.

Table 8 shows the estimation result. The results for models 1 and 2 suggest that in countries with a noncore currency, a 1% increase in sovereign CDS leads to a 0.76% increase in bank CDS. This magnitude is substantial given that the same 1% increase in sovereign CDS in countries with a core currency is associated only with a 0.29% increase in bank CDS. That is, the empirical result is consistent with hypothesis 2. Model 3 explicitly tests whether the difference in the magnitude is statistically significant using an interaction term for a noncore currency and sovereign CDS, with the result showing significance at the 1% level. Korea is one of the eight advanced countries here whose currency is not a core currency. This empirical analysis implies that Korea should place more emphasis on maintaining fiscal stability than countries with a core currency because the same weakening of fiscal stability could cause greater damage to the financial system.

Although I assume that US dollar, euro, UK pound sterling, Japanese yen, Swiss franc, and Canadian dollar as core currencies following the opinion of the Bank of Korea (2018), there is no exact and official standard of a core currency. Some may suggest that the Australian dollar should also be considered as a core currency. Others may believe that only the two most important currencies (i.e., the US dollar and euro) are core currencies. In addition, the euro may be special because it is commonly used by a number of different EU countries. Given this disparity in the

TABLE 8—NONCORE CURRENCY

Variables	Model 1 (Noncore)	Model 2 (Core)	Model 3 (All)
Δ log Sovereign CDS	0.7575*** (0.0655)	0.2862*** (0.0437)	0.2884*** (0.0351)
Noncore currency dummy	-	-	Yes
Δ log Sovereign CDS × Noncore currency dummy	-	-	0.3596*** (0.0493)
NA and EU CDS Indices × Bank Fixed Effect	Yes	Yes	Yes
Year-Quarter Fixed Effect	Yes	Yes	Yes
Bank Fixed Effect	Yes	Yes	Yes
Observation (Total/Bank/Country)	480/16/5	1,282/62/14	1,762/78/19
R-squared	0.7660	0.5911	0.6092

Note: 1) The dependent variable is the log difference of bank CDS, 2) The standard error is the Huber-White-Sandwich robust standard error. *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 3) Model 1 considers only countries with a noncore currency, 4) Model 2 considers only countries with a core currency, 5) Model 3 considers all countries.

TABLE 9—NONCORE CURRENCY: ROBUSTNESS CHECK

Variables	Model 1 (Baseline model)	Model 2 (Australia added)	Model 3 (US Dollar, Euro only)	Model 4 (Euro is missing)
Δ log Sovereign CDS	0.2884***	0.3052***	0.2508***	0.3494***
Noncore currency dummy	Yes	Yes	Yes	Yes
Δ log Sovereign CDS × Noncore currency dummy	0.3596***	0.3158***	0.2276***	0.2606***
NA and EU CDS Indices × Bank Fixed Effect	Yes	Yes	Yes	Yes
Year-Quarter Fixed Effect	Yes	Yes	Yes	Yes
Bank Fixed Effect	Yes	Yes	Yes	Yes
Observation (Total/Bank/Country)	1,762/78/19	1,762/78/19	1,762/78/19	918/36/9
R-squared	0.6092	0.6028	0.6057	0.6647

Note: 1) The dependent variable is the log difference of bank CDS, 2) The standard error is the Huber-White-Sandwich robust standard error. *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 3) In model 1, I define core currencies as the US dollar, euro, UK pound sterling, Japanese yen, Swiss franc, and Canadian dollar, 4) In model 2, I add the Australian dollar to the group of core currencies, 5) In model 3, I assume that only the US dollar and euro are core currencies, 6) In model 4, I use the same definition of core currencies used in model 1. However, Eurozone member countries are excluded in the regression analysis in this case.

definition of a core currency, I conduct a number of robustness checks while varying the group of core currencies. Table 9 shows that a positive and significant association between sovereign and bank CDSs is still observed when the Australian dollar is included (Model 2), when only the US dollar and euro are considered (Model 3), and when euro member countries are excluded from the sample (Model 4).

5. Banking Sector Size to Government Revenue

According to hypothesis 3, if the risk of sovereign default rises, the risk of bank default increases at a greater magnitude in countries where the size of the banking sector relative to government revenues is relatively large. To test this hypothesis, I divide countries into two groups: a large size group consisting of countries where the ratio of total assets in the banking sector to government revenues is higher than

TABLE 10—BANKING SECTOR SIZE TO GOVERNMENT REVENUES

Variables	Model 1 (Large size)	Model 2 (Small size)	Model 3 (All)
Δ log Sovereign CDS	0.2978*** (0.0405)	0.2476*** (0.0805)	0.0051 (0.1095)
Size to revenue	-	-	Yes
Δ log Sovereign CDS × Size to revenue	-	-	0.0013*** (0.0003)
NA and EU CDS Indices × Bank Fixed Effect	Yes	Yes	Yes
Year-Quarter Fixed Effect	Yes	Yes	Yes
Bank Fixed Effect	Yes	Yes	Yes
Observation (Total/Bank/Country)	1,324/77/19	438/47/14	918/66/17
R-squared	0.6300	0.6610	0.6788

Note: 1) The dependent variable is the log difference of bank CDS, 2) The standard error is the Huber-White-Sandwich robust standard error. *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 3) Model 1 considers only countries in which the ratio of total assets in the banking sector to government revenues is above the country average of 353%, 4) Model 2 considers only countries in which the ratio of total assets in the banking sector to government revenue is below the country average of 353%, 5) Model 3 considers all countries.

the country average of 353%, and a small size group formed by countries whose assets-to-revenue ratio is lower than the average.

Models 1 and 2 in Table 10 show the estimation result, which suggests that an 1% increase in sovereign CDS leads to a 0.30% increase in bank CDS in the large size group, while the same increase in sovereign CDS leads to only a 0.25% increase in bank CDS in the small size group. That is, the empirical result is consistent with hypothesis 3. Model 3 finds that this difference in magnitude of the fiscal-financial relationship is statistically significant at the 1% level. This finding is meaningful for Korean economy because the Korean banking sector (relative to government revenues) is larger than that of any other advanced country. That is, if government revenues were the only source of funding for bank bailouts, the Korean banking sector would be too big to save and, hence, a weakening of fiscal stability could impose a greater cost on the financial system.

6. Private Credit

A characteristic of the Korean economy is also its high level of private credit. According to hypothesis 4, with a greater intensity of the fiscal-financial relationship, a greater ratio of private credit to GDP exists. To test this hypothesis, I divide countries into a large credit group and a small credit group, where the ratio of private credit to GDP is higher than the country average of 177% in the former while it is lower than the average in the latter.

Models 1 and 2 in Table 11 demonstrate that a weakening of fiscal stability causes relatively more damage to the financial system in the large credit group as opposed to the small credit group.

TABLE 11—PRIVATE CREDIT TO GDP

Variables	Model 1 (Large credit)	Model 2 (Small credit)
Δ log Sovereign CDS	0.4207*** (0.0541)	0.2994*** (0.0524)
Private credit to GDP	-	-
Δ log Sovereign CDS × Private credit to GDP	-	-
NA and EU CDS Indices × Bank Fixed Effect	Yes	Yes
Year-Quarter Fixed Effect	Yes	Yes
Bank Fixed Effect	Yes	Yes
Observation (Total/Bank/Country)	854/69/18	908/53/11
R-squared	0.6355	0.6192
Standard error	Robust	Robust

Note: 1) The dependent variable is the log difference of bank CDS, 2) The standard error is the Huber-White-Sandwich robust standard error. *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively, 3) Model 1 considers only countries in which the ratio of private credit to GDP is above the country average of 177%, 4) Model 2 considers only countries in which the ratio of private credit to GDP is below the country average of 177%, 5) Model 3 considers all countries.

C. *Determinants of the Implicit Government Guarantee*

Thus far, I have empirically examined how fiscal instability influences financial instability. At the core of this fiscal-financial link are the implicit government guarantees in the banking sector. In the preceding analysis, I use the coefficient of sovereign CDS as a measure of this implicit guarantee.

Another measure of the implicit guarantee is the *uplift*, which is the difference between the final credit rating and the stand-alone credit rating, where the latter is the assessment of the creditworthiness of a debtor assuming away the possibility of a government bailout. With regard to assessing bank credit ratings, Moody's initially assesses the stand-alone rating and then adjusts the credit rating by considering the possibility and magnitude of a government bailout in case of a bank failure. Therefore, the uplift metric is a direct measure of the implicit guarantee.

Figure 3 shows the uplift for each advanced country. I calculate the bank uplift value by taking the time-average of uplifts of banks for the period of 2011 to 2020 and then calculate the country uplift value by taking the average of bank uplifts for each country. The uplift for Korea is 4.5 notches, which is highest among all countries. The magnitude (4.5 notches) is substantial. Raising the credit rating by only one notch is a challenging task for debtors. Sometimes even very creditworthy debtors fail to raise this metric by one or two notches despite their diligent efforts over many years. This implies that the implicit guarantee by the Korean government for Korean banks is very powerful. In other words, Korean banks rely heavily on the government.

In Figure 4, I focus on countries whose currency is a noncore currency in order to draw more Korea-specific implications. As of 2020, Korea remains the country with the highest uplift. Interestingly, the Korean banking sector's stand-alone rating is low though not the lowest in its comparison group. This implies that Korean banks are least capable of overcoming crises on their own but rely heavily on the government.

One can argue that according to the construction, uplift is decreasing in the stand-alone rating. However, this argument is flawed. Let u, f, and s denote the uplift,

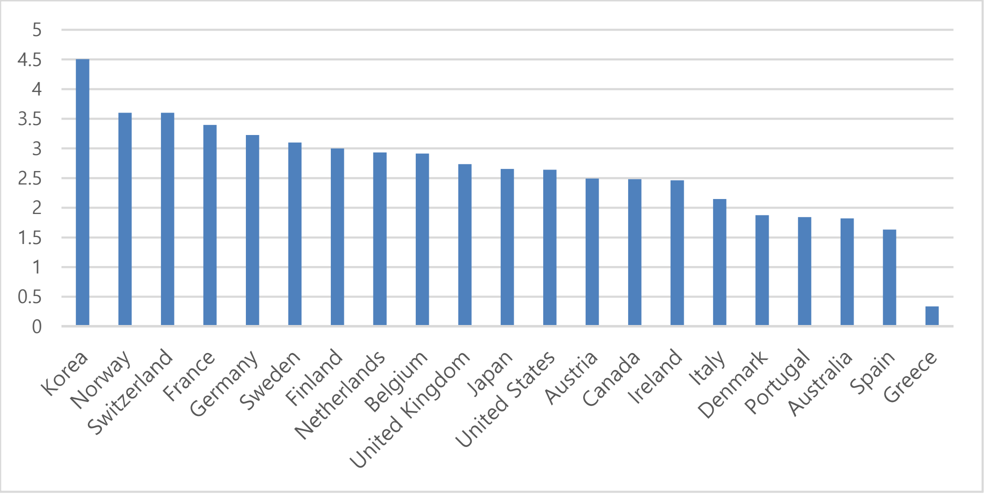

FIGURE 3. UPLIFT FOR 2011-2020

Note: I calculate the bank uplift metric by taking the time-average of uplifts of banks for the period from 2011 to 2020 and then calculate the country uplift by taking the average of the bank uplifts for each country.

Source: Author's calculation based on Moody's credit ratings.

