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Crisis and Consumption Smoothing*

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We study how the 1997-1998 Asian Financial Crisis affected consumption smoothing across households in Korean prefectures. The crisis caused the cross-sectional mean and volatility of household consumption to fall substantially. We show that such falls bias two standard tests towards rejecting consumption risk sharing. Exploiting the different sizes of bias in these two tests, we find that full risk sharing during crisis at the prefecture level could not be rejected for a consumption measure that includes nondurable goods and some services. In addition, prefecture level full risk sharing before crisis cannot be rejected at conventional significance levels in at least twelve of all fourteen prefectures. National risk sharing is, however, rejected throughout the whole sample period.

Key Words: Consumption smoothing; Risk sharing; Relative risk aversion; Financial crisis; Generalized method of moments; Panel data.
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1. INTRODUCTION

Although previous research on consumption has examined data sets that cover several recessions, we are not aware of studies that formally test how economic downturns affect overall cross-sectional consumption smoothing,

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i.e. the sharing of idiosyncratic shocks to consumption among households.[1] The expansion of public or private transfers during a recession, along with market mechanisms of risk pooling, should help smooth out consumption risk. However, the effects of a downturn may be very unevenly distributed across individuals. And such uneven distributional effects are likely to lead to lesser cross-sectional consumption smoothing. An interesting question, then, is which of these two effects dominates. In this paper, we shed light on this question by comparing risk sharing before and during the 1997-1998 Asian Financial Crisis using the data from the Korean Household Panel Study (KHPS).[2]

A key econometric issue in such a study is that a major crisis may lead to substantial variations in the cross-sectional means and variances of household idiosyncratic consumption, income and asset variables.[3] Such variations may distort the standard test statistics in tests of consumption smoothing models. In this paper, we show that a large decline in the cross-sectional mean and volatility of household consumption biases two standard tests toward rejecting the benchmark of risk sharing. We exploit the difference in the bias size in these two statistics to determine whether our benchmark model is rejected for the crisis year.

A number of studies have rejected national full risk sharing [e.g. Asdrubali, Sorensen and Yosha (1996), Attanasio and Davis (1996), Hayashi, Altonji and Kotlikoff (1996)], though others report substantial amount of risk sharing among provinces/states of a nation [e.g. Crucini (1999)]. Furthermore, available partial risk sharing models have major counter-factual consumption implications [Krueger and Perri (2005)]. Therefore, we start with the plausible benchmark of prefecture-level full risk sharing, and use the deviation from it to measure the extent of cross-sectional consumption smoothing before and during the crisis.[4]

The econometric methods used in our paper are established in Ogaki and Zhang (2001, OZ henceforth) and extended in Zhang and Ogaki (2004, ZO) to test risk sharing hypothesis against the alternative of the permanent income hypothesis (PIH). In Section 3 below, we further improve these

[1]Attanasio and Browning (1995) presented evidence for intertemporal consumption smoothing over the life cycle and the business cycle. They did not consider cross-sectional smoothing.

[2]The KHPS survey discontinued in 1998.

[3]Idiosyncratic consumption (or income) can be defined as the deviation of household consumption (or income) in per adult-equivalent terms from the cross-sectional mean of household consumption (or income). With such a definition, the cross-sectional variance of idiosyncratic household consumption (or income) is the same as the cross-sectional variance of household consumption (or income).

[4]Our benchmark of prefecture-level full risk sharing can be viewed as a partial risk sharing "model" at the national level in that it allows inter-prefecture risk sharing through, for example, public or private transfers, but does not presume such sharing is complete.

methods not only by examining the effects of substantial variations in variable volatilities on test statistics, but also by more judiciously choosing instrumental variables. Furthermore, in addition to these methodological contributions, we reinforce the empirical findings in the two papers referenced above: the consumption data favors decreasing relative risk aversion (DRRA), and allowing for DRRA leads to non-rejection of full risk sharing at prefecture level (but not at national level). Since similar findings in OZ (2001) and ZO (2004) were obtained using data sets on rural households in developing countries, researchers have wondered if they will survive data from an industrialized economy.[5] It is encouraging to know that they do.

In the rest of this paper, we first describe in next section our data. We then present in Section 3 the consumption smoothing mechanisms used by the Korean households captured in the KHPS survey. In Section 4, we explore the effects on two standard test statistics of a large decline in cross-sectional volatility of household consumption during a crisis after we briefly describe our tests. We present our empirical results in Section 5, and conclude in Section 6.

2. DATA

The Korean financial crisis broke out in the last quarter of 1997, causing a sharp increase in unemployment rate from 2.6% to 8.7%, and a 6.9% drop in real GDP in a year.

The KHPS survey covered all Republic of Korea prefectures except Jeju-do, and was conducted during 1992-1998 by Daewoo Institute of Economic Research. However, the records for the first two sample years were incomplete. We therefore employ the last four rounds of the survey from 1994 to 1998, where each round covered the period from August of a year to July next year. In our tests, we include sample households with complete information on demographics, consumption, income, and asset for all the last four rounds. This criterion leaves us with 2,008 households in 14 prefectures, out of the 2,266 households that stayed through all six rounds of survey. The number of households in a prefecture in our sample ranges from 54 to 421.

