Measuring the Effects of Decollectivization on China's Agricultural Growth: A Panel GMM Approach, 1970-1987

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Abstract: The mainstream view that decollectivization significantly contributed to China's agricultural growth has recently been challenged by revisionists, who emphasize the positive effects of the socialist legacy, such as irrigation and mechanization. This study contributes to this debate by explicitly recognizing the endogeneity of institutional changes and uses lagged weather shocks as valid instruments. With improved data on irrigation and mechanization in a provincial-level dataset covering the1970-1987 period, the results of panel GMM estimations reveal that the Household Responsibility System had a significantly positive effect on China's agricultural growth, which is larger than that indicated by OLS estimates.

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1. Introduction

During the1979-1984 period, China's rural sector experienced dramatic institutional changes that marked the end of collective farming based on production teams (which were the equivalents of villages) and the rise and eventual predominance of household-based farming, commonly known as the "Household Responsibility System" (HRS hereafter). Although only 1.02 percent of all production teams had converted to HRS by the end of 1979, this share rapidly rose to 98 percent by the end of 1983 (see Figure 1).

Equally impressive was China's agricultural growth during the late 1970s and early 1980s, which was unprecedented, particularly when compared with the previous collective era.[1] Did the HRS cause this agricultural miracle in China circa 1980? Theoretically, this result may have ensued because the HRS solved the well-known free-rider problem inherent in production teams – which arises as a result of the extreme difficulty in monitoring agricultural production –by effectively making the peasants residual claimants (Alchian and Demsetz, 1972; Lin, 1988; Nolan, 1988).[2] However, the incentive advantage of the HRS is likely tempered by the alleged efficiency of agricultural collectives in exploiting returns to scale, particularly regarding the provision of public goods, such as irrigation infrastructure.

Empirically, causal identification is complicated by the confounding effects of concurrent

[1] According to Lin's (1992) calculation, China's agricultural sector grew at an annual rate of 7.7 percent during the 1978-1984 period, which was much higher than the 2.9 percent annual growth rate experienced during the 1952-1978 period.
[2] Lin (1990) offers a nuanced "exit rights" explanation for the collapse of agricultural productivity in China under agricultural communes during the 1958-1961 period. However, Kung (1993) finds scant evidence of the existence of exit rights before 1958. Moreover, Kung and Putterman (1997) provide partial evidence that the absence of a productivity collapse in 1957 may be due to the de facto devolvement of farm operations and organization to households.

market-oriented reforms, particularly price reforms. There were two distinct prices in the state commercial system before the reforms, i.e., quota prices for crops sold to fulfill procurement obligations and above-quota prices with a premium for crops sold in excess of the obligation. In 1979, average procurement prices, which included quota and above-quota prices, increased by 22.1 percent (State Statistical Bureau, 1984, p. 403). Only considering the marginal prices, i.e., the above-quota prices, yields an increase of 40.7 percent. Even if we control for the prices of rural industrial products, which were stable circa 1980, the above-quota prices increased by 40.4 percent (see Figure 1).[3] Moreover, the government began to encourage the revival of rural markets by lifting restrictions on products that could be traded privately in the free market. By 1984, more than 18 percent of all purchases of agricultural products were made at market prices (Sicular, 1988b). Prices in rural markets were generally stable during the 1979-1984 period, even relative to the prices of rural industrial products (see Figure 1).[4] Another complication is that the availability of manufactured inputs, particularly chemical fertilizers, increased substantially during this period (see Figure 1), in addition to other technological changes, such as the diffusion of high-yielding seed varieties (Stone, 1988; Huang and Rozelle, 1996).

[3] See Sicular (1988a) for a detailed chronology of price changes in 1979 and thereafter.
[4] Rural market prices declined slightly in 1979 and 1980, which might have been due to increased supply as more products became available in the free market.