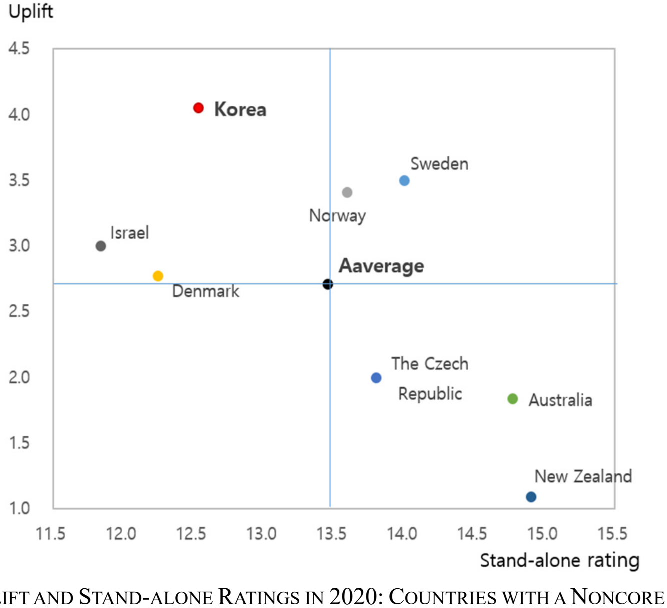

FIGURE 4. UPLIFT AND STAND-ALONE RATINGS IN 2020: COUNTRIES WITH A NONCORE CURRENCY

Note: For the stand-alone rating, one unit equals one notch in the credit rating. For instance, the highest credit rating (Aaa) is denoted as 20, the second highest (Aa1) is 19, and the lowest (C) is denoted as 0.

Source: Hwang (2021a).

final rating, and stand-alone rating. Because uplift u equals $f-s$ according to the construction, it appears to be decreasing with regard to the stand-alone rating s. However, Moody's in fact assigns the final credit rating only after adding some adjustments to the stand-alone rating, considering the possibility and magnitude of the implicit government guarantee. In this sense, the final rating f should equal $s+i$, where i denotes the intensity of the implicit guarantee. Then, the uplift u equals $s+i-s$, which is just equal to i. Therefore, the uplift does not directly depend on the stand-alone rating. However, the implicit guarantees $i=i(s)$ can indirectly depend on the stand-alone rating. For instance, if a systemically important bank becomes weaker and therefore its stand-alone rating is downgraded, the bank becomes more likely to fail and, hence, the ex-ante likelihood that the government provides a bailout for this bank may increase.

In order to examine what determines the implicit guarantee, I conduct an empirical analysis below. I manually collect stand-alone and final credit ratings from Moody's website. In this way, I construct a quarterly database of 29 countries and 84 deposit-taking commercial banks. The time span is 2011-2020 because the stand-alone rating is released only after 2011.

In the following sections, I examine how certain different independent variables are associated with the uplift metric. Table 12 shows the fixed-effect panel regression results. In every model specification, the estimation result suggests that the uplift is negatively associated with the stand-alone rating. The magnitude of this negative relationship is economically significant, as a one notch increase in the stand-alone rating is associated with a 0.24~0.33 notch decrease in the uplift. This finding implies that governments are more likely to save banks when the banks are relatively weak. If a bank's prudence is weaker on its own, it may need to rely more on a government guarantee to persuade investors that it is creditworthy. This finding is closely related to what is described in Figure 4, where in Korea the uplift is highest while the stand-alone rating is low.

The results for models 2 and 3 suggest that the implicit guarantees are weaker, the higher the government debt to GDP ratio. This is in line with the results of the previous analysis using CDS. If the fiscal space becomes weaker due to an expansion in government debt, fiscal stability is undermined. Therefore, the government's

TABLE 12—DETERMINANT OF IMPLICIT GOVERNMENT GUARANTEES

Variables	Model 1	Model 2	Model 3
Stand-alone rating	-0.2432*** (0.0310)	-0.2757*** (0.0351)	-0.3343*** (0.0359)
Government debt to GDP	-	-0.0280*** (0.0037)	-0.0119** (0.0054)
Log assets	-	-	0.8259*** (0.2383)
BIS	-	-	-0.0430*** (0.0126)
Leverage	-	-	0.0038 (0.0035)
ROA	-	-	-0.0134 (0.0501)
Year-Quarter Fixed Effect	Yes	Yes	Yes
Bank Fixed Effect	Yes	Yes	Yes
Observation (Total/Bank/Country)	2,437/84/21	2,084/84/21	865/47/17
R-squared	0.8562	0.8025	0.8444

Note: 1) The dependent variable is the uplift, 2) The standard error is the Huber-White-Sandwich robust standard error. *, **, and *** represent the 10%, 5%, and 1% levels of significance, respectively.

informal promise to bail out banks in distress is not trustworthy and, therefore, the value of the implicit guarantee is lower, causing investors to withdraw their trust from banks to some extent.

The coefficient estimates of assets and the BIS capital ratio also provide meaningful implications. The log assets result suggests that the value of the implicit guarantee is higher for a larger bank. This finding is consistent with the too-big-to-fail hypothesis. The BIS ratio result suggests that implicit guarantees are stronger for less prudent banks. This finding is in line with the observation that the uplift and stand-alone rating are negatively associated. That is, weaker banks can benefit more from the implicit guarantee.

The policy implications of the findings in Figure 4 and Table 12 are simple and clear: the Korean banking system is not very healthy on its own and relies heavily on the government's implicit guarantee. In order to improve their stand-alone competitiveness and health, Korean banks should improve their risk management and the Korean government simultaneously should maintain fiscal stability.

IV. Concluding Remarks

This paper analyzes how fiscal instability affects financial instability. Many are concerned with that the fiscal space in Korea will continue to shrink (IMF, 2021a). Fiscal instability is not merely a serious problem on its own but could also have negative spillover effects on the financial system according to the theoretical literature.

This paper empirically shows that an increase in the sovereign default risk leads to a rise in the bank default risk at an economically significant magnitude. I also take a closer look at the Korean fiscal-financial nexus and observe that Korea does not have a core currency, its banking sector is largest relative to government revenues among all advanced countries, and that private credit to GDP is higher than the average of other countries. The empirical analysis shows that all three of these characteristics contribute to magnify the intensity by which fiscal instability worsens financial instability.

The implicit government guarantee lies at the core of this adverse effect of fiscal instability on financial instability. I consider the difference between a bank's final credit rating and its stand-alone rating (i.e., uplift) as a direct measure of the implicit guarantee and find that a bank's reliance on the government is higher if the bank's stand-alone prudence is lower, its size is larger, or the fiscal space is smaller.

These empirical results imply that the Korean banking sector must reduce its reliance on the government's implicit guarantee. To this end, a special bank resolution regime based on the 'bail-in' concept should quickly be adopted. Under the bail-in regime, a failed bank will be resolved not with taxpayer's money but at the cost of the bank's creditors and shareholders. Given this bail-in regime, banks will not be bailed out by the government when they fail and, hence, the link of the fiscal-financial relationship will be cut. Therefore, even if fiscal stability is undermined, it may not cause a severe disruption in the financial system. Although Korea and other G20 countries agreed to adopt the bail-in regime in 2010 immediately after the 2008 Global Financial Crisis and considering that the United States, EU, Japan and other key jurisdictions adopted this new resolution regime thereafter, Korea remains hesitant with regard to its application.

This paper does not assert that the fiscal space in Korea is currently poor. Around the time of this publication, the CDS premium on Korean sovereign bonds is lower than the average of other advanced countries. However, it is inevitable that the fiscal space will be greatly reduced in the long run due to the low fertility and mortality rates in Korea and resulting spike in expenditures for pensions, health care, and other mandatory spending. Based on the empirical findings here, it should be stressed that we need to prepare to ensure a better future in terms of fiscal and financial stability in Korea.

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Job Creation during Korea's Transition to a Knowledge Economy†

By KYUNGSOO CHOI*

This paper analyzes job creation when the Korean economy transitioned to a knowledge economy from the 1990s to the 2010s. During this period, the ratio of service to manufacturing jobs increased, knowledge intensive industries grew, and job creation became geographically concentrated around Seoul. The changes slowed down in the 2010s, and overall job growth weakened. To analyze the effect of job creation driver industries during this period, the main part of which are knowledge intensive tradable service industries, on local service job creation, I use a modified version of the local labor market of Moretti (2010). I analyze the job changes during 1995-2005 and during 2006-2016 in 237 Si-Gun-Gu areas in the Census on Establishments datasets. I find that one manufacturing job creates 0.5 local service jobs and that one tradable service job creates 1.1 jobs within Gu areas of metro cities and 2.3 jobs in Si-Gun areas. The job creation relationship between the tradable and local service sectors was not altered in this period. As more jobs were created in the tradable sector driven by the transition to a knowledge economy, job creation overall remained active, with the opposite also being true.

Key Word: Korea, Job Creation, Local Labor Market Model, Knowledge Economy
JEL Code: J01, J21, J23, J24, R11, R12

I. Introduction

Job creation has emerged as a top-priority policy goal in Korea such that the incoming government referred to itself as a 'Jobs Government' in 2017.[1] Figure 1

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† This paper is a revised and updated version of Chapter 1 in Kyungsoo Choi (ed.), *The Jobs Creation Strategy in a Knowledge-Based Economy*, Research Monograph 2019-12, KDI, 2019 (in Korean).

[1]The incumbent government set up the 'Presidential Committee on Jobs' on May 17, 2017, where the chairman was the president himself. Setting up the committee was Presidential Order number 1 after the president's

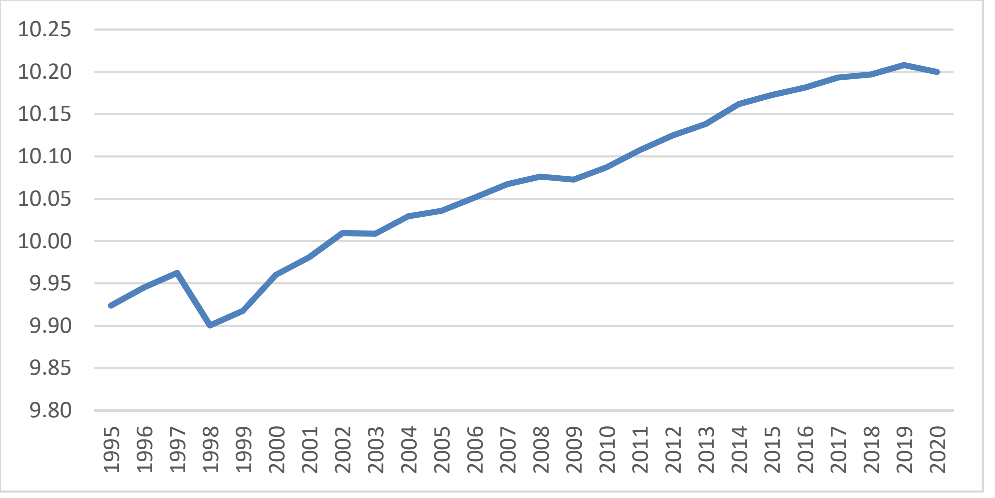

FIGURE 1. KOREA'S EMPLOYMENT TREND ON A LOG SCALE: 1995-2020

Note: Natural log of all industry workers in thousands.