The KHPS started with a representative sample of Korean households. Due to attrition and the removal of households with incomplete records, the sample we use in our empirical analysis cannot still be representative. Nonetheless, our sample quantities still track the national aggregates reasonably well. For example, while the U.S. Consumer Expenditure Survey captures slightly over 60% of the U.S. aggregate consumption, the KHPS consumption measure amounts to nearly 80% of the Korean national ac-

[5]Korea is a member of the Organization of Economic Cooperation and Development.

TABLE 1.

Consumption and Labor Income Per Adult-Equivalent in the KHPS and Korean National Account

Variables	1995	1996	1997	1998
Consumption				
KHPS	231.9	228.2	230.9	179.5
National Account	274.0	283.9	287.9	255.3
Labor Income				
KHPS	542.7	606.9	618.2	466.2
National Account	457.7	513.9	530.8	493.7

Note: 1. All quantities reported in this table are averages in 10,000 1995 Korean Won, which was equi-valent to about US$8. They are converted into per adult-equivalent terms by giving individuals of age 16 or above a weight of 1 and those younger than 16 a weight of 0.5. The KHPS consumption (labor income) in this table is the ratio of total consumption as defined on p. 2 (labor income) over the adult-equivalent population size in the sample.

2. The national account consumption measure we construct includes food, beverages and tobacco, clothing and footwear, housing, water, electricity, gas and other fuels, and transport expenses to match as closely as possible to the KHPS consumption measure that we use in our tests. The national account labor income measure is the compensation to employees.

3. The number of households is 2008 for every sample year in this and other tables.

count counterpart. In general, Deaton (2005) found that for most countries in his study, including many industrialized countries, household survey data and national account data on consumption and income differ substantially due to the different definitions used and the possibility that the rich are less willing to participate in household surveys.

On the other hand, the KHPS data has its limitations. The drops in the KHPS measures of consumption and labor income are both much larger than in the national account. See Table 1 for details.

In Table 2, we present the cross-sectional means and standard errors (i.e. volatilities) of household consumption, incomes, and asset in adult-equivalent terms. The most salient feature of these data is the declines in both the means and volatilities of consumption, labor income and its change, and total income when the crisis hit the Korean economy. In particular, the mean of consumption dropped by nearly a quarter, from 241.8 to 185.2, and its volatility drastically went down by 54% to 104.7. The mean of gross asset dropped, but its volatility went up in the last

TABLE 2.

The Means and Standard Errors Across KHPS Households of Consumption, Income, and Asset Per Adult-Equivalent

Variables	1995	1996	1997	1998
Consumption	246.6	239.0	241.8	185.2
	(248.0)	(211.8)	(228.2)	(104.7)
Labor Income	543.2	603.4	608.5	453.3
	(407.9)	(529.6)	(478.0)	(373.9)
Labor Income Change	-	60.2	5.1	−155.2
		(465.5)	(408.7)	(366.5)
Total Income	675.8	770.9	774.0	547.9
	(726.8)	(835.9)	(758.4)	(642.0)
Gross Asset	2158.6	2346.1	2509.8	2417.9
	(3407.3)	(3220.2)	(3156.5)	(3890.0)

Note: 1. The quantities reported in this table are denominated in 10,000 1995 Korean Won, which was equivalent to about US$8. Standard errors are in parentheses.
2. The consumption figures in this table are different from those in Table 1 because here the figures are the averages of household per adult-equivalent consumption, whereas in Table 1 the consumption figures are the total consumption from sample households divided by the adult-equivalent population size in the sample.
3. Total income is the sum of labor income and asset income.

sample year. Another salient feature of these data is that the volatilities of these variables remained stable before the advent of the crisis.

Our consumption measure includes expenditures on food (including eating out), housing, heating and other fuel, public utilities, clothes, and gasoline and car maintenance. Our asset measure includes savings deposits (including those held in insurance accounts), stock and bond holdings, and real estate (including land, building, and own housing). The income measure that we use as instrumental variable is total income, i.e. the sum of labor income and asset income (including interest income, dividend income, and income earned through leasing or selling land, house or other buildings). We justify the inclusion of asset income in p. 8. On the other hand, we use the change in labor income (net of any transfers) to proxy "endowment" shock in our tests.

3. MECHANISMS OF CONSUMPTION SMOOTHING

To understand intuitively how much consumption smoothing there was in the face of the crisis, we report in Table 3 households' uses of five smoothing mechanisms. For the information on what is included in each mechanism, see the notes to this table. Every cell of this table contains three numbers.

TABLE 3.

