Available online www.jocpr.com

Journal of Chemical and Pharmaceutical Research, 2014, 6(1):439-445

Research Article

ISSN : 0975-7384
CODEN(USA) : JCPRC5

Comparative advantage of food exports, rural-urban income disparity and agricultural employment: Empirical evidence from China

Yu Hong[1], Zhen Song[1], Jiayue Wang[1] and Hongwei Su[2*]

[1]Jilin University of Finance and Economics, Changchun, China
[2]College of Humanities and Sciences, Northeast Normal University, Changchun, China

ABSTRACT

This study assumes that 1) the agricultural employment is interactively associated with the rural-urban income disparity; and 2) the changing comparative advantage of the Chinese food exports has influences upon the income of rural residents and therefore upon the size of the agricultural employment. Using yearly data of 1984-2011, we employ vector error correction models to empirically examine the short-run and long-run Granger causal relations among the Chinese agricultural employment, rural-urban income ratio and the comparative advantage of food exports. We find that 1) rural-urban income ratio Granger causes the agricultural employment in both short-run and long-run with positive effects; 2) the agricultural employment has negative short-run and long-run effects on the rural-urban income ratio; and 3) the Chinese comparative advantage of food exports has significantly positive effect on the rural-urban income ratio, but the long-run effect is negative. We argue that although the food export facilitation can increase the rural income in short-run, the policies adversely hurt the rural-urban income equality in long-run.

Key words: Rural employment, Circular migrant labor, Rural-urban income disparity, Food exports

INTRODUCTION

China's rapid growth has provided more opportunities than challenges to the rest of the world [1]. Given that agriculture plays crucial but different roles for developing countries, issues about the Chinese agriculture have been widely scrutinized by scholars and policy makers [2-4].

The Chinese economy has been heavily dependent upon exports [5, 6]. This paper attempts to examine the effects of the agricultural exports on the Chinese rural development. Specifically, the research concern is how does the changing comparative advantage of the food/agriculture exports relate to the Chinese rural-urban income disparity and the agricultural employment.

The linkages of the variables can be complicated. On one hand, an increase in the comparative advantage implies more food exports. This would have positive impacts on the income of farm household [7] and thus narrow the rural-urban income disparity. Other agricultural policies, including the grain subsidy program [8], can also improve the income of the rural households and tend to keep the labor force at the rural origin [9]. On the other hand, some literatures document that the agricultural export elasticity of GDP is much lower than that of the non-agricultural exports [10]. Although an improved comparative advantage of food exports may benefit the agricultural labor in historical comparison, the lower sectoral productivity will lead up to a lower rural-urban income ratio and an even wider rural-urban disparity in horizontal comparison.

EXPERIMENTAL SECTION

Assumption and models

We start from an assumption of $AE=\alpha \cdot ID+\beta \cdot RCA+\mu$, where *AE* is the agricultural employment, *ID* stands for the rural-urban income disparity, and *RCA* represents the revealed comparative advantage of food exports. α and β are the coefficients to be estimated, and μ is the residual term that can not be explained. This model states a linear relationship that both rural-urban income ratio (*IR*) and the comparative advantage of food exports (*RCA*) determine agricultural employment (*AE*). A more comprehensive story tells that a lowered comparative advantage of food exports may reduce the income of farmers in relation to the urban residents, which may consequently drive more labor out of the agricultural sector [11, 12].

A problem inherent in the equation is that the causal relationship among the variables may be bi-directional. For instance, a change of agricultural employment may immediately cause a change in the factor endowments of the agricultural sector, which will generate a new agricultural labor cost and thus reshape the comparative advantage of food exports. Taking these into consideration, we allow for interactive relationship among the variables and obtain a system of equations

$$\begin{bmatrix} AE \\ ID \\ RCA \end{bmatrix} = \begin{bmatrix} \theta_1 \\ \theta_2 \\ \theta_3 \end{bmatrix} AE + \begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \end{bmatrix} ID + \begin{bmatrix} \beta_1 \\ \beta_2 \\ \beta_3 \end{bmatrix} RCA + \begin{bmatrix} \mu_1 \\ \mu_2 \\ \mu_3 \end{bmatrix} \tag{1}$$

where θs, αs and βs are the coefficients, and the restriction of $\theta_1=\alpha_2=\beta_3=0$ holds to avoid of self-explanation. We employ the first equation of the system as the co-integrating equation, which describes the long-run equilibrium relationship among the variables.