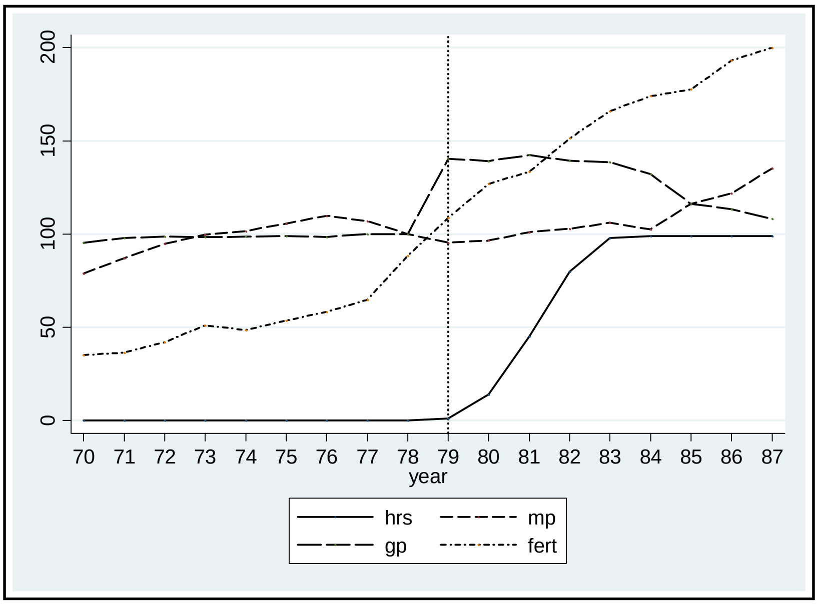

Figure 1. China's Rural Conditions during 1970-1987

Notes: *hrs* is the percentage of all production teams converted to the HRS by the end of each year; *gp* is the index of state above-quota prices relative to input prices (1978=100), *mp* is the index of rural market prices relative to input prices (1978=100), and *fert* is consumption of chemical fertilizers in 100,000 tons. Data source: Lin (1992) and State Statistical Bureau (2010) for *fert*.

A number of approaches have been applied to assess the effects of decollectivization on China's agricultural growth. The simplest approach is to compare total factor productivity (TFP) in China's farming sector before and after the reform period (Wen, 1993; Fan and Zhang, 2002). Most efforts to separate and disentangle the effects of the HRS, price reforms, and technological changes are undertaken through growth accounting (McMillan et al., 1989; Fan, 1991; Kalirajan et al., 1996; Zhang and Carter, 1997), which measures the contribution of the HRS indirectly as a residual. A more convincing approach exploits the tremendous variation in the adoption of the HRS across Chinese provinces and directly estimates the effects of the HRS as a right-hand-side variable in fixed effects panel regressions (Carter and Zhong,1991; Lin, 1992) or system regressions (Huang and Rozelle, 1996). Despite differences in samples and approaches, most studies report that the HRS significantly contributed to China's agricultural growth. For example, Lin (1992) finds that decollectivization accounted for approximately half of China's agricultural growth during the 1978-1984 period.

However, these positive studies are not without skeptics and critics (Riskin, 1987; Carolus, 1992; Bramall, 1995, 2000, 2008; Xu, 2012). A common objection is that agricultural growth in China stalled in the late 1980s after the HRS had been universally adopted.[5] The lackluster or even counterproductive performance of rural privatization worldwide (e.g., Eastern Europe) is linked to further suspicion (Rozelle and Swinnen, 2004). Perhaps the most serious challenge has come from Xu (2012), who focuses on the specification errors of Lin (1992), particularly the "wrong usage of HRS adoption rate". The main issue is that Lin (1992) measures the HRS using its adoption rate at the end of each year, whereas the dependent variable is the value of crop output in the current year, which indicates that some of the observed increases in agricultural production might have occurred before the HRS was adopted. To avoid this problem, Xu (2012) uses a lagged HRS variable in Lin's (1992) framework of two-way fixed effects and finds that the statistical significance of the HRS disappears completely.

However, all studies thus far have ignored an important issue in assessing the growth effects of the HRS –i.e., that institutional changes are typically endogenous (e.g., Acemoglu et al., 2001) – and China's decollectivization was no exception. It is possible that provinces with more to gain from decollectivization adopted the HRS earlier than other provinces (Lin, 1988). However, the government only initially allowed poor, remote and mountainous areas to experiment with the HRS (e.g., Du, 2005). Moreover, measurements of the HRS are likely imperfect (e.g., initial underreporting and subsequent overreporting as political conditions shifted), and errors-in-variable

[5] This result can be easily reconciled by recognizing that the transition to the HRS amounted to a one-time increase in peasants' efforts, i.e., moving from the interior to the frontier of the production function, which does not necessarily have any long-run growth effects.

problems might lead to attenuation bias. Finally, reverse causality is also possible because provinces experiencing substantial output growth rushed to adopt the HRS, and provinces with sluggish growth were slower to adopt the new system.

The primary contribution of this study is to explicitly address the endogeneity of the adoption of the HRS, thereby elevating the current debate to a new level. We use lagged weather shocks as valid instruments for the HRS and employ panel GMM (Generalized Method of Moments) for consistent estimations using Lin's (1992) two-way fixed effects model. Weather adversity is a key determinant of decollectivization in Bai and Kung (2014), a recent study on the endogenous selection of farming institutions in which two measures of weather adversity are both correlated with the HRS and are shown to be significant in the first-stage regressions. It is well known that the HRS was first introduced in Anhui province in late 1978 following a severe drought and was initially adopted as a life-saving measure. During a previous episode, a similar household contracting system also emerged across China following the Great Leap famine (1959-1961), before the government eliminated it (Yang, 1996). However, lagged weather shocks presumably had no impact on the current year's agricultural output other than through the channel of triggering institutional changes, thereby satisfying the exclusion restriction as valid instruments. A secondary contribution of this paper is that we improve the data utilized by Lin (1992) in several important dimensions, including measurement of the HRS, mechanical and animal power, irrigation, and weather adversity. Using improved provincial-level data for the1970-1987 period, the results of panel GMM estimations reveal that the HRS had a significantly positive effect on China's agricultural growth, which was larger than that reflected in OLS estimates.