Smoothing Mechanisms in Per Adult-Equivalent Terms: Averages and Standard Errors Across Participating Households and Percentage of Participation

Smoothing Mechanisms	1995	1996	1997	1998
Outstanding Loan	469.8	517.2	519.7	712.8
	(687.9)	(693.2)	(765.8)	(2972.9)
	[47.8%]	[50.4%]	[48.6%]	[48.1%]
Private Transfer Received	106.6	117.2	120.7	104.9
	(133.0)	(133.4)	(207.4)	(159.3)
	[14.1%]	[18.1%]	[20.1%]	[22.4%]
Public Transfer Received	103.0	93.5	74.9	55.8
	(141.0)	(142.7)	(139.3)	(137.5)
	[6.7%]	[7.6%]	[10.1%]	[17.1%]
Asset Liquidation	531.8	643.3	615.2	508.1
	(1142.1)	(1154.9)	(1150.7)	(798.3)
	[12.9%]	[11.0%]	[10.7%]	[11.6%]
Durable Goods Purchase	34.1	35.5	34.3	29.4
	(35.0)	(44.1)	(39.0)	(34.4)
	[34.3%]	[27.3%]	[28.9%]	[17.8%]

Notes: 1. The reported quantities are averages across participating households in a consumption smoothing mechanism in 10,000 1995 Korean Won per adult-equivalent. Standard errors are in parentheses. The percentages in brackets are the fractions of 2008 sample households utilizing a particular smoothing mechanism.
2. Outstanding loans include borrowings from financial and non-financial institutions for living expenses, purchase of land or house, purchase of car or household goods, and other purposes.
3. Private transfer includes support and gifts from relatives and friends in cash or in kind.
4. Public transfer includes unemployment insurance, support from government and social organizations in cash or in kind for the poor, disabled, disaster relief and others, national pension, pension for private-school teachers and government/military officials, and veteran's pension.
5. Liquidation of assets includes selling financial and real assets.

The top number is the average real Korean Won amount in per adult-equivalent terms across households that used a smoothing mechanism, the middle one is the associated standard error, and the bottom one is the percentage of sample households that used a certain mechanism. Each of these three numbers reveals a different aspect of the utilization of a smoothing mechanism, as we will see shortly.

Among these five mechanisms, borrowing was the most important in terms of the fraction of households relied on it as well as the average Won amount. While the percentage of sample households that had outstanding loans barely changed during the crisis year, the balance of such loans went up sharply from 519.7 to 712.8 (i.e. an increase equivalent to almost US$1,600 per adult-equivalent among those who borrowed). This sharp increase was perhaps mostly for cross-sectional smoothing because the drop in aggregate consumption shown in Table 1 suggests lack of intertemporal smoothing for Korea as a whole. This interpretation is reinforced by the nearly three-fold increase in the standard error of outstanding loans during the crisis year, which suggests that households were borrowing vastly different amounts to cope with the uneven distributional effects of the crisis.

The second row of Table 3 shows a decrease in the average private transfer during the crisis year from 120.7 to 104.9, accompanied by a slight increase in the fraction of households receiving it. In contrast, the share of households receiving public transfers went up dramatically from 10.1% before the crisis to 17.1% afterwards. But the declining amount of such transfer seems to indicate that the government was spreading it more thinly than before to cover more households in need of such assistance. The standard errors of these two transfers did not change significantly from pre-crisis years. Transfers obviously facilitate cross-sectional smoothing.

Asset liquidation (fourth row) can serve either cross-sectional or intertemporal smoothing, depending on purpose. The simultaneous decline in mean and standard error here is consistent with the well-known fact that real and financial assets commanded lower prices during the crisis. Selling assets therefore played a less significant role in cushioning consumption during the crisis than before.

The last row reports a mechanism that is perhaps mainly associated with intertemporal smoothing: delaying durables purchase. Though 11% of the households delayed or canceled durables purchase (as the fraction that purchased durables declined from about 29% to about 18%), the reduction in such purchase was only about US$40 per adult-equivalent, suggesting that this was unlikely to be an important mechanism.[6]

To summarize, there was obviously substantial amount of cross-sectional consumption smoothing going on during the crisis year. The question, again, is if it was adequate to fight off the impact of the crisis.

4. TESTS

OZ (2001) emphasized that the testable implications of the risk sharing hypothesis depends on consumers' risk attitudes. They found evidence for DRRA, and reported that allowing for DRRA in their tests changed some

[6]Browning and Crossley (2003) and Blundell, Pistaferri and Preston (2005) both reported that cutting or delaying expenditures on durables is one way to smooth non-durables and services consumption.

of the earlier results in the literature. In ZO (2004), their tests are shown to have power against several versions of PIH. We therefore adopt their tests and refer the reader to ZO (2004) for details. However, as we mentioned in the Introduction, one major innovation in this section is our inspection of two standard test statistics in the presence of substantial variations in the cross-sectional means and volatilities of consumption and other variables. In addition, we carefully choose our instruments that are valid under the null hypothesis of full risk sharing using economic theory as a guide.

The full risk sharing hypothesis can be tested by estimating the following equation:[7]

$$C_h(t+1) = \gamma[1-\phi(t+1)] + \phi(t+1)C_h(t) + \nu_h(t+1) \tag{1}$$

where C_h denotes the per adult-equivalent consumption for household h measured with error, γ is the common subsistence level and is estimated from data,[8] ϕ is the common growth rate of $C_h - \gamma$ across all households under full risk sharing and represents the aggregate shock to consumption. The disturbance term $\nu_h(t+1)$ under the hypothesis of full risk sharing consists of consumption measurement errors from period t and $t+1$. Due to the measurement errors in consumption, $C_h(t)$ and $\nu_h(t+1)$ are correlated across households. Therefore, an instrumental variable procedure is needed to estimate (1) consistently. An implication of full risk sharing that has been explored extensively in the literature is that individual consumption does not respond to idiosyncratic shocks after controlling for the aggregate shock. See e.g. Cochrane (1991). Eq. (1) can accommodate such a test by including idiosyncratic shocks such as the change in labor income and the changes in marital status or employment status.