Source: Korea Statistics, Economically Active Population Survey.

presents Korea's employment trend on a log scale from 1995 up to 2020.[2] Korea has overcome two major economic crises, in 1998 and in 2008, and sustained its job growth path, but job growth has weakened noticeably during the last decade. There were heated debates over the proper strategies to use for job creation in 2017. This research is motivated by these debates.

The Korean economy has been a typical manufacturing economy, and its comparative advantage has shifted from low wages to technology over the past few decades. At the same time, the driving force of job creation has moved from manufacturing to knowledge industries. The objective of this paper is to show how jobs were created during the transition to a knowledge economy in Korea in the period from the mid-1990s to the 2010s.

Labor demand functions presume a stable relationship between output and labor requirement and derive labor demand levels from production. However, in a knowledge economy, knowledge-intensive jobs in the tradable sector create more jobs outside the sector through job creation spillover effects.[3] The labor demand framework does not take into account these job-creation effects, because it derives the amount of labor needed for production, which is a rather small percentage in its total job-creation effect. Input-output tables calculate labor requirement in related industries. They do not address the spillover effects on unrelated industries. Thus, I use a local labor market job creation model, suggested by Moretti(2010), to analyze job creation during this transition period. This framework divides the economy into two sectors: the tradable sector and the non-tradable sector. Jobs in the tradable sector are created when they are productive enough, while jobs in the non-tradable sector are created by the spillover effects of job in the tradable sector. This paper

inauguration on May 10, 2017.

[2]The unit in Figure 1 is the natural log of thousands of workers in all industries.

[3]See Moretti (2012) for reference.

addresses how the tradable sector changed as the economy transitioned to a knowledge economy and how the spillover effect acted during this period.

This paper has two parts. In the first part, Section II, I address the question of how one can infer that there has been a job creation mechanism change in Korea. The answer to this question is naturally related to the issue of how the tradable sector changed as the Korean economy transitioned to a knowledge economy. There have been conspicuous changes, such as the ratio of service to manufacturing jobs, geographical job creation patterns, and the growth of knowledge-intensive industries. One question long in the author's mind concerns why more jobs are created when manufacturing jobs decrease. This question cannot be answered if the decrease in the number of manufacturing workers is accompanied by a transformation of manufacturing and an increase in knowledge-intensive industry workers, and the latter has a larger spillover effect. Topics in Section II may appear to be less interrelated; nonetheless, the author had this question in mind.

The second part, in Sections III and IV, addresses the question of how the spillover effects changed during this transition period. In order to answer this question, I introduce the local labor market model in Section III and show how the elements are constructed along with their trends. Section IV performs a regression analysis of the model proposed in Section III and discusses the estimation results.

This research has the following policy implications. If both the strong job creation in the 2000s and its weakening in the 2010s are related to the growth of knowledge-intensive industries, which are closely related to manufacturing, which policies affect their sustained growth? There are many proposals in the literature domestic and abroad, but further investigations of our own problem appear to be necessary. According to the model in this paper, local services such as restaurants and small shops have little spillover effects on other industry jobs, allowing room to manoeuver by policymakers on issues such as protection of traditional markets and the minimum wage, among others. Section V presents the summary and conclusion.

II. Shifts in Job Creation Patterns

A. *Employment by Industry*

At an early stage of economic development, manufacturing usually serves as the major driving force of job creation. As the economy matures, this role moves to the service industries. Table 1 and Figure 2 show the numbers of manufacturing and service jobs and their ratios between 1970 and 2020. The 1970s and 1980s represented an industrialization and growth spurt era for Korea. The number of manufacturing jobs nearly doubled in a decade, as indicated by the dotted line in Figure 2. For each manufacturing job, two service jobs were created in a parallel fashion such that the ratio between the two remained around 2.0 in this period (Table 1, row 3). From the 1990s, service jobs continue to grow without being accompanied by manufacturing job growth. Consequently, the ratio rises, reaching 4.5 in 2010. Afterwards, the rise becomes stagnant and overall job growth weakens.

The growth of service jobs in the 1990s and 2000s suggests that some force other than manufacturing may have driven job creation. This shift is the main subject of

TABLE 1—MANUFACTURING AND SERVICE JOBS: 1970-2020

(Unit: millions)

	1970	1975	1980	1985	1990	1995	2000	2005	2010	2015	2020
Manufacturing	1.3	2.2	3.0	3.5	4.9	4.8	4.3	4.1	4.1	4.6	4.4
Service	3.4	4.1	6.0	7.6	9.9	13.2	14.6	16.9	18.3	20.2	21.1
Ratio	2.7	1.9	2.0	2.2	2.0	2.7	3.4	4.1	4.5	4.4	4.8

Note: Service is all industry except for agriculture, quarrying, mining, and manufacturing.

Source: Author's calculation from Korea Statistics, Economically Active Population Survey.

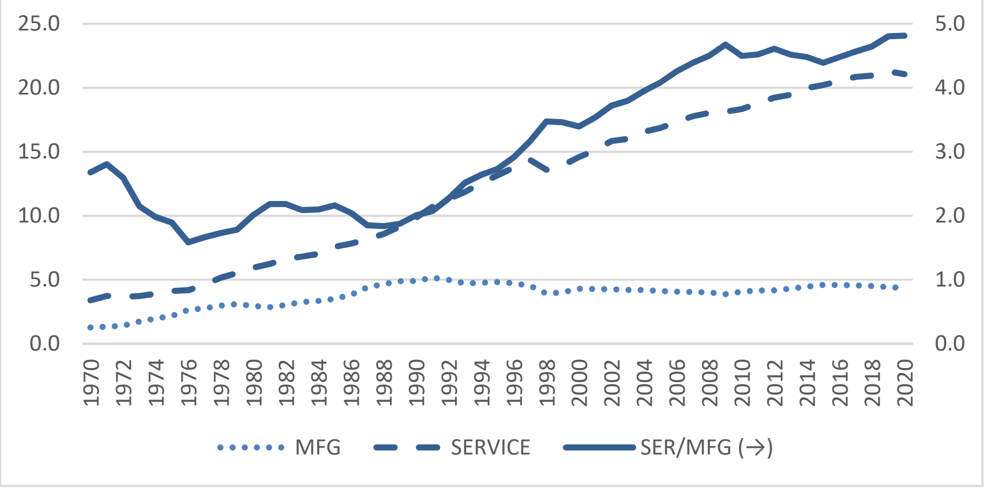

FIGURE 2. SERVICE AND MANUFACTURING EMPLOYMENT: 1970-2020

Source: Author's calculations from Korea Statistics, National Accounts, GDP by Economic Activities.

this paper. On the production side, noticeable is the growth of knowledge-based industries in this period. The growth of knowledge-based industries is one of the major changes that occurred in advanced economies in the late 20th century.[4] The OECD notes that 'economies are increasingly based upon information and knowledge as drivers of productivity and growth' (OECD, 1996, p.3). The quantitative content of this claim is the rising share of knowledge-based industries, which are technology-intensive manufacturing and knowledge-intensive service industries, in overall output. The OECD report further notes, "Indeed, it is estimated that more than 50 percent of Gross Domestic Product (GDP) in the major OECD economies is now knowledge-based" (p.9).[5]

In the literature, knowledge-based industries are broadly defined without a single consensus. Cermeño (2018) observes that the share of knowledge-based business services (KIBS) in GDP increased consistently over the very long period of 1950 to 2010. Cermeño (2018, Figure 1, p.146) shows that its share rose from 18 percent in

[4]The literature is not explicit with regard to the timespan of the 'knowledge-based economy.' Figure 1 in the OECD report (1996, p.10) shows the share of OECD high-tech exports starts its rapid ascent in the 1980s. If this share is a proper measure, the transition accelerated in the 1980s and 1990s.

[5]However, the definition of knowledge-based industries is not very specific.

TABLE 2—SHARES OF KNOWLEDGE INTENSIVE BUSINESS SERVICES IN TOTAL VALUE ADDED: 1985-2020

(Unit: %)

	1985	1990	1995	2000	2005	2010	2015	2020
Manufacturing	26.9	27.7	28.3	29.3	28.4	30.2	29.0	27.1
Tech-intensive	11.4	13.6	16.0	17.4	17.3	19.5	19.1	18.6
KIBS	12.9	15.5	17.9	18.9	22.0	22.7	23.0	24.3
Health, social work	2.0	1.8	2.0	2.4	3.1	3.8	4.2	5.3
Knowledge-based	26.3	30.9	35.9	38.6	42.4	46.0	46.3	48.2

Note: KIBS industries are financial and insurance, information and communication, business services, and education. Knowledge-based is the sum of tech-intensive, KIBS and health, social work values.

Source: Author's calculations from Korea Statistics, National Accounts, GDP by Economic Activities.

1950 to 35 percent in 2010, while the shares of manufacturing and non-KIBS service remained basically stagnant, hovering between 10 to 15 percent. Her definition of KIBS is very broad and includes finance and insurance, real estate, business services, and professional services.[6]

In order to review a corresponding trend in Korea's context, I need a definition of knowledge-based industries. The OECD STI (Science, Technology and Innovation) directorate regularly updates the list of technology-intensive manufacturing and knowledge -intensive service industries in its STI scoreboard publication.[7] I classify high-tech and medium-high-tech industries as technology-intensive manufacturing industries.

The series of long-run value added by industry in Korea's National Account System is available from 1970 in 13 manufacturing and 21 service industries.[8] I regard the following as tech-intensive manufacturing: chemicals (including pharmaceuticals); computers; electronic and optical instruments; electrical equipment, machinery and equipment; and transport equipment (motor vehicle and other transport). With regard to knowledge-intensive services, I select financial and insurance, information and communication, business services, and education. The health industry is clearly knowledge intensive but it is bound together with social work in the output statistics. Therefore, I separate out this industry.

Table 2 shows the shares of technology- and knowledge-intensive industries out of total value added in five-year intervals between 1985 and 2020. The first row is the share of manufacturing, and it remains between 27 and 30 percent throughout in the past 35 years. However, within manufacturing, the share of technology-intensive industries consistently rises up to 2010.[9] This trend is graphically shown in Figure 3 over a longer horizon. This rise explains the gap between the constant output share and the falling employment share of manufacturing. The share of KIBS rises faster

[6]Public administration is, as a rule, not included as a knowledge-intensive service.

[7]STI classifies technology-intensive manufacturing industries according to the R&D intensity. In ISIC3, aircraft and spacecraft (353); pharmaceuticals (2423); computers and electronic instruments (30); communication and broadcasting equipment (32); and medical, precision, and optical instruments (33) have the highest R&D intensity and are classified as high-tech (HT) industries. The next highest R&D intensity groups are motor vehicles, trailers and semitrailers (34); chemicals and chemical products, excluding pharmaceuticals (24 excl. 2423); electrical machinery and apparatuses (31); other machinery and equipment (29); and railroad equipment and other transport equipment (352+359). These are classified as medium-high-tech (MHT) industries (OECD, 2001, p.124, p.139).

[8]In KSIC rev.10 industrial classification.

[9] The share of tech-intensive (HT and MHT) manufacturing workers in all manufacturing has been approximately 15% continuously since the mid-1990s to the present.