We examine the causal relationship among the variables by Granger causality test. Granger causality assumes that the past information about the cause variables helps to improve the prediction of the dependent variable [13]. This implies that dynamic models with time lags are necessary. If the time series of the three variables are stationary, we employ vector auto-regression (VAR) models for further Granger causality tests. If they are non-stationary but first-order integrated, we turn to vector error correction models (VECM) in the form of

$$\begin{bmatrix} \Delta AE_t \\ \Delta ID_t \\ \Delta RCA_t \end{bmatrix} = \begin{bmatrix} e_1 \\ e_2 \\ e_3 \end{bmatrix} ECT_{t-1} + \begin{bmatrix} \theta_{11} \Lambda\ \theta_{1k} \\ \theta_{21} \Lambda\ \theta_{2k} \\ \theta_{31} \Lambda\ \theta_{3k} \end{bmatrix} \Delta AE_{t-k} + \begin{bmatrix} \alpha_{11} \Lambda\ \alpha_{1k} \\ \alpha_{21} \Lambda\ \alpha_{2k} \\ \alpha_{31} \Lambda\ \alpha_{3k} \end{bmatrix} \Delta ID_{t-k} + \begin{bmatrix} \beta_{11} \Lambda\ \beta_{1k} \\ \beta_{11} \Lambda\ \beta_{1k} \\ \beta_{11} \Lambda\ \beta_{1k} \end{bmatrix} \Delta RCA_{t-k} + \begin{bmatrix} u_1 \\ u_2 \\ u_3 \end{bmatrix} \tag{2}$$

where Δ stands for first differentiation, the subscript t identifies time period, and the subscript k denotes the maximum order of lag. *ECT* is the error correction term, which is the residual of the co-integrating equation. It reflects the long-run equilibrium relationship among the variables.

There are three points to be noted: Firstly, the optimum lag is left unknown; secondly, there may be intercepts and linear deterministic time trends in the models; and thirdly, the co-integrating equation may also contain an intercept and (or) a time trend. Because subjective forejudging about these specifications might lead up to wrong results, we specify the optimum models by objective tests.

Measurements and Data

We measure *AE* by $AE = \dfrac{E_R - E_T}{E}$, where E_R is the total rural employment, E_T is the employment in township and village enterprises, and E stands for the total employment. So the *AE* indicator captures the percentage of agricultural employment. Its value ranges from 0 to unity.

We obtain rural-urban income disparity by $ID = \dfrac{NI_R}{DI_U}$, where NI_R refers to net personal income per capita of rural residents, and NI_U represents disposable personal income per capita of urban residents. An increase in *ID* ratio implies a reduction of the rural-urban income disparity.

$RCA_{ik} = \frac{X_{ik} / X_i}{X_{wk} / X_w}$ is an indicator that explores the comparative advantage revealed in exports. X refers to exports. The subscript of k is for the k-th product, that of i stands for the i-th country, and that of w denotes the world. Thus X_{ik} and X_{wk} refer to the exports of the k-th product in country i and in the world respectively, X_i and X_w are the total exports of country i and the world [14]. The *RCA* indicator compares the share of country *i*'s exports of the k-th product with that of the world. The higher the value of the indicator, the more specialized the i-th country is in product k.

We obtain the annual data of China's total employment (E), rural employment (E_R), the employment in township and village enterprises (E_T), and the personal income per capita of rural and urban residents (NI_R and DI_U) from "China Statistics Yearbook" of the National Bureau of Statistics of China. Trade amount data are compiled from United Nation's COMTRADE database. We employ the second revision of Standard International Trade Classification (SITC Rev.2) to identify agricultural products. The category of "food and live animals chiefly for food", which is under 1-digit SITC code, is selected as the definition of food exports. Taking the availability of data into consideration, the sample period of our data is from the year of 1984 to 2011.

CO-INTEGRATION ANALYSIS

ADF Unit Root Test

The time series may be either stationary or non-stationary. Because ordinary least square regression of non-stationary time series may suffer from spurious regression, we conduct augmented Dicky-Fuller (ADF) unit root tests to examine the stationary. Tab. 1 reports the test results.

Tab. 1 Augmented Dicky-Fuller Unit Root Test Results

Level Series			First Differenced Series		
Variable	Test Type[1]	ADF	Variable	Test Type	ADF
AE_t	N,N,0	-0.641 (0.968)[2]	ΔAE_t	C,N,0	-3.582 (0.014)
ID_t	C,N,4	-1.331 (0.856)	ΔID_t	N,N,0	-2.048 (0.041)
RCA_t	C,N,1	-3.465 (0.065)	ΔRCA_t	N,N,0	-5.135 (0.000)

1. The test types of "C, T, p" are the model specification, which stands for intercept, trend, and the lag length respectively. The symbol of "N" is used when there is no an intercept or a trend.