According to standard microeconomic theory, under full risk sharing (see e.g. ZO (2004, eq. (5)) depends on household h's welfare weight in the social planner problem. For the solution to this problem to be equivalent to a competitive equilibrium, a household's welfare weight should depend on its initial wealth when the risk sharing pool is set up, which is certainly unknown.[9] However, as long as households' initial wealths are sufficiently

[7] This equation is derived in ZO (2004) under the following assumptions. Households in a risk sharing pool share the same time preference rate and identical probabilistic belief about the state of the world. A social planner maximizes the weighted, discounted, household lifetime expected (power) utility defined over adult-equivalent consumption net of subsistence level, subject to the aggregate resource constraint for the risk sharing pool at every period. Leisure can enter the utility function if it is separable from consumption.

[8] The subsistence parameter, despite being so called, does not have to be positive because its role is to control how relative risk aversion (RRA) changes with consumption. If it is positive, DRRA holds. If it is negative, increasing RRA holds. If it is zero, it implies the familiar constant RRA case. See ZO (2004).

[9] The initial wealth in our data set is not necessarily the initial wealth in the theory.

correlated with their assets of the sample periods in the cross section, we can use the latter to instrument C_h. The asset measure we adopt is gross asset (i.e. not net of debt) because net asset informs a household's welfare weight in the social planner problem, and borrowing is an important channel of pooling consumption risk as we see in Table 3. Therefore under the null of risk sharing gross asset is a valid instrument.

Another possible instrument is total income that includes asset income and labor income. Total income may be cross-sectionally correlated with consumption under the null of risk sharing for three reasons. First, there may be positive correlation between asset income and initial wealth, i.e. asset income may inform welfare weights. Second, asset income may be a channel to share consumption risk. Third, labor income may also inform initial wealth if the latter impinges on human capital accumulation.

Let A_h and Y_h represent household h's financial assets and total income, and let $Z_h = (1, A_h, Y_h)'$ be the vector of instruments. If the measurement errors in consumption are not cross-sectionally correlated with A_h and Y_h and their measurement errors, then under full risk sharing the cross-sectional sample mean of $\nu_h(t+1)Z_h(t)$ should converge in probability to a zero vector. These orthogonality conditions can be used to estimate the unknown parameters γ and ϕ's in (1) and to test the risk sharing hypothesis with the 2-step generalized method of moments (GMM) of Hansen (1982) that is robust to conditional heteroscedasticity and serial correlation in $\nu_h Z_h$.

One reason that the contemporaneous measurement error in consumption and that in income are correlated across households is that there may be production for own consumption, compensation in kind, and employer-paid consumption. This may especially be relevant for farmers, fishermen and other rural households, which account for 40% of the KHPS sample. If some of these households failed to include such consumption/ production in their reported consumption/income measures, they became common measurement error in reported consumption and income. In addition, the measurement error in consumption and that in gross asset can be correlated as well, if reported assets do not include the storage of output from own production or ignore non-financial assets such as livestock, and consumption from them is not reported. Therefore, once the measurement errors are taken into account, $\nu_h(t+1)$ and $Z_h(t)$ can be correlated across households through the measurement error at period t. Consequently, the sample mean of $\nu_h(t+1)Z_h(t)$ may not converge in probability to a zero vector even under full risk sharing. The common measurement errors in consumption, income, and asset may also cause spurious correlation between the endogenous regressor $C_h(t)$ and the instruments in $Z_h(t)$, leading to spuriously valid instruments. To avoid these problems and the possible serial correlation in measurement errors, which may cause the consumption

TABLE 4.

Instrument Relevance

First-Stage Regression	Estimates/Statistics
1995 Consumption on	
1997 Gross Asset	0.0078 (4.39)
1997 Total Income	0.0711 (9.59)
Constant	171.9 (21.3)
Adjusted R^2	0.07
F(2, 2005) with (p-value)	76.4 (0.000)
1996 Consumption on	
1998 Gross Asset	0.0043 (3.63)
1998 Total Income	0.0761 (10.58)
Constant	187.0 (28.9)
Adjusted R^2	0.06
F(2, 2005) with (p-value)	68.6 (0.000)
1997 Consumption on	
1995 Gross Asset	0.0039 (2.51)
1995 Total Income	0.0626 (8.71)
Constant	191.1 (27.3)
Adjusted R^2	0.05
F(2, 2005) with (p-value)	52.9 (0.000)

Note: The t-ratios are in the parentheses after parameter estimates.

measurement error in $C_h(t)$ to be correlated with those in $Z_h(t+1)$ or those in $Z_h(t-1)$, we use $Z_h(t+2$ or $Z_h(t-2)$ to instrument $C_h(t)$. For example, the 1995 total income and total asset are used to instrument the 1997 consumption. Both instruments are significant at the 1% level. See Table 4.