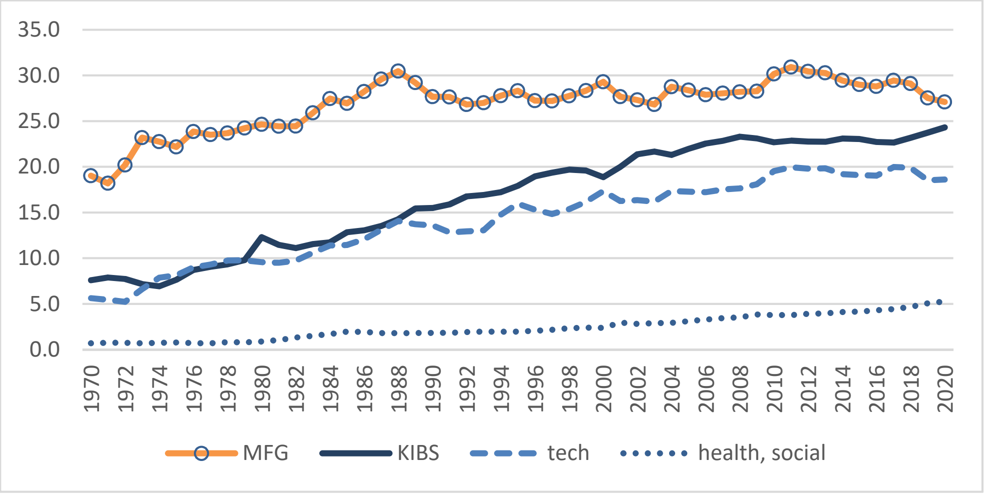

FIGURE 3. SERVICE AND MANUFACTURING EMPLOYMENT: 1970-2020

Note and Source: Identical to that in Table 2.

than the share of tech-intensive manufacturing. Starting from a similar level, the share of KIBS stands above that of tech-intensive industries in 2010 as a consequence of its quicker ascent in the 1990s and 2000s. The increase would have been quicker if medical services were added to it. Again, it flattens from 2010 onwards. When tech-intensive industries, KIBS, and health and social work are considered as knowledge-based industries, the sum of their shares reaches 48.2 percent as of 2020, slightly below one half (row 4).

B. *Geographical Concentration of Job Creation*

Another important evidence indicating a shift in the job creation pattern is the geographical concentration of job creation areas. Figure 4 compares the locations of the top 10 job creation Si-Gun-Gu areas in the decade from 1995 to 2005 with those in the decade from 2006 to 2016. The sizes of the circles are proportional to the numbers of jobs created. In the previous decade, the areas were either new residential areas or manufacturing centers located along the Seoul-Pusan line (See Column 3 in Table 3). In the latter decade, all are within or near the Seoul metropolitan area. Some are new residential areas but the majority are knowledge-intensive service industry centers. Clearly, job creation has gravitated towards Seoul, especially to the Gangnam area, which is the new and prosperous part of Seoul. Among the three Gu's in the Gangnam area, one is ranked at the second and another is at the fourth place. The Gu at the fifth is bordering with the Gangnam.

Table 3 lists the names of the top 10 areas in Figure 3. The administrative district units are Si-Gun-Gu. A 'Gu' is a district within a large city. It corresponds to, for example, a borough in the U.S. A 'Si' is a small to medium-sized city which does not belong to a metropolitan area. A 'Gun' is a district in less urbanized areas. There are approximately 250 Si-Gun-Gu areas in Korea with an average population size of around 200 thousand.

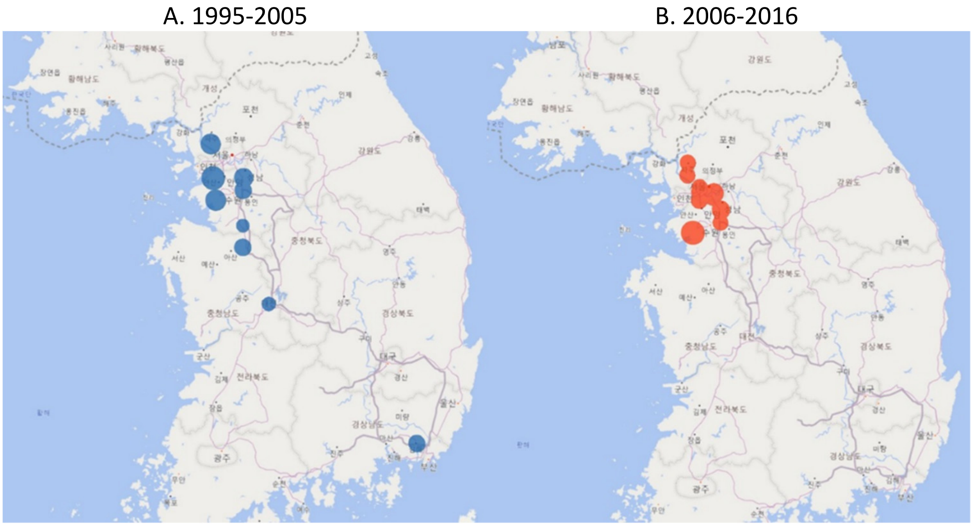

FIGURE 4. TOP 10 JOB CREATION AREAS: 1995-2005 VS 2006-2016

Note: The sizes of circles are proportional to the number of jobs created.

Source: Author's calculations from the Korea Statistics, the Census on Establishments, micro datasets.

TABLE 3—TOP 10 JOB CREATION SI-GUN-GU AREAS: 1995-2005 VS 2006-2016

Ranking	1995-2005 Si-gun-gu	Job Growth (thousands)	Mfg Share* (1995, %)	2006-2016 Si-gun-gu	Job growth (thousands)	Mfg Share* (2006, %)
1	Gyeonggi Hwaseong	108.5	72.5	Gyeonggi Hwaseong-si	179.9	62.5
2	Gyeonggi Goyang-si Ilsan-gu	103.4	21.8**	Seoul Gangnam-gu	121.3	6.4
3	Gyeonggi Seongnam Bundang-gu	92.5	15.2	Gyeonggi Seongnami Bundang-gu	108.1	6.0
4	Gyeonggi Siheung-si	86.3	68.9	Seoul Seocho-gu	85.2	7.1
5	Gyeonggi Yongin-si	76.1	51.4	Seoul Geumcheon-gu	84.9	33.9
6	Chungnam Cheonan-si	74.8	38.4	Seoul Yeongdeungpo-gu	82.4	10.7
7	Gyeongnam Gimhae-si	68.2	48.5	Gyeonggi Paju-si	81.2	43.0
8	Gyeonggi Suwon-si Paldal-gu	51.8	42.4	Gyeonggi Goyang-si, Ilsan-gu	80.4	12.9**
9	Daejeon Seo-gu	49.7	8.0	Gyeonggi Yongin-si	78.8	26.6
10	Gyeonggi Pyeongtaek-si	45.8	40.7	Seoul Mapo-gu	77.4	6.9

Note: 1) * Percent of workers in manufacturing in all industries exclusive of agriculture, fishery, quarrying, and mining, 2) ** Average of Gyeonggi Goyang-si.

Source: Author's calculations from the Korea Statistics, the Census on Establishments, micro datasets.

A detailed review of the new jobs in the Gangnam area by industry reveals that the majority are professional service jobs such as business headquarters, legal service, and programming jobs (See Table 4). These jobs are knowledge intensive and have agglomeration effects. Hence, the geographical concentration is interpreted as a consequence of the transition to a knowledge-based economy. Knowledge activities in highly complicated industries are very specialized such that cooperation becomes important and participators congregate in specific areas (Balland *et al*., 2008). As these effects are strong at very short distances, such as within miles, knowledge industries tend to be highly concentrated geographically (Rosenthal and Strange, 2004).

C. *Geographical Concentration of Job Creation*

Knowledge-intensive services mean that the service industry requires a high level of knowledge activity for production or delivery of the services. The OECD classifies knowledge-intensive services according to the share of the highly educated among the workers in the industry. The OECD classifies this at the two-digit level. In ISIC3, communication (64), financial and insurance (65, 66, 67), and business support (71-74, excluding real estate) belong in this category. Education (80) and health (85) are intermediate, and they are usually categorized separately.[10]

The NSB (National Science Board, 2018) uses a more detailed classification and excludes some service industries that do not require a high level of knowledge activity. For example, it excludes social work and the renting of household goods (713, in ISIC3 henceforth). This classification takes into account the level of R&D intensity in services and includes the publishing of recorded media (221) and news agency activity (922) as knowledge-intensive services. Since the OECD classifies this at a two-digit level and both health and social work are in ISIC3 85 industries, social work is classified as knowledge intensive as health is an important knowledge-intensive service. However, social work is not treated as knowledge-intensive when a lower level of classification is used, as done by the NSB.

The countrywide version of the Census on Establishments datasets contains industrial classification codes up to five digits.[11] I select five-digit industries corresponding to ISIC3 classification 71 (renting of machinery and equipment, household goods), 72 (information, computers), 73 (R&D), 74 (professional, scientific, and technical services), 65-67 (financial, insurance), 642 (communication), 80 (education), and 851(medical), and exclude rentals of consumer goods (713) and include the publishing of recorded media (221) and news agency activity (922). This result is shown in Figure 5.

The share of knowledge-intensive service workers among all industry workers rises robustly from 1995 to the end of the 2000s and then levels off. The rise is sharpest in the 2000s. Industrial classification codes were revised between 2005 and 2006, but a break is not apparent.

The trends shown in Figures 1 to 5 are all interrelated. The Korean economy's transition to a knowledge economy was rapid in the 2000s, and both technology-intensive manufacturing and knowledge-intensive services grew rapidly. Knowledge-intensive services played an important role in job creation, and jobs became concentrated within or near the Seoul metropolitan area. As the pace of the transition decelerated, job creation weakened. The next section focuses on the spillover effects of the knowledge sector on non-tradable, local service jobs.

[10]See OECD (2001), Table D.5.1, p.203.

[11]Statistics Korea deletes all establishments in an industry in the area if there are two or fewer establishments in the industry within the area when providing the Census on Establishments datasets. Hence more observations are lost when more detailed information about areas is requested. The countrywide version contains minimal information about areas but complete information about establishments.

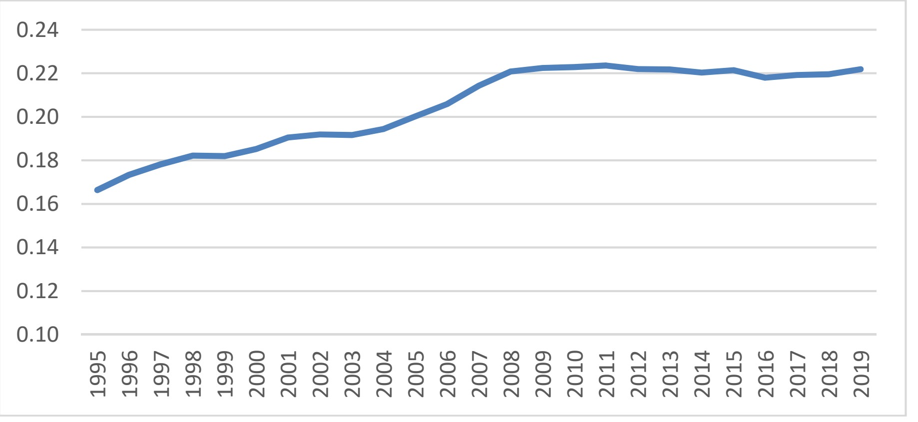

FIGURE 5. SHARE OF KNOWLEDGE-INTENSIVE SERVICE WORKERS AMONG ALL INDUSTRY WORKERS

Source: Author's calculations from Statistics Korea, Census on Establishment micro datasets.