The remainder of this paper is organized as follows. Section 2 reviews the literature, and

Section 3 describes how the relevant data were improved. Section 4 discuss our empirical strategy and report the estimation results. Section 5 concludes.

2. Literature Review

There is now a sizable literature that evaluates the effects of the HRS on China's agricultural growth.[6] Wen (1993) uses national aggregated time-series data for the 1952-1989 period and finds that TFP in China's farming sector increased rapidly after 1978, which coincided with the adoption of the HRS across China. With improved data and a more sophisticated Divisia index approach, Fan and Zhang (2002) obtain similar results with smaller but respectable estimates of TFP during the reform period.[7]

McMillan et al. (1989) offer perhaps the first serious attempt to disentangle the effects of the HRS and price reforms. In their growth accounting approach aided by a simple model of peasants' optimal choice of effort, these authors find that 78 percent of productivity growth in Chinese agriculture during the 1978-1984 period can be attributed to the HRS and 22 percent to higher prices. However, the decomposition found in McMillan et al.(1989) relies upon a few simplifying assumptions, e.g., that there was no technological change and that peasants' income took a simple linear form.

Due to the substantial comovement between decollectivization and price reforms shown in Figure 1, national aggregated time-series data are not ideal for separating the effects of the HRS and market-oriented reforms. Therefore, Fan (1991) employs growth accounting through a stochastic frontier analysis on panel data for 29 provinces in 1965, 1970, and the period spanning

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[6] See Chapter 7 of Putterman (1993) for an excellent early review.
[7] Fan and Zhang (2002) also estimate TFP at the provincial level but do not compare it with provincial variations in the adoption of the HRS.

from 1975 through 1986 and finds that 26.6 percent of the growth in production was attributable to institutional changes and 15.7 percent to technological change. Similarly, Kalirajan et al. (1996) apply a varying coefficients production function to conduct growth accounting on provincial panel data during the 1970-1987 period and find that TFP growth during the pre-reform period was negative in 20 of 28 provinces and was positive in nearly all provinces during the reform periods. Zhang and Carter (1997) use county-level data to study the growth of grain output and introduce weather conditions (an aridity index) into their regression-based growth accounting; these authors find that economic reforms contributed approximately 38 percent of the growth in grain output during the 1980-1985 period. However, the use of county-level data limits their sample period to only 1980, 1985, and the period 1987-1990, which only partially covers the transition to the HRS.

However, a substantial disadvantage of the growth accounting approach is that it only measures the contribution of the HRS indirectly as a residual, which has been dubbed a measure of our ignorance (Abramovitz, 1956). As noted by Lin (1992), the customary criticisms of Dennison-type growth accounting are applicable to these studies. A more promising approach is to measure the effects of the HRS directly as an explanatory variable in regressions, although this method introduces a new problem because the HRS may then be correlated with the error term due to endogenous institutional changes. Carter and Zhong (1991) were pioneers in this approach but take a shortcut in directly using the grain yield as the dependent variable and only measuring the HRS as a dummy variable taking the value of one after 1981. In hindsight, the contribution of Lin (1992) represented a landmark breakthrough by exploiting the tremendous variation in HRS adoption across Chinese provinces and using a two-way fixed effects model to estimate China's provincial agricultural production function during the 1970-1987 period with the HRS being a key right-hand-side variable. Lin (1992) finds that the HRS accounted for approximately half of China's agricultural growth during the 1978-1984 period. It has since become a widely cited classic paper, and even critics such as Bramall (2008) and Xu (2012) acknowledge that Lin (1992) is the most influential contribution in this area. In another study using provincial panel data, Huang and Rozelle (1996) focus on the rice sector – devoting particular attention to endogenous technology adoption – and find that nearly 40 percent of the growth in rice yield during the 1978-1984 period was caused by technological change, whereas institutional reform accounted for 30 percent of the growth.