We now turn to the behavior of two tests, Hansen's J test of model specification (i.e. a chi-square test) and the usual t test of the significance of a parameter estimate. It is useful to start with the covariance matrix of $\nu_h Z_h$, i.e.[10]

$$S \equiv E_H[\nu_h^2(t+1)Z_h(t \pm 2)Z_h'(t \pm 2)], \tag{2}$$

because its sampling counterpart plays a central role in both the J statistic (i.e. the minimized GMM criterion function) and the covariance matrix for GMM estimates of model parameters. Furthermore, for the ease of exposition we focus on the case of conditionally homoskedastic $\nu_h(t)$ so

[10] E_H below denotes expectations operator defined over the cross-sectional dimension.

that (2) simplifies to

$$S = E_H\lfloor v_h^2(t+1)\rfloor \cdot E_H[Z_h(t\pm 2)Z_h'(t\pm 2)].$$

Denote the sampling counterpart of S by $\hat{S}$, the sampling counterpart of $E_H[\nu_h^2(t+1)]$ by $\hat{\sigma}^2_{\nu_h}$, and the sampling counterpart of $E_H[Z_h(t\pm2)Z_h'(t\pm2)]$ by $S_{Z_hZ_h'}$. Let N be the number of sample households (in a prefecture). Let $\hat{\nu}_h(t+1)$ be the residual obtained from eq. (1) by plugging in the GMM estimates $\hat{\gamma}$ and $\hat{\phi}(t+1)$. In addition, let $\overline{C}_h(t)$ be the cross-sectional mean of consumption at period t, and $\hat{\sigma}_{C_h(t+1,C_h(t)}$ be the cross-sectional sampling covariance between consumption of time $t+1$ and time t. We obtain

$$\begin{aligned}
\hat{\sigma}^2_{\nu_h} &= \frac{1}{N-2}\sum_h \hat{\nu}_h(t+1)^2 \\
&= \hat{\sigma}^2_{C_h(t+1)} + \hat{\phi}(t+1)^2\hat{\sigma}^2_{C_h(t)} + \overline{C}_h(t+1)[\overline{C}_h(t+1) - 2\hat{\phi}(t+1)\overline{C}_h(t)] \\
&- 2\hat{\phi}(t+1)\hat{\sigma}_{C_h(t+1),C_h(t)} \\
&+ \overline{C}_h(t)[\hat{\phi}(t+1)^2\overline{C}_h(t) - 2(1-\hat{\phi}(t+1))^2\hat{\gamma}] + (1-\hat{\phi}(t+1))^2\hat{\gamma}^2.
\end{aligned}$$

Now imagine what the data on consumption in Table 2 imply about the crisis-year $\hat{\sigma}^2_{\nu_h}$, given the previous year's mean $\overline{C}_h(t)$ and variance $\hat{\sigma}^2_{C_h(t)}$ of consumption. Since the cross-sectional mean $\overline{C}_h(t+1)$ and variance $\hat{\sigma}^2_{C_h(t)}$ have declined sharply during the crisis year, we should tend to expect a large decline in $\hat{\sigma}^2_{\nu_h}$.[11] This in turn drives up Hansen's (1982) J statistic because $\hat{S}^{-1} = S^{-1}_{Z_hZ_h'}/\hat{\sigma}^2_{\nu_h}$ is the weighting matrix in the GMM criterion function, and $S^{-1}_{Z_hZ_h'}$ does not depend on $\hat{\sigma}^2_{\nu_h}$. Nor does $S^{-1}_{Z_hZ_h'}$ involve the crisis-year data, since the instruments are two years apart from the endogenous regressor $C_h(t)$, and one year apart from the crisis-year consumption $C_h(t+1)$. In other words, even if full risk sharing is true for the crisis year so that $E_H[\nu_h(t+1)Z_h'(t\pm 2)] = 0$ still holds at the true parameter values, simply because of the substantial decline in $\hat{\sigma}^2_{\nu_h}$ the J test of risk sharing for that year can become very large relative to the J statistic based on previous year's data, yielding a rejection of the benchmark model for the crisis year.

Of course, the weighting matrix is just the central component of the GMM criterion function that is a quadratic form in the sample analogue of

[11] And this is what happens in the empirical results. See next section. Conceptually, however, there is the possibility that the decline in $\hat{\sigma}^2_{\nu_h}$ may not be large, or $\hat{\sigma}^2_{\nu_h}$ might even increase, for two reasons. First, as $\overline{C}_h(t+1)$ drops, $\hat{\phi}(t+1)$ decreases. Second, the sampling covariance $\hat{\sigma}_{C_h(t+1),C_h(t)}$ tends to decrease as $\hat{\sigma}^2_{C_h(t+1)}$ declines. This possibility does not need to concern us because as an empirical matter, it only happens in one of fourteen prefectures in our data, namely, Gwangju.