III. Local Labor Market Job Creation Model

This section establishes a local labor market model to estimate the job-creation effects of a knowledge economy. An important feature of a knowledge economy is its geographical concentration. By comparing localities where knowledge industries are concentrated with those in which they are not, a meaningful inference about the job-creation effects can be made. This is one of the reasons why a local labor market model is particularly useful when analyzing a knowledge economy. Another advantage is that one can disregard institutional effects, which often complicate a labor market analysis, as localities within a country have largely the same institutions.

On the other hand, an international comparison or an input-output analysis is not appropriate in this context because between countries there are wide differences in institutions whereas their stages of transition towards a knowledge economy are similar. An I-O model presumes a national economy and quantifies transactions between industries. On a national scale, one cannot identify the effects of a knowledge sector on local service job creation. Moreover, definitions of knowledge or tradable industries are not so clear-cut such that detailed industry information is needed, but I-O sectors are defined broadly for data reasons. General equilibrium aspects such as rising housing prices at knowledge centers cannot be addressed either.[12]

This section is organized as follows: Subsection A builds a local labor market model, subsection B discusses how to determine local labor markets, subsection C defines tradable services, and finally subsection D compares tradable services with knowledge-intensive services.

A. *Model Specification*

I set up a simple local labor market model to estimate job creation in a knowledge economy. A local labor market model presumes two sectors within an economy—

[12]See Moretti (2010; 2011) for a more in-depth explanation.

the tradable sector and the non-tradable sector, which is equivalent to the local service sector. The sectors are divided based upon industry characteristics. The tradable sector does not exactly overlap with knowledge-intensive industries, but as the tradable sector growth is driven by knowledge industries, the effects of its growth virtually reflect the effects of knowledge-sector growth.

Goods and services in the tradable sector are produced to meet outside demands, while services in the non-tradable sector are produced locally to meet local demands. Typical tradable sector industries are manufacturing and agriculture, where goods are traded to meet outside demands. There are service industries in the tradable sector as well, and they are called 'tradable services.' They are usually professional and knowledge-intensive services, such as broadcasting, universities, medical hospitals, and legal services, among others. The industries in the non-tradable sector are local services by nature, such as restaurants, lodging establishments, retail shops, laundry and cleaning businesses, construction companies, medical clinics, and primary schools. Local services are mostly not highly knowledge intensive, but there are exceptions, such as clinics and schools. However, they do not constitute the main growing part in a knowledge economy.

The model I use is a modified version of Moretti (2010) with the results of Van Dijk (2015; 2018) taken into consideration. The basic model is as follows:

$$(1) \qquad \frac{\Delta L_{ct}^{NT}}{L_{ct}} = \alpha + \beta_1 \frac{\Delta L_{ct}^{T1}}{L_{ct}} + \beta_2 \frac{\Delta L_{ct}^{T2}}{L_{ct}} + C_{ct}\gamma + \varepsilon_{ct}$$

In equation (1), the superscript NT denotes the non-tradable sector. Unlike in Moretti (2010), the tradable sector consists of two parts: manufacturing (T_1) and tradable services (T_2). Figure 4 and Table 3 provide a hint of the rising importance of tradable services as a driver of job creation, and I choose to separate the two. Primary industries (agriculture, fishery, forestry, and mining) are dropped from the sample throughout.

The subscript c denotes city, which corresponds to a 'Si-Gun-Gu' in this study, and t is time, which is a ten-year period. The variable L represents the number of jobs in the sector. The growth rates of L are defined using the initial period total, L_{ct} $(= L_{ct}^{NT} + L_{ct}^{T1} + L_{ct}^{T2})$ as denominators. The reason I use the total instead of its own initial period value is to control excessive volatility.[13] Local labor markets are small and the growth rates tend to vary widely when their own initial values are used.[14] For example, when a plant moves out from inside a city to its vicinity, the city's manufacturing employment growth rate then shows a large negative number while that of the affected vicinity becomes very large, magnifying the estimation error. The initial period total suppresses such large variations. Van Dijk (2015) performs a sensitivity test and recommends the use of linear differences over the total as growth

[13]Moretti (2010)'s model uses log differences, i.e., $\ln\left(\frac{L_{ct}^T}{L_{cs}^T}\right), s < t$ instead. This term takes a very large positive or negative value if L_s is very small.

[14]An auto plant was set up in 1997 in Gangseo-gu, Pusan, resulting in an increase of manufacturing employment from 7 to 40 thousand in 1995-2005. Hwaseong is a new city in a manufacturing area. The number of its manufacturing workers jumps from 50 to 120 and to 200 thousand in 1995, 2005, and 2016, respectively.

rates. The variable C_{ct} is a set control variables for area characteristics. I use the shares of those educated at a four-year university or more among the population aged between 25 and 64 in the Population and Housing Census data. Dummies for metropolitan areas are not used as they are statistically insignificant. A full fixed-effect model is not used, as in this case I lose one of the two periods.

The parameters β_1 and β_2 are the central parameters and measure how many non-tradable sector jobs an additional tradable sector job creates. For this reason, Moretti (2010) calls them 'local multipliers.'[15] These parameters reflect general equilibrium effects, which should be taken into account when interpreting estimates.[16] For example, if there is a productivity shock in an area, it pushes up housing costs and reduces local service job-creation effects.

Another issue is whether to use weights in the estimations. I use the initial period total employment sizes L_{ct} as weights. Moretti (2010) uses total employment sizes in 1990 as weights both for the 1980-1990 and 1990-2000 periods. Van Dijk (2018) recommends OLS instead. He explains that WLS and OLS estimates have different interpretations. OLS estimates the job-creation effect in an average area, whereas WLS estimates the effect of an average job. In the Census on Establishments sample, the sizes of Si-Gun-Gu areas are very diverse and metropolitan Gu areas are much larger. The WLS estimation is driven by large areas, whereas the OLS results are determined by many small areas. As tradable services are concentrated in large areas, and WLS produces larger R^2 values, I choose to use WLS. As for the weights, population sizes can be alternatives, but as populations must be retrieved from the Census dataset[17] and they exist in five-year intervals, I choose to use total worker sizes. Experimentation with population sizes as weights gives just slightly different results.

B. *Areas of Local Labor Markets*

To implement the model, local labor markets must be determined among others. An obvious choice is a city. A city is a geographical unit composed of business centers and residential areas. Workers are much more mobile within a city than between cities. The difficulty of commuting produces housing price differentials across cities, which is a major reason of wage differentials between cities in a local labor market model.

The residential area of workers in a city is called a commuting zone (CZ), and a local labor market is commonly identified with a CZ. Commuting zones are determined from an travel-to-work pattern analysis in census data. Tolbert and Sizer (1996) selected 741 CZs in the U.S. using the U.S. Census data, which are commonly used in U.S. studies. For Korea, Yoon *et al*. (2012) at the Korea Labor Institute adopt the same method and construct 130 local labor markets using a 10% subsample of

[15]In economics, a multiplier is defined to include its own effect such that a fiscal multiplier becomes one if the output is equal to the input. However Moretti's multiplier does not count its own effects. Van Dijk (2015; 2018) points this out and claims that true multiplier value should be 1 plus Moretti's local multiplier.

[16]See Moretti (2011) for an explanation of general equilibrium comparative statistics.

[17]The use of the census dataset necessitates another round of administrative district adjustments, which can never be done perfectly.

the Korea population and housing census data. They use Coombe's algorithm, which determines zones from workers' travel-to-work flows. In their construction, the entire Seoul metropolitan area plus some adjacent cities constitute a single local labor market, as the residences of workers in Seoul are very widespread geographically.

A problem with this construction is that the Seoul labor market covers a very large part of the Korean economy—one-fifth of the total population and a quarter of total production—which makes inference based on comparisons of local labor markets less meaningful. Furthermore, job concentration in the 2000s is towards the Gangnam area, which is located in the southern part within the Seoul metropolitan area, with a population size less than one-fifth that of Seoul.[18] When the entire Seoul metropolitan area is regarded as a single labor market, job concentration within Seoul cannot be analyzed. Cermeño (2018) uses counties in the U.S. to analyze a century-long pattern of localization of job creation in the U.S. There were 3,143 counties in the U.S. as of 2010, and a county is a much smaller unit than a commuting zone. Cermeño finds that increasing returns-to-scale effects of a knowledge economy are a major cause of the job concentration effect in the 2000s in the U.S. As the increasing returns-to-scale effect is very sensitive to distance,[19] she uses a smaller unit of areas.

As an alternative, I use a 'Si-Gun-Gu' as a local labor market. A 'Si-Gun-Gu' is the smallest administrative entity with local administration in Korea. A 'Si' is a small to medium-sized city that is not part of a metropolitan area. A 'Gun' is a district in a rural area. A 'Gu' is a district in a metropolitan city. There are six metropolitan cities, and Seoul has 25 Gu areas. As of the end of 2017, there were 260 Si-Gun-Gu in Korea.[20] These consist of 75 Si, 82 Gun, and 69 Gu with local administrations and two other non-autonomous Si and 32 Gun. The Census on Establishment samples have 245 Si-Gun-Gu areas in 2000, 250 in 2005, 251 in 2010, and 250 in 2016. I use a version of the Census on Establishments data that has information on Si-Gun-Gu as well as Eup-Myon-Dong, which is a lower unit, and industrial classification codes up to three digits. Administrative districts have been reorganized several times, and I match the districts in the 1995 and 2005 and the 2006 and 2016 samples using Eup-Myon-Dong information and each Si-Gun-Gu's history on the Namuwiki websites.[21] I drop very small districts so that the sample is composed of 237 Si-Gun-Gu areas. Many metropolitan Gu areas are adjacent, and job creation spillover effects may occur in other Gu areas. This can diminish spillover effect estimates, but I use the locations of establishments and not residences, and more importantly, if the entire Seoul area is a single labor market, I cannot address job concentration in a knowledge economy properly.

[18]As of 2000, the population size of the Gangnam 3 Gu area is 1.7 million, 17% of Seoul's total population size of 9.9 million.

[19]The knowledge economy increasing returns-to-scale effect of is known to be very sensitive to distance. It drops to a fifth in five miles and disappears at more than 10 miles (Rosenthal and Strange, 2003).

[20]According to the Ministry of Public Administration and Security (2018).

[21]http://namu.wiki

C. *Tradable Services*

In equation (1), the tradable sector consists of manufacturing and tradable services. Tradable services constitute an important part given that the spillover effects are large. Job concentration in large cities cannot be explained without tradable services. By definition, the tradable sector is composed of industries that supply outside areas, but as transactions are not recorded, a different way of identifying tradable industries is needed.

Professional services such as broadcasting, advertising, universities, and hospitals are doubtlessly tradable in nature. Primary and secondary schools as well as local medical clinics are knowledge-intensive local services. Industries in financial services, communication, health, and education retain both tradable and non-tradable characteristics, and a clear-cut borderline is difficult to draw. Bank branches meet local demands but their headquarters function as a tradable service.

Jensen and Kletzer (2005) propose a simple way of identifying a tradable service. By definition, the sizes of local services are likely to be proportional to the area sizes in terms of population or employment. For tradable services, this is not necessarily the case. Thus, if an industry is geographically concentrated and its distribution is unrelated to the sizes of the areas, it can be regarded as a tradable industry. The measures of geographical concentration can be diverse, and Jensen and Kletzer (2005) use a locational Gini index. The index measures how unequally the industry workers are distributed compared with distribution of the population or employment. When the industry's worker distribution is perfectly proportional to the population or employment distribution, the industry has a Gini value of zero, whereas when all workers in the industry are concentrated in a single area, its Gini value is one.