Most studies reviewed above have reached similar conclusions that decollectivization significantly contributed to China's agricultural growth. However, this apparent consensus is not without its skeptics and revisionists. Riskin (1987) argues that some of the reported increase in Chinese agricultural production is simply a statistical fluke because pre-reform output was underreported to reduce procurement quotas. Putterman (1989) reports an idiosyncratic case in Dahe Township, Hebei Province, where grain yields rose in the 1970s, but these trends were replaced by reversal and stagnation during the1980s.[8] Carolus (1992) contends that no more than 20 percent of the increase in total crop value can be attributed to HRS adoption under the most plausible set of data. Dong and Dow (1993) estimate that monitoring costs only absorb 10-20 percent of total labor time for Chinese agricultural teams but concede that this estimate is likely a lower bound of the efficiency gain under the HRS. Drawing on data from local gazetteers in Sichuan Province, Bramall (1995) finds that the growth rates of grain yields in counties that

[8] Putterman (1989) proposes a possible reconciliation by quoting Butler (1985) that "fertilizer usage declined at Dahe in 1980 when teams and brigades were given greater autonomy to determine input levels, strongly suggesting that higher authorities had driven Dahe to raise yields by economically irrational methods during the 1970s."

decollectivized late were no different from counties that decollectivized early.[9] In a monograph devoted to China, Bramall (2000) is critical of TFP estimations in the absence of precise information on labor time, organic fertilizer, and draft animals. In yet another monograph, Bramall (2008) raises doubts concerning the validity of Lin’s (1992) approach, which ignored the lagged effects of irrigation projects implemented during the collective era and did not consider weather conditions.

The most serious challenge perhaps came from Xu (2012), who focuses on the "wrong usage of HRS adoption rate" by Lin (1992). Whereas the dependent variable is the value of crop output in a year, Lin (1992) measures the HRS using its adoption rate at the end of that year, which creates the possibility that some results (agricultural growth) occurred before the cause (the HRS). To correct this specification error, Xu (2012) uses a lagged HRS variable in Lin’s (1992) two-way fixed effects model and finds that the statistical significance of the HRS disappears completely. Xu (2012) asserts that more intensive application of modern inputs and other conditions accounted for most of the increased growth. In another chapter of his dissertation, Xu (2012) constructs a socialist legacy index and demonstrates that socialist legacies – including irrigation, mechanization and human capital from the collective era – had a substantial impact on China's agricultural development after decollectivization.

In summary, a direct regression approach that exploits the tremendous cross-provincial variation in HRS adoption appears to be the most convincing to date, as exemplified by the pioneering work of Lin (1992). However, none of the previous studies address the important issue of endogenous institutional changes, which is generally acknowledged among development

[9] Counties in which "baogan dao hu", a form of HRS, was adopted by a majority of production teams in the fall of 1982 or later are classified as "late" decollectivizing counties; others are classified as "early". The selection bias discussed above may help explain certain of the puzzling observations in Sichuan counties in Bramall (2000).

economists (e.g., Acemoglu et al., 2001). The current study fills a gap in the literature precisely at this juncture in the assessment of China's decollectivization.

3. Data

The primary data source is Lin (1992), which uses provincial-level panel data from 1970 to 1987 for 28 of the 29 provinces in mainland China.[10] Nevertheless, this study extends the data from Lin(1992) in several important respects and exploits data sources that have become available since 1992.

3.1. Household Responsibility System

First among these improvements are the data for *hrs*, which is defined as the percentage of production teams converted to the HRS at the end of each year. Lacking data for *hrs* during the 1979-1980 period, Lin (1992) eliminates the 1980 observations from the data, and sets all *hrs* values to zero for 1979 because only 1.02 percent of all production teams nationwide had been converted to the HRS by 1979. This result is unsatisfactory because the national mean of 1.02 percent places no constraint on the possibility of uneven provincial distribution in HRS adoption rates.[11] Fortunately, additional data have become available in subsequent sources, such as the monograph from Chung (2000), *A Compendium of Chinese Agriculture* (1999), and in the *China Agriculture Yearbook* (1981). Nevertheless, even with these additional data sources, we continue to lack data for *hrs* for eight provinces in 1980, and these are thus set to missing.

[10] Tibet (Xizang) Autonomous Region was not included due to a lack of data. Moreover, Hainan province was spun off from Guangdong province in 1988, and Chongqing municipality was spun off from Sichuan province in 1997. In our data, Hainan was included in Guangdong, and Chongqing was included in Sichuan.

[11] According to our new data sources, three provinces had positive values of *hrs* in 1979, i.e., 0.1 for Anhui Province, 0.024 for Shandong Province, and 0.016 for Zhejiang Province.

3.2. Irrigation

One socialist legacy emphasized by revisionists is the set of irrigation projects constructed before decollectivization, which increased the amount of irrigated land. The omission of irrigation is likely to result in omitted variable bias because provinces with proportionally more irrigated land were less vulnerable to weather adversity, which hastened the adoption of the HRS (Bai and Kung, 2014).[12] To address this issue, we define *irrig* as irrigated land as a percentage of total cultivated land. The data for irrigated land are primarily taken from State Statistical Bureau (2010), supplemented by State Statistical Bureau (1999) and Ministry of Water Resources (1989).