$E_H[\nu_h(t+1)Z'_h(t\pm2)]$. However, it can be shown that under the assumption $E_H[\nu_h(t+1)Z'_h(t\pm2)] = 0$ the criterion function evaluated at the efficient GMM estimates can eventually be written as

$$\begin{aligned} J \;\approx\; & \frac{w_{11}}{N} + \frac{1}{\hat{\sigma}^2_{\nu_h}} \left[w_{11} \frac{2\sum_{i\neq j} \hat{\nu}_i \hat{\nu}_j}{N^2} + (w_{12}+w_{21}) \frac{\sum_h \hat{\nu}_h}{N} \cdot \frac{\sum_h \hat{\nu}_h Y_h}{N} \right. \\ & + \; (w_{13}+w_{31}) \frac{\sum_h \hat{\nu}_h}{N} \cdot \frac{\sum_h \hat{\nu}_h A_h}{N} + w_{22} \left(\sum_h \frac{\hat{\nu}_h Y_h}{N} \right)^2 \\ & \left. + \; (w_{23}+w_{32}) \frac{\sum_h \hat{\nu}_h Y_h}{N} \cdot \frac{\sum_h \hat{\nu}_h A_h}{N} + w_{33} \left(\frac{\sum_h \hat{\nu}_h A_h}{N} \right)^2 \right], \end{aligned} \tag{3}$$

where Y_h and A_h are from year $t \pm 2$, w_{ij} is the i-jth element of the weighting matrix $\hat{S}^{-1}$ and only depends on instruments of period $t \pm 2$, and none of the terms in the square bracket can be written as a function of $\hat{\sigma}^2_{\nu_h}$ or $\overline{C}_h(t+1)$ and $\overline{C}_h(t)$. The J statistic is therefore a decreasing function of $\hat{\sigma}^2_{\nu_h}$.

We now look into the t statistic that tests the significance of an individual parameter estimate, especially the estimate of the slope coefficient (called η) on a shock variable such as labor income change.[12] Denote $\theta = (\gamma, \phi(t+1), \eta)'$. The covariance matrix for the GMM estimates of model parameters is $(\hat{G}'\hat{S}^{-1}\hat{G})^{-1} = \hat{\sigma}^2_{\nu_h}(\hat{G}' S^{-1}_{Z_h Z'_h} \hat{G})^{-1}$ under conditional homoskedasticity, where $\hat{G}$ is the sampling counterpart to $E_H\{\partial[\nu_h(t+1)Z'_h(t\pm2)]/\partial\theta'\}$ and is evaluated at $\hat{\theta}$, the GMM estimates. Since $\hat{G}$ is independent of $\hat{\sigma}^2_{\nu_h}$ and does not involve crisis-year consumption,[13] a smaller $\hat{\sigma}^2_{\nu_h}$ leads to smaller standard error for the slope estimate on the shock variable (and other parameters). This raises the t statistic for the slope coefficient, and biases the t test against the null hypothesis of full risk sharing when the crisis-year data is used.

Although the J test and the t test are both biased towards rejecting the benchmark of risk sharing, the size of bias can be larger in one test than in

[12] When a shock variable with coefficient η is added to eq. (1), ν_h is understood to be the disturbance in eq. (1) adjusted for this additional term. And $\hat{\sigma}^2_{\nu_h}$ will include three additional terms, namely, the variance of this shock term with the coefficient η^2 and the covariances between the shock and $C_h(t+1)$ and $C_h(t)$, with $-\eta$ and $-\eta\phi(t+1)$ being the respective coefficients. However, the cross-sectional volatility of labor income change/shock for the crisis year does not change much from its previous values as shown in Table 2. Hence its impact on J and t statistics may be very small.

[13] When the test is done for the crisis year, $\hat{G}$ includes the cross sectional means of a shock from the crisis year and of the product between this shock and each of the two instruments. As such, the standard errors of parameter estimates are affected by the second moments of the shock variable in an ambiguous way. However, as an empirical matter, their effects are either very small or they reduce $\hat{\sigma}^2_{\nu_h}$ as well, otherwise, the results in Table 6 below cannot be rationalized.

the other. Therefore, these two tests may produce different results. Such difference is nonetheless informative. If one of these two tests rejects the benchmark model, but the other does not, the model is considered non-rejected.[14] This is because if the model does not hold, both tests should reject it, given they are both biased towards rejecting.

Although the discussion above is based on assuming conditional homoskedasticity, such biases should carry through to the case of conditional heteroskedasticity because $\hat{\sigma}^2_{\nu_h}$ is in $\hat{S}$ even then.[15] Finally, since our instruments are two periods away from the endogenous variable $C_h(t)$, the tests for the 1995-1996 and the 1997-1998 sample years will not involve instrument values from the crisis year. Therefore, even though $\hat{G}$ and $\hat{S}$ include the second moments of instrumental variables, the somewhat large variations in the volatilities of the instruments during the crisis year do not interfere with the J test and the t test when they are conducted using the 1995-1996 and the 1997-1998 sample years.

5. EMPIRICAL RESULTS

In the row labeled "1995-1998" in Table 5, we report the results based on the whole sample. Hansen's J test of over-identifying restrictions rejects the orthogonality between ν_h and Z'_h at the 5% significance level, indicating that the null hypothesis of full risk sharing at prefecture level for the entire 4-year period is at odds with data. It also implies that full risk sharing at national level for the whole sample period is rejected. The next row tests risk sharing during the crisis year. Again, we find that prefecture-level full risk sharing is rejected by the J test at the 5% level.

However, when we test this hypothesis for the 1995-1997 period, the model is not rejected at conventional significance levels: the J statistic now implies a p-value of 19.5%.[16] Therefore, the rejection in the first row is clearly driven by the rejection of prefecture-level full risk sharing during the crisis year in the second row. In the rest of this section, we call the test in the third row of this panel the base run. The non-rejection here is notable because the number of data points used in the base-run test for the pre-crisis sample is twice that used in the test for the crisis-year sample, due to the fact that one more year's data is used in the base run.