The version of the Census on Establishments dataset I use to estimate the local labor market model has detailed location information up to the Eup-Myon-Dong level but only three-digit industry codes. The three-digit classification information is adequate for dividing industries into tradable and non-tradable types. It distinguishes primary, middle, high schools, and colleges. Departments, supermarkets, and convenience stores are different industries. Use of very fine five-digit classification does not greatly improve the quality of the division. For example, bank branches and headquarters are in the same industry even at the five-digit level. One complication is that the Eup-Myon-Dong sample does not contain full observations, as Statistics Korea drops observations if there are two or fewer establishments in the same industry in the same area for confidentiality reasons. This omission impairs the accuracy of Gini indices. For this reason, I calculate the industry Gini indices from older version datasets that were produced before this omission rule has been implemented. I use the 2005 sample for KSIC8 and the 2010 sample for the KSIC9 classification.[22]

Regarding the choice of a threshold Gini index value, Jensen and Kletzer (2005) simply suggest guidelines. They report that Gini values of 0.1 and 0.3 divide the sample roughly into thirds. They recommend the choice of a threshold value that sets all manufacturing as tradable, construction as non-tradable, and the majority of

[22]As shown in Table 7, the 1990s samples show a very different distribution across industries from the samples in the 2000s. Thus, I choose the most recent (2005) one from among the samples in the KSIC8 codes.

TABLE 4—SHARES OF TRADABLE AND NON-TRADABLE SECTOR WORKERS

	1995	2000	2005	2006	2010	2016
G < 0.1	.324	.357	.332	.352	.322	.341
G < 0.3	.485	.583	.582	.578	.577	.588
≤ Water supply	.832	.684	.645	.629	.623	.651
Country	1995	2000	2005	2006	2010	2016
manufacturing	.272	.245	.228	.217	.194	.189
Tradable services	.094	.102	.110	.128	.144	.154
Local service	.627	.648	.659	.652	.659	.658
Seoul	1995	2000	2005	2006	2010	2016
manufacturing	.188	.159	.123	.100	.061	.055
Tradable services	.174	.198	.210	.248	.282	.291
Local service	.636	.642	.666	.658	.661	.654
Seoul, Gangnam*	1995	2000	2005	2006	2010	2016
manufacturing	.114	.101	.088	.066	.025	.017
Tradable services	.221	.291	.265	.307	.317	.349
Local service	.660	.606	.646	.626	.658	.651

Note: 1) Shares in all industries excluding primary industries. See text for definitions of sectors, 2) *Average of the three Gu areas (Seocho-gu, Gangnam-gu, and Songpa-gu).

Source: Statistics Korea, author's calculations from Census on Establishment micro datasets.

public utilities as non-tradable.

The first and second rows in Table 4 show the shares of non-tradable sector workers for the threshold values 0.1 and 0.3 in the respective years. As in Jensen and Kletzer (2005), each constitutes approximately a third of the total. Typical local services such as haircuts, retail stores, restaurants, primary and middle schools, post offices, police and fire stations, and local administrative offices have Gini values of less than 0.1. Local medical clinics are usually located close to this threshold value. When the value is set to 0.3, the share of industries below the threshold is slightly less than 0.6. Jensen and Kletzer (2005) recommend the use of a public utility industry as a borderline industry. If I use water supply industry (360 in KSIC9) as the borderline,[23] the shares of local service industries are approximately two-thirds, as given in row 3. In this classification, some public utilities, wholesale companies, long-distance transportation companies, certain financial and insurance services, colleges and special high schools, and movie and broadcasting companies are in the tradable sector.

In classifying industries into tradable and non-tradable types, I consider the industry characteristics in addition to the Gini value. Industries in the sections of business support (N), real estate (L), arts, sports, leisure (R), and associations and membership organizations (S) are local services by nature. Some of them are distributed very unevenly, such as labor unions (942), business support services (751), worker dispatch agencies (759) and real estate activity agencies (681). These industries meet local demands but because their customers are distributed unevenly, they are distributed unevenly. For example, labor unions are close to factories but

[23]Another candidate is electricity (351 in KSIC9). The Gini values in the 2010 samples are 0.392 for the water supply and 0.548 for electricity. I choose the water supply as its Gini value is more proper for the purposes here.

factories are concentrated. Real estate activity agencies are concentrated where real estate transactions are active. These industries are classified as local services as jobs in the industries are created by spillover effects of other tradable industries. Some specialized services such as travel agencies (752) and security companies (753) are located at the centers of large cities. I classify them as tradable services.

The rows of row 4 and below show the sector's shares by region. Countrywide, the shares of tradable services show a rising trend. This increase is strong in Seoul, and it is strongest in the Gangnam area. The trend for the non-Seoul areas rises very modestly, which is not shown here, but as can be inferred. The gaps in the trends explain job creation disparity between the Seoul metro area and others. Local services' shares do not show any clear trends in all cases. The increase in the shares of tradable services is matched by the decline of manufacturing's shares.

D. *Tradable Services and Knowledge-intensive Services*

Tradable and knowledge-intensive services have different definitions and have different component industries, but their growth characteristics are common in a knowledge economy such that the effects of the tradable service sector growth reflects the effects of growth the knowledge sector. While a tradable service industry can be selected empirically, the definition of a knowledge-intensive industry is less clear-cut. This subsection reviews their relationship and show that tradable service growth is largely due to knowledge-intensive service growth.

Table 5 shows the composition of tradable and knowledge-intensive industries in the service industry for the years 1995, 2005, 2006, and 2016 and the corresponding changes. In column (4), the share of knowledge-intensive industries in the services rises from 0.21 in 1995 to 0.25 in 2016 (in rows 3, 6, 12, and 15).[24] Within tradable services, its share rises much more quickly from 0.34 in 1995 to 0.59 in 2016 (rows 1, 4, 10, and 13).[25] Such trend implies that the growing part of tradable services is likely to be the knowledge-intensive part.

Rows 7 to 9 and 16 to 18 show the changes in the ten-year periods. Between 1995 and 2005, tradable services increase by 322 thousand. Among them, the knowledge-intensive component accounts for 280 thousand, i.e., 87%. Between 2006 and 2016, tradable services increase by 1,204 thousand, with 819 thousand or 68% stemming from knowledge-intensive services. In addition to aggregate numbers, the high-tradable-service-growth areas are at the same time shown to be high-knowledge-industry-growth areas, as in the Gu areas in the Gangnam area. In the estimation, I use changes in the numbers of workers in the tradable and local service sectors and interpret the effects of tradable service growth practically as the effects of knowledge-intensive service growth.

[24]The shares of tradable services in service industry, which is the column ratios of column (3), is not show in Table 5. They are 0.12 (1995), 0.13 (2005), 0.15 (2006), and 0.18 (2016).

[25]There are many industries geographically concentrated but do not require high knowledge activity. Examples are wholesale, long-distance transportation, storage and public utilities such as gas and electricity.

TABLE 5—DISTRIBUTION OF TRADABLE SERVICES AND KNOWLEDGE-INTENSIVE SERVICE

		Workers (thousands)			Row ratios	
		(1) Knowledge	(2) non-KI	(3) Service total	(4) Knowledge	(5) non-KI
1995	Tradable	357	686	1,043	0.34	0.66
	Local Service	1,445	6,125	7,570	0.19	0.81
	Total	1,801	6,812	8,613	0.21	0.79
2005	Tradable	637	728	1,365	0.47	0.53
	Local Service	1,909	7,277	9,186	0.21	0.79
	Total	2,546	8,004	10,550	0.24	0.76
Change in 1995-2005	Tradable	280	41	322	0.87	0.13
	Local Service	465	1,151	1,616	0.29	0.71
	Total	745	1,193	1,937	0.38	0.62
2006	Tradable	841	751	1,592	0.53	0.47
	Local Service	1,666	7,335	9,001	0.19	0.81
	Total	2,507	8,086	10,593	0.24	0.76
2016	Tradable	1,660	1,136	2,796	0.59	0.41
	Local Service	2,214	10,450	12,664	0.17	0.83
	Total	3,874	11,586	15,460	0.25	0.75
Change in 2006-2016	Tradable	819	385	1,204	0.68	0.32
	Local Service	548	3,115	3,663	0.15	0.85
	Total	1,801	6,812	8,613	0.21	0.79

Note: See the text for definitions.

Source: Author's calculations from Statistics Korea, the Census on Establishments micro datasets.

IV. Empirical Estimates

Section IV reports and discusses the estimation results. In a local labor market model, tradable sector jobs create non-tradable sector jobs through the spillover effect. Actually, much more jobs are in the non-tradable sector than in the tradable sector. An estimation of the spillover effects will show how the job-creation mechanism has changed as the tradable sector grew and became more knowledge intensive.

I analyze the change in the 1995 to 2005 period (in KSIC8) and in the 2006-16 period (in KSIC9) from the Eup-Myon-Dong level and three-digit version of the Census on Establishments samples. A ten-year period is a natural choice as spillover effects take a very long time to fully materialize.[26] I drop agriculture, fishery, forestry, and mining from the sample. The weight variables are the number of workers in all industries (excluding primary industries) at the initial periods, in 1995 and in 2006.

The sample consists of 237 Si-Gun-Gu countrywide. There are 25 Gu in Seoul, 43 Gu in five other metropolitan cities, and 169 Si-Gun, which are composed of basically small to medium-sized cities. Table 6 shows the sample summary statistics.

[26]Five-year intervals produce very small and often statistically insignificant estimates. Furthermore, two major economic crises in the intervening years, the Asian Financial Crisis of 1998 and the Global Financial Crisis of 2008, make five-year intervals practically unusable.

TABLE 6—SAMPLE SUMMARY STATISTICS: AVERAGES

Averages of	237 Si-Gun-Gu			25 Seoul Gu		
Periods	(1) 1995-2005	(2) 2006-2016	(3) 1995-2016‡	(4) 1995-2005	(5) 2006-2016	(6) 1995-2016‡
Manufacturing	-1,088	486	-301	-10,217	-5,816	-8,017
Tradable services	1,623	3,413	2,518	5,384	17,275	11,329
Local service	6,029	11,313	8,671	3,930	21,370	12,650
Initial L size*	57,069	64,861	60,965	154,570	155,742	155,156
Share of univ grads†	.234	.373	.308	.327	.522	.425
Averages of	43 Gu in Other Metro Areas			169 Si-Gun		
Periods	1995-2005	2006-2016	1995-2016‡	1995-2005	2006-2016	1995-2016‡
Manufacturing	-4,775	15	-2,380	1,201	1,538	1,370
Tradable services	1,356	2,346	1,851	1,135	1,634	1,384
Local service	5,885	11,983	8,934	6,376	9,654	8,015
Initial L size*	76,182	80,400	78,291	37,783	47,464	42,623
Share of univ grads†	.217	.345	.283	.187	.314	.258

Note: 1) ∗ Workers in all industries exclusive of agriculture, forest, fishery, and mining, 2) †Among the population aged 25-64, 3) ‡Average of two ten-year periods.

Source: Author's calculations from Statistics Korea, the Census on Establishments micro datasets.