2. Probabilities for the ADF values are in parentheses.

All the level time series are non-stationary. However, the first differences are all stationary at 0.05 confidence level, implying that OLS approach is vulnerable to spurious regression. Given that all the time series are first-order integrated, co-integration tests are necessary to identify whether there is a long-run equilibrium co-integrating relation among these variables.

Johansen Co-integration Test

We conduct Johansen co-integration tests based on vector error correction models by three steps.

The first step of Johansen co-integration is to select the optimum lag interval. We addressed this problem by estimating the unrestricted vector autoregressive model (VAR) with a maximum lag of 4. We determined the optimum VAR lag length "p" by Schwarz criterion (SC). The lag interval of Johansen co-integration test as well as the vector error correction models (VECM) is given by "1, p-1". In this study, we had p=4 and the optimum lag interval is 1 to 3.

The second step is to select the optimum test type of co-integration. By estimating the five possible VECMs and according to the information criteria of AIC and SIC, we confirmed that the optimum co-integration test type (as well as the optimum model specification of VECM) has linear deterministic trends in both the co-integrating equation and the VAR.

For the third step, we conducted Johansen co-integration test. Tab. 2 presents the summary of the test results. Both Trace test and Max-eigenvalue test indicate that there is only one co-integrating relation with model V as the optimum test type. This enables us to perform short-run and long-run Granger causality tests.

Tab. 2 Augmented Dicky-Fuller Unit Root Test Results

Test Type[1]	I	II	III	IV	V
Trace	3[2]	3	3	3	1
Max-eigen	3	3	3	3	1
AIC	-14.486	-14.496	-15.290	-15.843	-16.544*[3]
SC	-12.857	-12.818	-13.513	-14.016	-14.618*

1. There are five possible test typs: ① model I has neither intercept nor trend in co-integrating equation and VAR; ② model II has an intercept in co-integrating equation; ③ model III has intercepts in both of co-integrating equation and VAR; ④ model IV has intercept and trend in co-integrating equation, but has only an intercept in VAR; ⑤model V has quadratic deterministic trends in co-integrating equation and VAR.

2. The figure indicates the number of co-integrating relations as identified by Trace test and Max-eigenvalue test.

3. * indicates the optimum test type as selected by AIC and SC criteria.

GRANGER CAUSALITY TESTS

When the inclusion of the lags of an independent variable helps to improve the explanatory power of the model, we say that the independent variable Granger causes the dependent variable. Hong and Su (2012) argues that the length of physical time is not the criterion of distinguishing short-run and long-run causality relations. The short-run causality assumes "other conditions keeping unchanged", while the long-run causality allows for the interactive feedbacks of effects through the long-run equilibrium co-integrating relationship [15]. In this study, we perform both short-run and long-run Granger causality tests because they can be dramatically different.

***Short-run Granger causality test*:** This study conducts short-run Granger non-causality tests by employing VECM-based block exogeneity Wald test. If an exclusion of the three lagged differenced terms of an independent variable significantly decreases the explanatory power of the VECM, Granger causality runs from the independent variable to the dependent variable. For example, in the VECMs of equation system (2), we implement the Wald restriction of $\alpha_{11}=\alpha_{12}=\ldots=\alpha_{1k}=0$ to test whether there is a short-run Granger causality running from ΔID_{t-k} to ΔAE_t, and implement the restriction of $\theta_{21}=\theta_{22}=\ldots=\theta_{2k}=0$ to test whether ΔAE_{t-k} Granger cause ΔID_t. When identifying a significant Granger causal relationship by χ^2 statistics, we sum the coefficients of the three lag terms of the independent variable to plot the short-run effects. Tab. 3 reports the empirical results.

Tab. 3 Short-run Granger Causality Test Results

Independent variable:[1]	$\triangle AE_t$	$\triangle ID_t$	$\triangle RCA_t$
$\triangle AE_{t-1}$, $\triangle AE_{t-2}$, $\triangle AE_{t-3}$	——	16.291 (0.001)	6.865 (0.076)
$\triangle ID_{t-1}$, $\triangle ID_{t-2}$, $\triangle ID_{t-3}$	13.725 (0.003)[2]	——	7.213 (0.065)
$\triangle RCA_{t-1}$, $\triangle RCA_{t-2}$, $\triangle RCA_{t-3}$	6.114 (0.106)	10.079 (0.018)	——

1. The independent variables of the vector error correction models are in the first line; the dependent variables are in the first column.
2. The values in the cell are $\chi 2$ statistics with probabilities in parentheses.