[14] Or if both tests fail to reject our benchmark model, the model is considered non-rejected. On the other hand, if both tests reject risk sharing, it is not informative. This is because it could be that the model holds, but is rejected by these over-rejecting tests; meanwhile it could also be that the model does not hold, and is rejected by these two tests.

[15] Our tests reported in the next section do not assume conditional homoskedasticity.

[16] This of course raises the question of whether national risk sharing holds for 1995-1997. We have tested it, and have found that it is rejected at the 10% level.

TABLE 5.

Risk Sharing at Prefecture Level

Period/ Shock	γ (s.e.)	γ_L (s.e.)	η (s.e.)	J (p-value)	C (p-value)
Tests of Equation (1)					
1995-1998	99.2 (7.4)			108.2 (0.033)	
1997-1998 (Crisis Year)	107.4 (9.3)			41.1 (0.041)	
1995-1997	82.0 (20.0)			63.8 (0.195)	
1995-1997	94.3 (51.4)	81.8 (19.6)		63.7 (0.172)	0.05 (0.823)
Tests of Equation (1) Augmented with Shocks One at a Time, 1995-1997					
Labor Income Change Added	57.6 (30.7)		0.08 (0.03)	57.1 (0.362)	6.70 (0.010)
Marital Status Change Added	81.0 (19.8)		−120.8 (111.3)	62.6 (0.198)	1.18 (0.277)
Employment Status Change Added	83.2 (20.3)	−50.6 (56.3)	63.0 (0.189)	1.55 (0.213)	
Tests of Equation (1) Augmented with Shocks One at a Time, 1997-1998					
Labor Income Change Added	110.9 (11.8)		−0.08 (0.04)	75.0 (0.000)	33.9 (0.000)
Marital Status Change Added	96.0 (13.9)		155.7 (109.5)	38.7 (0.053)	2.4 (0.121)
Employment Status Change Added	129.0 (11.8)		−27.2 (42.7)	80.6 (0.000)	39.5 (0.000)

Notes: 1. The number of data points used in a test is 2008 times the number of involved sample years.

2. The numbers in parentheses below the parameter estimates are standard errors. J and C are chi-squared statistics. The numbers in parentheses below J and C statistics are p-values.

3. In the last row of the upper panel, the γ estimate is for households with gross asset per adult-equivalent above the median, and the γ_L estimate is for the remaining households. They are obtained by using the 2nd step GMM estimates of the third row as initial values and the weighting matrix there as the weighting matrix.

4. The C statistic tests $\gamma = \gamma_L$ in the fourth row, and $\eta = 0$ in the following rows. It is the difference between the J statistic of a row and that of the third row (for pre-crisis years) or fourth row (for the crisis year) in the top panel, depending on the sample period.

5. The η estimates are for coefficients on the three shocks. The coefficients are assumed to be constant across prefectures. This is especially necessary for the latter two shocks because of their low incidences during 1995-1997.

6. When η is allowed to be different across prefectures, only one η estimate was significant at 5% level (Gyeongbuk) and one at 10% level (Chunbuk) for the pre-crisis period.

We would expect that the more years of data (i.e. the more data points) are used, the more power our test should have, and the more likely it is to reject the prefecture-level full risk sharing benchmark for the pre-crisis years if risk sharing was far from being complete. However, despite that fact, this benchmark is not rejected for the period 1995-1997. The results in the second and the third rows therefore constitute strong evidence for substantial amount of consumption smoothing across households for pre-crisis years, but the opposite seems to hold for the crisis year.

In the base run, the subsistence level, γ, is estimated to be 82.0 with a standard error of 20.0. The positive γ estimate implies DRRA as we mentioned in Footnote 9. Furthermore, since the γ estimate is about 1/3 of the pre-crisis average basic consumption presented in Table 2, it cannot possibly be interpreted as the bliss point in consumption. This result alone is evidence against three versions of an alternative model on consumption smoothing, the PIH based on the quadratic preferences. See ZO (2004) for details.

To explore the possibility that γ varies with wealth class, we allow γ for households with above-median gross asset to be different from that (γ_L) of the remaining households, and redo the tests. The results are in the fourth row of Table 5. The γ estimate for the wealthier half of the sample is 94.3, and is close to the estimate for the other half, 81.8. Not surprisingly, the C test does not reject the restriction $\gamma = \gamma_L$: the p-value of this test is 82.3%.

As a further check on the robustness of the non-rejection of risk sharing at prefecture level for the pre-crisis years, we test if consumption then responded to three idiosyncratic shocks in the lower panel of Table 5. The slope coefficient (η) estimate for labor income change (ΔY_h^L) is significant, and the C statistic for this case rejects $\eta = 0$. However, when η is allowed to be different across regions, the significance of ΔY_h^L occurred in just two prefectures at the 10% level (not reported to conserve space; however, see Note 6 under this table for further details.) Furthermore, Cochrane (1991) pointed out that full risk sharing is compatible with a significant income change term if consumption and leisure are not separable in the utility function.[17] So the evidence based on the effect of labor income change alone mostly favors our benchmark. In addition, the results based on marital status and employment status changes support the risk sharing benchmark even more strongly, as neither slope estimate is significant. Therefore, the overall empirical evidence here is consistent with the base run.