An average Gu in Seoul has 155 thousand workers, twice as large as an average Gu in other metro areas which is 89 thousand workers. These are again twice the size of an average Si-Gun, which has 43 thousand workers. Although there are much more Si-Gun areas in the sample, a Gu is much larger, consequently the totals of all Si-Gun and all Gu are approximately of the same size. The numbers of manufacturing workers increase in Si-Gun but decreases in metropolitan Gu areas. The tradable service worker increase is concentrated in the Seoul metro area. A Seoul Gu adds 11 thousand tradable service workers over a decade, while a Gu in other metros or a Si-Gun adds less than two thousand over a decade. Tradable service job creation is closely related to the shares of the highly educated. A Gu in Seoul has much more university graduates and experiences larger job increases in tradable services.

Countrywide, the size of local service job growth is 3.9 times that of tradable sector job growth (column 3). In Seoul, this ratio is 3.8 (column 6). In non-Seoul metro areas, the tradable sector job growth is negative and a ratio is not calculated. In Si-Gun areas, this ratio is 2.9 (column 6). The summary statistics suggest a larger job creation spillover effect of tradable sector jobs in Seoul than in Si-Gun areas, but a conclusion cannot be reached unless a regression analysis is conducted.

I start with a one tradable sector variable model with one β parameter. Single variable models are more common in the literature, and as such, the results can be more straightforwardly compared with existing results. A multi-collinearity problem between manufacturing and tradable services variables can also be avoided in a single variable case. [27] Table 7 reports the estimation results.

The sample consists of 237 Si-Gun-Gu areas with two decadal changes, meaning that there are 474 observations in total. When there is only one explanatory variable — tradable sector job growth — the coefficient estimate is 0.835 (column 1), which

[27] If I regress tradable services (X2) against manufacturing (X1) and education (EDU), I obtain, X2=-.23 + .065 X1 + .286 EDU (R^2=.250). The coefficient estimate of X1 has a s.d. of .023 and a p-value of .005.

TABLE 7—SINGLE TRADABLE SECTOR VARIABLE REGRESSION

	Specification			Tradable sector		Split samples		
	(1) All areas	(2) + education	(3) + period dummy	(4) Manufacturing	(5) Tradable Services	(6) Seoul	(7) Other metro areas	(8) Si-Gun
α	6,820	5,130	6,215	5,092	5,851	10,046	1,217†	1,206**
s.d.	590	629	869	693	673	2,851	2,729	578
β	.835	.835	.820	.449	1.122	.879	.623	.881
s.d.	.053	.051	.052	.071	.078	.128	.151	.049
EDU	-	.137	.134	.277	-.001	-.007†	.500	.674
s.d.	-	.022	.022	.024	.031	.039	.146	.043
D1995	-	-	-2,034**	-	-	-	-	-
s.d.	-	-	1,127	-	-	-	-	-
R^2	.344	.394	.398	.269	.305	.500	.257	.629

Note: 1) EDU is orthogonalized, 2) ** Significance at 5%, † No significance at 10%. All others are significant at 1%.

Source: Identical to that in Table 6.

means that one tradable sector job creates 0.8 local service jobs, with an R^2 value of 0.344. The summary statistics in Table 6 show that there are approximately four local service jobs are created for each new tradable sector job. However, the portion explained by only the quantitative changes of tradable sector jobs is much smaller. The R^2 value indicates that only a third of the total change can be explained by the equation. The coefficient estimate of the constant variable is as large as 6,820. As the average of local service job growth is 8,671, this means that much is left to be explained by variations not included in the equation. However, this coefficient estimate, 0.835, is not small. I return to this point later.

The education variable, which is the share of university graduates among the population aged between 25 and 64 in an area, is an important control variable for area characteristics, but this variable is correlated with the explanatory variable, especially with tradable services.[28] As the correlation is usually positive, the inclusion of the education variable lowers the estimates. I orthogonalize the education variable to eliminate this collinearity. That is, I regress education against the explanatory variable and subtract the correlated part from the variable. In this way I prevent the addition of the education variable from affecting the tradable sector variable coefficient estimate-s. When education is added (column 2), the R^2 value increases slightly to 0.394 and the constant estimate drops to 5,130. The education coefficient estimate is 0.137, which is very large. Suppose that the share of university graduates has risen by 0.1, then local service jobs increase by 1.37% of total employment. As the average size of total employment is 61 thousand (column 3, Table 6) 1.37% of this is 840 workers. On the other hand, the growth of the tradable sector added 1,885 workers. As the growth of the tradable sector is 2,257 (=2,518 – 301) workers on average and multiplied by the β estimate, 0.835, is 1,885. Thus, in this case education

[28]When education is regressed against the tradable sector, the estimated coefficient value is .654 (s.d.=0.107), and when it is regressed against tradable services, it is 2.338 (s.d.=0.116). Education is negatively correlated with manufacturing.

upgrading explains 45% of total local service job increase. The share of university graduates rose by 0.139 in the second decade relative to the first decade. When multiplied by the β coefficient estimate, 0.137, and the employment size in the second period, which is 65 thousand (column 2, row 4 in Table 6), the local service job creation effect is 1,235 workers. Local service job creation in the second period is larger than that in the first period by 5,284 (columns 1 and 2, row 3, Table 6). Educational upgrading accounts for 23% of the total change. The education variable proxies the quality of the tradable sector jobs. The result shows importance of the quality of the jobs in local service job creation.

Column (3) adds a dummy variable for the first period. The coefficient estimate is negative and statistically significant only at the 5% level, implying that the job-creation effect is smaller in the first period. This may be due to the smaller share of tradable services in the tradable sector, but the difference is not very significant. In columns (4) and (5), the dependent variable is the same, but only manufacturing or tradable services workers are included in the of tradable sector. The coefficient estimate for manufacturing is 0.449 in column (4) and that for tradable services is 1.122 in column (5), showing that a tradable service job has a larger local service job-creation effect. The education coefficient becomes very large at 0.277 in the case of manufacturing. Tradable services have a larger job-creation effect and education is correlated with this variable. The large coefficient value of education is a consequence of this relationship.

Columns (6), (7), and (8) are the results from subsamples with the same specifications. I split the sample into three: Seoul, other metro areas, and Si-Gun areas. Main differences between the samples are that Seoul has a much larger share of tradable services within the tradable sector and a higher ratio of the highly educated. As I have orthogonalized the education variable, its effect on the tradable service variable is controlled. The coefficient estimates are similar at 0.879 (Seoul, col 6), 0.623 (other metros, col 7), and 0.881 (Si-Gun areas, col 8). In other metro areas, the change in the tradable service is in the negative and the estimation is imprecise. The R-square values are correspondingly 0.500 and 0.629 in Seoul and Si-Gun areas but only 0.257 in other metro areas. The coefficient for education is small in the Seoul sample but very large in other metro areas and in the Si-Gun area samples. This occurs because the larger job creation effects of tradable services in Si-Gun areas appears in correlated education variables. This is confirmed by the results in Tables 9 and 10.

By period (in Table 8), the β coefficient estimate in the second decade at 1.0 is larger than that in the first decade which is 0.6 (in columns 1-2 and 3-4). The second period estimation produces a larger education coefficient estimate and a larger R-square value. The share of tradable service in the tradable sector is higher in the second period (See Figure 5), and this explains the larger β coefficient in the second decade. In addition, the first period has seen an economic crisis and consumption has been constrained. The larger estimates of β and the education efficient in column 4 both stem from the higher tradable service shares. Addition of the education variable in the second period estimation greatly improves the R-square value (in column 4), showing that education actually proxies the share of tradable services. Columns (5) and (6) are results from a stacked sample, where each

TABLE 8—SINGLE TRADABLE SECTOR VARIABLE REGRESSION, BY PERIOD

Period	1995-2005		2006-2016		1995-2016	
Column	(1)	(2)	(3)	(4)	(5)	(6)
α	5,705	5,752	7,371	4,686	13,275	10,026
s.d.	873	965	771	728	1,425	1,501
β	.605	.605	1.011	1.011	.917	.917
s.d.	.083	.083	.066	.057	.075	.072
EDU	-	-.004†	-	.227	-	.269
s.d.	-	.037	-	.025	-	.053
R^2	.185	.185	.498	.629	. 388	. 448

Note: 1) EDU is orthogonalized, 2) † No significance at 10%. All others are significant at 1%.

Source: Identical to that in Table 6.

observation corresponds to a two-decade long change. The estimation gives an intermediate value for the β parameter, at 0.917, and a large coefficient estimate for the education variable, supporting the importance of the qualitative aspect of the tradable sector. The larger second period estimates do not necessarily imply a structural shift in the second period. If I separately include manufacturing and tradable services in the tradable sector and add a period dummy, the dummy is not statistically significant, as shown in Table 10.

The estimate from the Census on Establishments sample is 0.835 (columns 1 and 2, Table 7), which is merely half of the multiplier estimate by Moretti (2010) of 1.6.[29] However, when differences in the samples and specifications are considered, the estimate is not entirely incompatible with Moretti's result. Van Dijk (2017, 2018) performs an extensive sensitivity test and claims that when the tradable sector variable is extended to include tradable services as well as manufacturing, the multiplier estimate from decadal changes lies in the range of 0.17 and 0.93 in the U.S. Census sample in the period between 1980 and 2000.[30] Three reasons can be put forward as to why my estimate is smaller than that of Moretti (2010). First, the Si-Gun-Gu areas are much smaller than the MSAs in the U.S.[31] When areas are small, some job spillover effects occur outside of the boundary, lowering the estimates. Second, when a broad definition of the tradable sector is used, the job creation multiplier value naturally falls. Manufacturing is the only tradable sector industry in Moretti (2010), but I include both manufacturing and tradable services in the tradable category.[32,33] Third, place of work data produce smaller multiplier

[29] Moretti (2010), Table 1, p.376. Moretti uses a single tradable sector variable specification, which is manufacturing. Control variables are not used, except for period dummies, resulting in larger estimates.

[30] Van Dijk (2018), p.281. Van Dijk suggests a range between 1.17 and 1.93, but this range includes its own effects, resulting in a value of 1.

[31] The samples of Moretti (2010) and Van Dijk (2017; 2018) consisted of 226 (1980) and 238 (1990, 200) MSAs in the U.S. Census data.

[32] Column 4 in Table 7 uses the same non-tradable sector definition used in columns 1 to 3. If all industries exclusive of manufacturing are regarded as non-tradable, the estimate is much larger. Van Dijk (2017) classifies a group of industries as 'medium-tradable.' When the tradable sector is extended to include this group, the estimate almost halves (p.476). The IV estimate drops from 1.69 to 0.72 (Table 2, p.475 and Table 4, p.477).

[33] Moretti (2010) uses log differences, whereas Van Dijk (2018) uses linear differences. The effect of this change on the size of estimates is, however, in the opposite direction. Van Dijk (2018) obtained a slightly larger estimate (1.60 vs. 1.99) in the linear difference specification (Table 5, p.291, column 2 and 7). The advantage of using linear

TABLE 9—TWO TRADABLE SECTOR VARIABLE MODEL: MANUFACTURING AND TRADABLE SERVICES

	All areas		Seoul & other metro areas		Si-Gun	
Column	(1)	(2)	(3)	(4)	(5)	(6)
α	5,836	5,313	7,223	7,928	4,942	1,000*
s.d.	581	624	1,404	1,483	622	593
β_1	.565	.565	.450	.450	.573	.573
s.d.	.064	.064	.144	.143	.073	.060
β_2	1.193	1.193	.952	.952	1.653	1.653
s.d.	.073	.072	.105	.105	.123	.100
EDU	-	.066**	-	-.056†	-	.727
s.d.	-	.029	-	.039	-	.056
R^2	.403	.410	.385	.395	.448	.632
No. of obs.	474	474	136	136	338	338

Note: 1) EDU is orthogonalized, 2) * Significance at 10%. ** Significance at 5%, † No significance at 10%. All others are significant at 1%.