Firstly, rural-urban income disparity (ΔID) Granger causes the agricultural employment (ΔAE) at 0.05 level in short-run. The short-run effect is positive, implying that an increase in the personal income rural residents can attract more Chinese circular migrant workers back to the rural areas. Changes in the revealed comparative advantage of food exports (ΔRCA), however, have no significant short-run effect on the Chinese agricultural employment.

Secondly, in short-run, both ΔAE and ΔRCA Granger cause rural-urban income disparity (ΔID). Specifically, an increase in the agricultural employment (ΔAE) will reduce the relative personal income of the Chinese rural residents (ΔID) and thus widen the rural-urban income disparity. An increase in the revealed comparative advantage of Chinese food exports (ΔRCA), however, has a positive effect on the personal income of rural residents. This implies that the improvement of international competitiveness of Chinese food exports may narrow the rural-urban income disparity.

Finally, neither ΔAE nor ΔID exerts significant effect on ΔRCA at 0.05 level. We thus argue that the comparative advantage of Chinese food exports is driven by technical changes instead of structural changes in short-run.

***Long-run Granger causality test*:** We perform VECM based combined Wald tests to examine long-run Granger causalities. The approach is featured by restricting the coefficients of the lagged differences of the explanatory variable and the coefficient of the error correction term to be zero. Specifically, as shown in the VECM system, we implement the restriction of $e_1=\alpha_{11}=\alpha_{12}=\ldots=\alpha_{1k}=0$ to test whether ΔID_{t-k} Granger cause ΔAE_t in the long-run, and we implement the restriction of $e_2=\theta_{21}=\theta_{22}=\ldots=\theta_{2k}=0$ to test whether the long-run Granger causality runs from ΔAE_{t-k} to ΔID_t.

The restriction on the coefficient of the error correction term (ECT_{t-1}) is necessary because it contains the past information about the long-run equilibrium relationship among the three variables. An impulse in the concerned independent variable will cause a change in the error correction term, which has consequent influences upon the other two variables. In other words, the combined Wald test allows "other conditions" to change. By doing this, the Granger causality test is a long-run test. Tab.4 presents the long-run Granger causality test results.

Tab. 4 Long-run Granger Causality Test Results

Independent variable:	$\triangle AE_t$	$\triangle ID_t$	$\triangle RCA_t$
$ECT_{t-1}, \triangle AE_{t-1}, \triangle AE_{t-2}, \triangle AE_{t-3}$	——	9.781 (0.001)	1.719 (0.216)
$ECT_{t-1}, \triangle ID_{t-1}, \triangle ID_{t-2}, \triangle ID_{t-3}$	3.715 (0.038)[1]	——	2.147 (0.143)
$ECT_{t-1}, \triangle RCA_{t-1}, \triangle RCA_{t-2}, \triangle RCA_{t-3}$	2.052 (0.156)	9.548 (0.001)	——

1. The values in the cell are F-statistics with probabilities in parentheses.

The directions of long-run Granger causalities are the same with those of the short-run Granger causalities. The positive/negative signs of the long-run effects, however, may be different from those of the short-run effects. We thus employ the convergence values of VECM-based generalized impulse-response functions to judge the long-run effects. Fig. 1, Fig. 2 and Fig. 3 show the locus of the function values when the long-run Granger causal relationship is significant at 0.05 level. Each of the three functions essentially converges to a constant at 50 periods after the impulses occur.