3.3. Power

In Lin (1992), agricultural capital measured by the horsepower of tractors and draft animals is insignificant in some specifications, which indicates potential measurement errors because this measure fails to account for machines other than tractors. To construct a more comprehensive measure, we utilize data on the "total power of agricultural machinery" in kilowatts from State Statistical Bureau (2010) and convert it into horsepower. In addition, the number of draft animals at the end of each year is taken from State Statistical Bureau (1990). Following Li and Yang (2005), we compute the simple arithmetic mean of the two end-of-year numbers for a more accurate proxy for draft animals for the corresponding calendar year. Following Lin (1992), each draft animal is counted as 0.7 horsepower. Our measure of agricultural capital, denoted *power*, is simply the sum of the total power of agricultural machinery and draft animals in thousands of

[12] Bai and Kung (2014) demonstrate that "more effectively irrigated acreage constructed during collectivization had the paradoxical effect of hastening a province's (village's) exit from collective agriculture when bad weather struck, simply because, once durable public goods were in place, there was less incentive to maintain collective agriculture in the face of negative weather shocks."

horsepower.

3.4. Weather Shocks

By its nature, agricultural production is susceptible to weather shocks, particularly given agricultural conditions in China during the 1970-1987 period. For example, Zhang and Carter (1997) find that weather significantly contributes to the variability of China's grain production during the 1980-1990 period. If weather was generally favorable during the 1979-1984 period when decollectivization occurred, we might overestimate the HRS's contributions to agricultural growth. As a remedy, we use two measures of weather adversity. The first is based on disaster-affected acreage (shouzai mianji), da_{it}, which is defined as acreage suffering at least a 10 percent loss in output due to drought or flood as a percentage of total cultivated land in province *i* and year *t*. The other measure is based on disaster-ravaged acreage (chengzai mianji), dr_{it}, which is defined as acreage suffering at least a 30 percent loss in output due to drought or flood as a percentage of total cultivated land in province *i* and year *t*. The raw data are taken from Ministry of Water Resources (1989). For each province, we normalize da_{it} and dr_{it} using their deviations from the provincial means and define $wa_{it} \equiv (da_{it} - \overline{da}_i)/\overline{da}_i$, and $wr_{it} \equiv (dr_{it} - \overline{dr}_i)/\overline{dr}_i$, where $\overline{da}_i$ and $\overline{dr}_i$ are the corresponding provincial means. We then use *wa* and *wr* as two measures of weather adversity with increasing severity. Table 1 summarizes the variable definitions and data sources.

Table 1. Definitions of Key Variables and Date Sources

Variable	Observation	Definition	Data Source
y	504	Value of crop output in constant 1980 prices	L
hrs	496	Percentage of production teams converted to the HRS by the end of each year	L, C, A, Y
land	504	Cultivated land in thousand mu	L
labor	504	Labor force in the cropping sector	L
power	503	Total power of agricultural machinery plus 0.7* number of draft animals in thousands of horsepower	S60, S40
fert	504	Gross weight of chemical fertilizers consumed in thousands of tons	L
irrig	473	Ratio of irrigated land to cultivated land	S60, S50, W
mci	504	Multiple cropping index (total agricultural sown acreage divided by cultivated land)	L
ngca	504	Percentage of area devoted to nongrain crops	L
mp	504	Index of market prices relative to input prices	L
gp	504	Index of state above-quota prices relative to input prices	L
wa	504	Standardized disaster-affected acreage suffering at least a 10 percent loss of output due to drought or flood as a percentage of total cultivated land	W
wr	504	Standardized disaster-ravaged acreage suffering at least a 30 percent loss of output due to drought or flood as a percentage of total cultivated land	W

Notes:
A: *A Compendium of Chinese Agriculture* (1999)
C: Chung (2000)
L: Lin (1992)
W: Ministry of Water Resources (1989)
S40: State Statistical Bureau (1989)
S50: State Statistical Bureau (1999)
S60: State Statistical Bureau (2010)
Y: *China Agriculture Yearbook* (1981)

4. Empirical Strategy and Results

Following Lin (1992), our initial specification is to estimate a provincial-level production function using a two-way fixed effects model with year dummies:

$$\begin{aligned}\ln y_{it} = \beta_0 + \beta_1 hrs_{i,t-1} + \beta_2 \ln land_{it} + \beta_3 \ln power_{it} + \beta_4 \ln fert_{it} + \beta_5 irrig_{it} \\ + \beta_6 \ln mci_{it} + \beta_7 ngca_{it} + \beta_8 wa_{it} + \beta_9 wr_{it} + u_i + \lambda_t + \varepsilon_{it}\end{aligned}, \quad (1)$$