Source: Identical to that in Table 6.

values than place of residence data do. The Census on Establishments data record places of work, while the Population Census records both places of work and places of residence. Van Dijk (2018) obtains a multiplier value of 1.60 from places of residence data and that of 1.49 from places of work data.[34]

Tables 9 and 10 are the results when there are two tradable sector variables: manufacturing and tradable service. Columns 1 and 2 in Table 9 apply basic specifications to the full sample. The manufacturing's coefficient estimate is approximately 0.6 and that of tradable services is close to 1.2. The education variable has a small value and is statistically significant only at the 5% significance level. A period dummy does not have statistical significance, and I do not use the dummy variable. The education variable is orthogonalized against explanatory variables, as before. Manufacturing and tradable service variables are correlated, but the multi-collinearity problem is not serious. If the dependent variable is regressed against manufacturing, the coefficient is 0.45, while the regression against tradable service produces an estimated value of 1.12. Therefore, I do not attempt to orthogonalize the variables.

Estimates vary greatly by areas when I split the sample (Table 9). The results from samples for Seoul, other metro areas, and Si-Gun areas are given correspondingly in columns 3 to 4 and 5 to 6 in Table 9. In the sample containing Seoul and other metro areas, the estimates are slightly smaller than those from the overall sample. Manufacturing's coefficient is 0.45 instead of 0.6 in the full sample, and the tradable services variable has an estimate of 0.95 instead of 1.2. Education is not statistically significant. The Si-Gun sample produces larger estimates: Manufacturing's coefficient is 0.57 and the coefficient for tradable services is as large as 1.65. When the education variable is added, its coefficient is very large and the R-square value jumps to 0.632 (in column 6). The Si-Gun sample has larger coefficients likely due

differences is to control volatility (p.290).

[34]Moretti (2010), p.291, Table 5.

TABLE 10—TWO TRADABLE SECTOR VARIABLES AND SWITCHING REGRESSION SCHEME

	WLS			2SLS		
Column	(1)	(2)	(3)	(4)	(5)	(6)
α	5,651	4,946	6,172	4,896	4,451	4,691
s.d.	574	617	846	617	647	904
β_1	.507	.507	.492	.512	.512	.509
s.d.	.065	.064	.064	.068	.068	.068
β_{21}	1.020	1.020	.997	1.109	1.109	1.102
s.d.	.084	.083	.084	.110	.109	.111
β_2	1.632	1.632	1.622	2.259	2.259	2.245
s.d.	.132	.131	.131	.182	.181	.185
EDU*	-	.086	.082	-	.089**	.090**
s.d.	-	.029	.029	-	.040	.040
D1995	-	-	-2,305**	-	-	-447†
s.d.	-	-	1,091	-	-	1,170
R^2	.422	.433	.438	.380	.386	.387

Note: 1) EDU is orthogonalized, 2) ** Significance at 5%. † No significance at 10%. All others are significant at 1%.

Source: Identical to that in Table 6.

to the differences in the definitions of the areas between the samples. In Seoul and other metro areas, the areas are Gu areas in the city, and they are located adjacent to others, while in the Si-Gun sample, the areas are distanced. Why education has a very large coefficient value in the Si-Gun sample is not clear. The variable is orthogonalized and the multi-collinearity effect is eliminated before the estimation. Compared to metropolitan areas, the Si-Gun areas have smaller tradable sectors and lower ratios of university graduates in the population. Moreover, there are wider differences among the Si-Gun areas than among the Gu areas in a metropolitan city. The large coefficient value for education in the Si-Gun sample appears to reflect the differences among the Si-Gun areas. Some of them are industrial centers with a high ratio of university graduates, while others are simply rural areas.

To incorporate the wide gap in the estimates of tradable services between the subsamples, I introduce a sort of switching regression scheme into the model and allow the tradable service variable to have different parameter values depending on the region. Suppose D is a dummy variable that takes a value of one if the observation is in the Si-Gun area and takes a value of 0 if it is in a metro area. I use two variables $(1-D)\Delta L_{ct}^{T2}$ and $D\Delta L_{ct}^{T2}$ in place of ΔL_{ct}^{T2}. An alternative is to add a Si-Gun dummy to the equation. However, in this case, interpretation of the dummy variable parameter is unclear because Gu and Si-Gun areas are different in many respects.

The estimation result is given in Table 10. Manufacturing has an estimated coefficient value of 0.5 and the coefficients for tradable services are 1.0 in metro areas and 1.6 in Si-Gun areas (columns 1 and 2). The period dummy is statistically significant only at the 5 percent level (column 3). Education's coefficient is small and positive, as expected. Columns 4 to 6 report the 2SLS estimation results. The instrumental variables used are the Bartik instruments, commonly used in local models. The instruments are constructed from three-digit industry-level changes and the shares of each Si-Gun-Gu in the industry in 1995 and 2006. Instrumental variables are used only for the tradable service variable. When predicted values are used both for manufacturing and tradable services, a multi-collinearity problem between the two arises again.[35] The first stage of the estimation is done independently in each subsample.[36] The first-stage estimation F-statistics are large enough at 858.9 and 620.8 in the metro area and Si-Gun subsamples, with R-square values of 0.65 and 0.57, respectively. The 2SLS estimates of manufacturing are approximately 0.5 and are only slightly larger than the WLS estimates. The 2SLS estimates for tradable services are 1.1 in Seoul and other metro areas and 2.3 in Si-Gun areas (columns 4 and 5). For Seoul and other metro areas, the 2SLS result is slightly greater than the WLS result, but for the Si-Gun areas it is significantly larger than the WLS value of 1.6. The 2SLS estimates can be either smaller or larger than the WLS estimates depending on the direction of correlation between tradable services and the second-stage error term, and both results can be found in the literature. However, when Bartik instruments are used, education variable coefficients become statistically significfsant only at the 5% confidence level. The period dummy is not statistically significant (column 6). I take the result in column 6 as the baseline result, as it controls the endogeneity of the variables as well.

V. Summary and Conclusion

According to the empirical results shown here, a manufacturing job creates 0.5 local service jobs, and a tradable service job creates 1.1 local service jobs within the same Gu in metro areas and 2.3 jobs in Si-Gun areas. Tradable service jobs have larger spillover effects than manufacturing jobs as they are likely to be higher wage jobs in which knowledge-intensive industries.

Many studies report even larger spillover effects. Moretti (2010) estimates that while an average job in the tradable sector creates 1.6 jobs, a job in the skilled[37] tradable sector creates 2.5 local service jobs.[38] Van Dijk (2018) obtains an estimate of 2.9 for a job in the skilled tradable sector with the same specifications.[39] For high-wage jobs, Van Dijk (2017) reports a multiplier of 3.7.[40] Moretti (2012) reports a multiplier value of 5 for innovation jobs.[41] These results indicate that spillover effects are larger when the qualitative aspects of tradable sector jobs are taken into account as well. As I use the Census on Establishments dataset, only the quantities

[35]Such sensitivity is found in other studies as well. For example, in Moretti (2010), Table 1 (p.376), when there are tradable durable and tradable nondurable variables, the tradable durable's coefficient estimate becomes statistically insignificant in the IV estimation. In Van Dijk (2017) in Table 2 (p.13), when two variables (tradable durables / nondurables) are used, the tradable durable's coefficient estimate varies considerably when the IV estimation is used. In Table 3 (p.15), when skilled tradable and unskilled tradable are used at the same time, all estimates lose statistical significance in the IV estimation.

[36]The first-stage estimates are 0.541 (s.d.=.018) and 1.31 (s.d.=.052) in the metros and Si-Gun samples. Tradable services shares are larger in metro areas than in Si-Gun areas, making the estimates less and greater than one.

[37]In the labor economics literature, 'skilled' usually means 'highly educated.'

[38]Moretti (2010), Table 2, p.377.

[39]Van Dijk (2018), Table 4, p.290.

[40]Van Dijk (2017), Table 6, p.479.

[41]Moretti defines innovation jobs as jobs that make intensive use of human capital and human ingenuity. Gene rally speaking, these jobs are high-skill jobs in innovation industries.

of jobs are considered in the estimation.

Tradable service jobs have larger job-creation effects, but many of them are created by the demand derived from manufacturing. The empirical estimates should not be interpreted simply to imply that service industry is replacing manufacturing in job creation. Usually, a tradable service job pays a higher wage than a manufacturing job. The larger job-creation effect is derived from the higher wage and higher level of consumption. We observe that when a manufacturing area is transformed into a service industry city, the number of local jobs increase. Advanced economies usually experience a decrease in manufacturing jobs, but they have more and better jobs as manufacturing is upgraded and more knowledge-intensive jobs are created, while the number of production workers diminishes. It is more important how the manufacturing industries are upgraded than how many workers they hire in an economy's job creation.

This research shows that as Korea transitions to a knowledge economy, job creation has become more active as knowledge-intensive service jobs grow and consequently the tradable sector expands. As the number of manufacturing jobs decreased, the total number of jobs did not decrease as knowledge-intensive jobs in the tradable sector grew and added local service jobs. The relationship between the tradable sector and local service job creation seems to have been stable throughout. The changes within the tradable sector, which are the decrease in manufacturing jobs and growth of tradable service jobs, largely explain the growth of local service jobs. Likewise, the slowdown of knowledge sector growth explains the slowdown in local job creation. The growth of knowledge-intensive services decelerated since the 2010s (Figure 5), and overall job creation has weakened (Figure 1).

The rise of the service-manufacturing jobs ratios, the geographical concentration of jobs, and the growth of knowledge-intensive services discussed in Section I are all typical phenomena in an economy's transition to a knowledge economy. A unique pattern in geographical concentration in Korea is that the gravitation is towards a single pole, the Seoul metropolitan area. The main message from Moretti's book (Moretti, 2012) is that the growth of certain jobs, which he calls innovation jobs, and where they are located are important factors for economic development. An implication of the book is that for regional development, a knowledge hub must be built within the region so that an agglomeration effect can work in the region. Relocating government agencies outside Seoul is not sufficient, and the regions should be able to provide amenities if they are to have a knowledge hub within them. As for the Seoul metropolitan, the Gangnam area has now become the knowledge center for the metropolitan as well as for the whole Korean economy, with an inevitable consequence for housing prices. Taxes and supply policies on housing need to be reviewed from a perspective of economic growth. When supply is limited or transactions are suppressed, returns from agglomeration effects accrue to housing owners instead of knowledge workers who contributed to the housing price rise. Such an objective seems to be missing in housing policy designs. At a recent KDI conference[42] Professor Dani Rodrik stressed that good jobs are created by good companies and that industrial regional policies and innovation policies are needed to

[42]National Economic Advisory Council and Korea Development Institute conference, "Towards a Virtuous Cycle of Growth-Jobs-Welfare," October 7, 2021, Four Seasons Hotel, Seoul, Korea.

foster good companies. Those are policies for sustained job creation in a knowledge economy as well.

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