However, when these same tests on individual shocks are conducted using the crisis-year data, the results are different from the rejection that the J

[17]Alternatively, the noisiness of the measured income change could also cause the significance in these two prefectures at the 10% level. This may explain the significant but negative slope estimate for labor income change in the bottom panel.

TABLE 6.

Comparison of Residual Variance $\hat{\sigma}^2_{\nu_h}$

Prefecture	$\hat{\sigma}^2_{\nu_h}$ for 1995-96	$\hat{\sigma}^2_{\nu_h}$ for 1997-98	% Change in $\hat{\sigma}^2_{\nu_h}$
Seoul	37400139.0	8177130.6	−78%
Busan	4430401.8	1435583.8	−68%
Daegu	3300634.5	783246.9	−76%
Incheon	1559151.5	1275937.8	−18%
Gwangju	1687573.6	6186583.0	27%
Daejeon	742549.1	238678.7	−68%
Gyeonggi	10602613.0	5146218.4	−51%
Gangwon	1764131.7	809492.8	−54%
Chungbuk	1368885.9	395274.5	−71%
Chungnam	9182093.2	608801.1	−93%
Jeonbuk	4065148.9	597616.0	−85%
Jeonnam	4684252.9	514905.8	−89%
Gyeongbuk	4746747.1	969186.0	−80%
Gyeongnam	6253823.3	1061115.6	−83%

Note: We compare these two years because the instruments for both sets of tests are from pre-crisis years only.

test produces in the second row of Table 5. Rather, they are consistent with the null hypothesis of full risk sharing. See the lower panel of this table. Here again, only labor income change is significant, but with the wrong sign.[18] The other two shocks are insignificant. If risk sharing is to be rejected as the J test for 1997-1998 indicates, the effects of these three shocks should be significantly different from 0 and have the negative sign.

That these three shocks do not significantly affect the crisis-year consumption is strong evidence for prefecture-level risk sharing for that year. This is because the standard errors of the η estimates would be even larger for that year if the cross-sectional mean and volatility of consumption did not drop substantially to lower $\hat{\sigma}^2_{\nu_h}$, as we have explained in Section 4. On the other hand, such lower $\hat{\sigma}^2_{\nu_h}$ produced by the decline in the mean and volatility of consumption across households explains the rejection of risk sharing by the J test for the crisis year, according to eq. (3).

[18]The standard errors of parameter estimates based on the 1997-1998 data should be smaller than those obtained from the data of a pre-crisis year, according to our theory in Section IV. This prediction is largely confirmed by comparing the standard errors of the 1997-1998 period with those of the 1995-1997 period in Table 5. The only exception is the standard error of η for the change in labor income due to the fact that more data points are used in the estimation presented in the middle-panel of this table. This exception disappears when we restrict the pre-crisis sample to a one-year sample.

To check if $\hat{\sigma}^2_{\nu_h}$ does decline for the crisis year, we report in Table 6 the $\hat{\sigma}^2_{\nu_h}$ constructed using the parameter estimates obtained using 1995-1996 data and the crisis- year data separately. For 13 of all the 14 prefectures, there is substantial drop in $\hat{\sigma}^2_{\nu_h}$. And for 12 of these 13 prefectures, the drop is over -50%. Therefore, the evidence here supports our theory on how the J and t statistics can be biased by large declines in the cross-sectional mean and volatility of consumption.

6. CONCLUSIONS

We have conducted a "stress test" of the workings of the consumption smoothing mechanisms in the Republic of Korea using a panel data set that covers the period 1995-1998. We do not reject full risk sharing at prefecture level for pre-crisis years for at least twelve of the fourteen Korean prefectures using both the standard model specification test in the GMM framework and the usual t test. For the crisis year, the test results from these two standard tests diverge: the prefecture-level full risk sharing is rejected by the model specification test, but judged by the t test the consumption in that year did not respond to idiosyncratic shocks. We attribute this discrepancy in the two sets of test results for the crisis year to the sensitivity of the standard test statistics to the variations in cross-sectional means and volatilities of consumption and other variables. Therefore, the uneven distributional effects of the crisis across individuals within a prefecture seemed to be sufficiently insulated by the expansions of public and private transfers and borrowing. On the other hand, since our findings do not support national full risk sharing even for pre-crisis years, they are still consistent with the studies that reject full risk sharing at the national level cited in the Introduction.

In addition, it is worth mentioning that we confirm the previous empirical evidence for DRRA uncovered from data of low-income countries, because our subsistence estimate is positive, almost always statistically significant, and does not vary significantly across wealth class. Since Korea has an industrialized economy, the case for DRRA seems more compelling.

Our study is subject to one important limitation. Although the crisis broke out in the fourth quarter of 1997, the Korean economy showed signs of stabilization only in 1999. So the "crisis period" in our tests does not include the latter part of the actual crisis period, due to the fact that the KHPS survey was discontinued in 1998. Whether our test results for the crisis period would change, had more data been available, is a tantalizing question.

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