where subscript *i* refers to province *i*, subscript *t* refers to year *t*, u_i represents the provincial fixed effects, and λ_t represents the year fixed effects. Two national price indices (*gp* and *mp*) are not included, as they have no cross-provincial variation. Our specification differs from Lin (1992) in the following nontrivial aspects. First, we use $hrs_{i,t-1}$ instead of hrs_{it} in response to Xu's (2012) critique but use hrs_{it} and $hrsmid_{it} \equiv (hrs_{it} + hrs_{i,t-1})/2$ as robustness checks. Second, we add three new variables (*irrigation*, *wa*, and *wr*) and improve the measurement of *hrs* and *power*. Third, we use ln(*mci*) instead of *mci* because the former makes more theoretical sense (the product of cultivated land and MCI is the sown area) and has a slightly better fit with virtually identical empirical results. Finally, Lin (1992) divides output and traditional inputs (*land*, *labor*, *power*, *fert*) by the number of production teams in a province to alleviate heteroskedasticity. However, we are uncertain whether heteroskedasticity takes this specific form. Instead, we use cluster-robust standard errors that are robust to heteroskedasticity and within-panel autocorrelation of arbitrary forms.[13]

As noted above, the key variable $hrs_{i,t-1}$ in equation (1) is likely endogenous due to selection bias, measurement errors, and reverse causality. We identify the causal effect of the HRS on crop output by using lagged weather shocks ($wa_{i,t-1}$ and $wr_{i,t-1}$) as valid instruments. Lagged weather shocks are correlated with the HRS as triggers of institutional changes (Bai and Kung, 2014), and they supposedly had no effect on the current year's agricultural output other than by impacting farming institutions. Under overidentification, equation (1) can be efficiently estimated by panel GMM by first removing the provincial fixed effects u_i through the within transformation and then instrumenting $hrs_{i,t-1}$ with lagged weather shocks ($wa_{i,t-1}$ and

[13] Lin (1992) uses standard errors because cluster-robust standard errors for the fixed effects model were being developed at the time(Arellano, 1987). Our results are robust if we follow Lin (1992) and divide output and traditional inputs by the number of production teams.

$wr_{i,t-1}$).[14]

However, because $hrs_{i,t-1}$ can easily be explained by year dummies,[15] we encounter the problem of weak instruments, where the first-stage *F*-statistic testing the joint significance of $wa_{i,t-1}$ and $wr_{i,t-1}$ is only 1.32 (and we do not wish to use year dummies to explain institutional changes). Fortunately, none of these year dummies are significant at the 5 percent level, and most of these are highly insignificant. Moreover, econometricians (e.g., Arellano, 2003, p.61) suggest the exclusion of time dummies in short panels when the effects of macro-level explanatory variables are of substantive interest. Therefore, we drop the year dummies and replace them with a time trend:

$$\begin{aligned} \ln y_{it} = {} & \beta_0 + \beta_1 hrs_{i,t-1} + \beta_2 \ln land_{it} + \beta_3 \ln power_{it} + \beta_4 \ln fert_{it} + \beta_5 irrig_{it} \\ & + \beta_6 \ln mci_{it} + \beta_7 ngca_{it} + \beta_8 wa_{it} + \beta_9 wr_{it} + \beta_{10} gp_{it} + \beta_9 mp_{i,t-1} + u_i + \lambda t + \varepsilon_{it} \end{aligned}, \qquad (2)$$

With the year dummies eliminated, equation (2) now includes two national price indices (gp_{it} and $mp_{i,t-1}$) that only vary over time but not across provinces.[16] Estimation results from the panel GMM are presented in Table 2.

[14] The within transformation applies to all variables, including instrumental variables, which transforms them into deviations from their provincial means.
[15] A simple regression of lagged *hrs* on year dummies yields a R^2 value as high as 0.93.
[16] The use of these national price indices is justified because price reforms were generally implemented uniformly across Chinese provinces following directives from the central government.

Table 2. Estimation Results with Time Trend

Dependent Variable: ln(value of crop output in constant prices)

	(1) GMM	(2) GMM	(3) GMM	(4) FE
L.hrs	0.300**			0.181***
	(0.138)			(0.0339)
hrs		0.316*		
		(0.177)		
hrs_mid			0.332**	
			(0.165)	
ln(*land*)	0.735***	0.733***	0.737***	0.726***
	(0.119)	(0.118)	(0.120)	(0.114)
ln(*labor*)	0.123*	0.117	0.114	0.143***
	(0.0640)	(0.0726)	(0.0698)	(0.0477)
ln(*power*)	0.131**	0.137*	0.142**	0.0844
	(0.0622)	(0.0700)	(0.0658)	(0.0610)
ln(*fert*)	0.152***	0.162***	0.156***	0.155***
	(0.0356)	(0.0350)	(0.0357)	(0.0376)
irrig	0.383**	0.316**	0.371**	0.326*
	(0.156)	(0.147)	(0.153)	(0.170)
ln(*mci*)	0.561***	0.708**	0.653***	0.438***
	(0.140)	(0.277)	(0.205)	(0.149)
ngca	0.174	0.253	0.219	0.133
	(0.153)	(0.170)	(0.159)	(0.154)
L.mp	0.0284	0.0498	0.0564	-0.0534
	(0.0994)	(0.126)	(0.118)	(0.0406)
gp	0.0303	-0.0195	0.0130	-0.0230
	(0.0624)	(0.0529)	(0.0606)	(0.0201)
t	-0.00779	-0.0109	-0.0130	0.0114
	(0.0209)	(0.0266)	(0.0243)	(0.00893)
wa	-0.0160	-0.0122	-0.0142	-0.0183
	(0.0118)	(0.0124)	(0.0119)	(0.0123)
wr	-0.0306***	-0.0272**	-0.0286**	-0.0310*
	(0.0117)	(0.0110)	(0.0113)	(0.0153)
_cons				1.448
				(0.895)
N	442	442	434	442
R^2	0.876	0.872	0.878	0.881
First-stage F-stat	9.56	7.15	9.25	
p-value for Hansen J	0.90	0.64	0.84	

Notes: Cluster-robust standard errors in parentheses. Operator *L.* refer to a one-year lag, and *hrs_mid* = (*hrs* + *L.hrs*)/2. $^{*}p< .1$, $^{**}p< .05$, and $^{***}p< .01$.

Column (1) of Table 2 reports the results when the key explanatory variable is *L.hrs* (lagged

HRS), which is positively significant at the 5 percent level with a magnitude larger than the fixed effects estimate (0.30 versus 0.19) in Lin (1992). All conventional inputs except labor are positively significant at the 5 percent level, although labor is positively significant at the 10 percent level. The coefficients of irrigation and MCI are positively significant at the 5 percent and 1 percent levels, respectively. The coefficient of the standardized disaster-ravaged area (*wr*) is negatively significant at the 5 percent level, confirming the common belief that Chinese agriculture was susceptible to weather shocks. Moreover, the *F*-statistic testing the joint significance of *L.wa* and *L.wr* in the first-stage regression is 9.56, which is close to the rule-of-thumb value of 10, suggesting that weak instruments are not a substantial concern. The *p*-value for Hansen's *J* statistic reaches 0.90, and we can thus easily accept the null hypothesis that both instruments are exogenous.

Column (2) of Table 2 reports the results when the key explanatory variable is *hrs* (current HRS), which is positively significant at the 10 percent level. However, the first-stage *F*-statistic decreases to 7.15, which is far below 10, and we may thus have a weak instrument problem, which explains the reduced significance of *hrs*. Nevertheless, according to Xu (2012), it is inappropriate to use the year-end measure of HRS to explain agricultural output in the current year. Column (3) of Table 2 reports the results when the key explanatory variable is *hrs_mid* (a simple average of HRS and lagged HRS), which is again positively significant at the 5 percent level, and the first-stage *F*-statistic rebounds to 9.25, bolstering this estimate’s credibility. For purposes of comparison, we also report results obtained via fixed effects estimation in column (4) of Table 2, where *L.hrs* is positively significant at the 1 percent level with a magnitude close to the original estimate in Lin (1992) (0.181 versus 0.19).

Note that the time trend t is highly insignificant in all the columns in Table 2. The inclusion of the time trend also weakens the instruments (none of the first-stage F-statistics exceeds 10), simply because the time trend is a strong predictor of lagged HRS (and we do not wish to explain institutional changes by the mere passage of time).[17] Therefore, we estimate equation (2) again without the time trend, and the results are reported in Table 3.

Column (1) of Table 3 reports the results when the key explanatory variable is *L.hrs* (lagged HRS), which is now positively significant at the 1 percent level with a magnitude still larger than the original estimate in Lin (1992) (0.267 versus 0.19). All conventional inputs, in addition to irrigation and MCI, are positively significant at least at the 5 percent level. The coefficient of the standardized disaster-ravaged area (*wr*) is now negatively significant at the 1 percent level. Moreover, the F-statistic testing the joint significance of *L.wa* and *L.wr* in the first-stage regression increases to 11.24, which is greater than 10, and we are assured of the absence of weak instruments. The p-value for Hansen's J statistic reaches 0.93, and thus we can again accept the null hypothesis that both instruments are exogenous.

Columns (2) and (3) of Table 3 report the results when the key explanatory variable is *hrs* (current HRS) and *hrs_mid* (average of current and lagged HRS), respectively. The results are qualitatively similar to those in column (1). For comparison purposes, we also report the results obtained via fixed effects estimation in column (4) of Table 3.

[17] A simple regression of lagged HRS on the time trend yields R^2 as high as 0.70.

Table 3. Estimation Results without Time Trend

Dependent Variable: ln(value of crop output in constant prices)

	(1) GMM	(2) GMM	(3) GMM	(4) FE
L.hrs	0.267***			0.226***
	(0.0562)			(0.0260)
hrs		0.281***		
		(0.0670)		
hrs_mid			0.273***	
			(0.0610)	
ln(*land*)	0.737***	0.750***	0.739***	0.718***
	(0.118)	(0.117)	(0.116)	(0.117)
ln(*labor*)	0.129**	0.117**	0.125**	0.135**
	(0.0539)	(0.0567)	(0.0545)	(0.0499)
ln(*power*)	0.104**	0.0956**	0.100**	0.131***
	(0.0459)	(0.0484)	(0.0466)	(0.0376)
ln(*fert*)	0.147***	0.148***	0.147***	0.160***
	(0.0350)	(0.0381)	(0.0366)	(0.0372)
irrig	0.383**	0.339**	0.375**	0.344**
	(0.158)	(0.158)	(0.156)	(0.164)
ln(*mci*)	0.541***	0.683***	0.601***	0.451***
	(0.123)	(0.183)	(0.146)	(0.146)
ngca	0.145	0.189	0.163	0.170
	(0.146)	(0.143)	(0.143)	(0.160)
L.mp	-0.00219	0.00670	0.00279	-0.0119
	(0.0468)	(0.0491)	(0.0480)	(0.0463)
gp	0.0151	-0.0301	-0.00853	-0.00298
	(0.0327)	(0.0299)	(0.0309)	(0.0244)
wa	-0.0163	-0.0131	-0.0159	-0.0195
	(0.0113)	(0.00923)	(0.0102)	(0.0124)
wr	-0.0299***	-0.0275***	-0.0271***	-0.0294*
	(0.0108)	(0.00948)	(0.00975)	(0.0146)
_cons				1.079
				(0.937)
N	442	442	434	442
R^2	0.879	0.875	0.881	0.880
First-stage F-stat	11.24	11.53	11.59	
p-value for Hansen J	0.93	0.98	0.86	

Notes: Cluster-robust standard errors in parentheses. Operator *L*. refers to a one-year lag, and *hrs_mid* = (*hrs* + *L.hrs*)/2. $^{*}p< .1$, $^{**}p< .05$, $^{***}p< .01$.

Overall, the results in Table 2 and Table 3 show that after we resolve the endogeneity issue associated with the HRS using lagged weather shocks as valid instruments, the HRS variable has a significantly positive effect on China's agricultural growth. Consistent estimates obtained via panel GMM were larger than OLS estimates that do not consider endogeneity, which may be explained by selection bias. It is acknowledged that only poor, remote and mountainous regions were initially allowed to experiment with the HRS and that these regions might have had less to gain under the HRS – although implementing the HRS might have been necessary to ensure the survival of their populations – because they were operating close to subsistence levels. Another possible source of underestimation is the attenuation bias due to the measurement errors in the HRS variable.

5. Conclusion

Thirty years have passed since China dismantled rural collectives (1979-1984) and witnessed dramatic agricultural growth in the early 1980s. Scholars continue to debate the causal relationship between these two events. Although positive researchers embrace decollectivization as one of the most successful rural institutional changes in China, critics remain nostalgic about the bygone collective era, when massive resources could be mobilized for public goods provision, such as irrigation projects. We contribute to this debate by explicitly recognizing the endogeneity of institutional changes and use lagged weather shocks as valid instruments. With improved provincial-level data from the 1970-1987 period, results from panel GMM estimations reveal that the HRS had a significantly positive effect on China's agricultural growth, which was larger than that reflected in OLS estimates. Moreover, our results controlled for irrigation and mechanization, factors emphasized by revisionists.

At a minimum, the transition to household farming did not hamper China's agricultural growth. The fact that decollectivization initially began as a spontaneous, grassroots movement despite the political risk involved indicates that the HRS was the choice of most Chinese peasants in the pursuit of survival and prosperity. Moreover, household farming also provides Chinese peasants with enormous freedom in agricultural production, as well as the option to relocate themselves out of agriculture (Putterman, 1989), which paved the way for China's accelerated urbanization and vibrant rural industry in the form of township and village enterprises.

However, the claims of revisionists are not entirely without merit, particularly their emphasis on the role of the socialist legacy regarding public goods provision, such as irrigation projects. Although household farming supplies the correct incentives, the collective era provided China's rural sector with irrigation and other types of rural infrastructure. These two aspects might be complementary. As Bai and Kung (2014) empirically demonstrate, "more effectively irrigated acreage constructed during collectivization had the paradoxical effect of hastening a province's (village's) exit from collective agriculture when bad weather struck, simply because, once durable public goods were in place, there was less incentive to maintain collective agriculture in the face of negative weather shocks". Moreover, the alleged disadvantage of household farming in providing rural public goods has been at least partially remedied through the introduction of elections at the village level since the late 1980s (Martinez-Bravo et al., 2012).

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