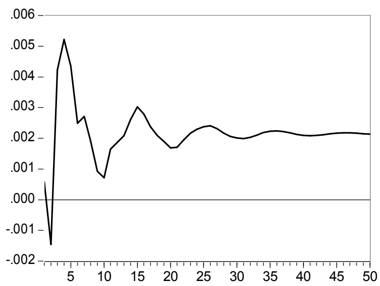

Fig. 1 Resdponses of *AE* to *ID*

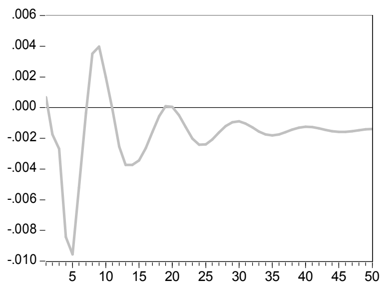

Fig. 2 Resdponses of *ID* to *AE*

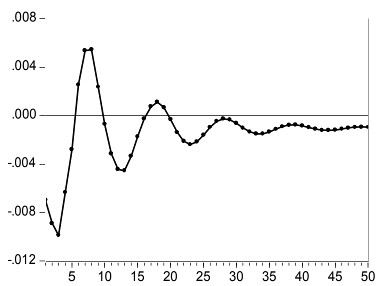

Fig. 3 Resdponses of *ID* to *RCA*

Firstly, the rural-urban income disparity (*ID*) has positive long-run effect on the Chinese agricultural employment (*AE*). A decrease in the relative personal income of rural residents may drive more rural labor forces out from the agricultural sector, both in short-run and in long-run.

Secondly, the agricultural employment (*AE*) has negative effects on the rural-urban income ratio (*ID*). The more labor forces are there in the agricultural sector, the wider the rural-urban disparity. This long-run effect is similar to that of the short-run effect in terms of the directions of Granger causality and the signs of the effects.

Thirdly, the revealed comparative advantage of Chinese food exports (*RCA*) has negative long-run effect upon the rural-urban income ratio (*ID*), which is opposite to the short-run effect. The finding is very interesting: although improving the comparative advantage of food exports can increase the income per capita of the rural residents and thus narrow the rural-urban income disparity in short-run, in long-run, it reduces the rural income and enlarges the rural-urban income disparity.

Finally, the revealed comparative advantage of Chinese food exports (*RCA*) is once again exogenous. The structural changes which are related to the abundance of labor have neither significant short-run nor long-run effects on *RCA*.

CONCLUSION

This study develops a dynamic econometric model to investigate the relationship among the Chinese ratio of agricultural employment, rural-urban income ratio and the comparative advantage of food/agricultural exports. Using Chinese annual data ranging from 1984 to 2011, we performed vector error correction model-based short-run and long-run Granger causality tests. We sum up the coefficients of the lagged differences of an explanatory variable to judge the positive/negative signs of its short-run effects upon different dependent variables. The signs of long-run effects are examined by the convergence values of the generalized impulse-response functions based on the vector error correction model. Empirical evidences show that:

1) Rural-urban income disparity Granger causes the Chinese agricultural employment in both short-run and long-run with positive effects. Considering that both of the indicators have shown downward trends, the consistency of the short-run and the long-run effects implies that the descending rural-urban income ratio has played a crucial role to drive the rural labor force out from the agricultural sector. Another implication is that the Chinese agricultural and rural policies, which target at improving the income of the Chinese farmers, may effectively retain the rural labor force in or attract them back to the agricultural sector. These policies can thus serve as strategic tools to control the mobility of labor across agricultural and industrial sectors, and thus mitigate the pressure of unexpected urban unemployment when recession occurs.

2) The ratio of agricultural employment in turn Granger causes the Chinese rural-urban income ratio. The short-run and the long-run effects, however, are negative. These findings are of little surprise because before the Lewis turning point, there should be surplus rural labor and thus reduce the marginal productivity of the agricultural labor in relation to other sectors. We would rather state that the decreasing agricultural employment have increased the marginal productivity of the agricultural labor, which have improved the relative income of rural residents [16, 17]. Additionally, the income of the increasing circular migrant workers from non-agricultural sectors can also contribute to the income improvement of the Chinese rural residents. We consequently conclude that an effective pathway to the reduction of the Chinese rural poverty is to facilitate the mobility of rural labor force and the development of non-agricultural sectors.

3) There are both short-run and long-run Granger causal relationships running from the comparative advantage of the Chinese food exports to the rural-urban income ratio. The effects, however, are inconsistent. The comparative advantage of food exports has positive short-run effect on the rural-urban income ratio, while the long-run effect is negative. The declining comparative advantage of the Chinese food exports in recent years may reduce the rural personal income in relation to that of the urban residents in short-run. But in long-run, this changing trade pattern of the food exports has facilitated the development of non-agricultural economy in the Chinese rural regions, and has driven more rural labor force out from the low-productivity agricultural sector to the more economically efficient industrial sectors. In other words, the Chinese rural residents have actually benefited from the weakening comparative advantage of food exports. The agricultural support policies, which aim at improving the international competitiveness of the agricultural sector by means of export facilitation, will adversely hurt the rural-urban income equality in long-run.

Acknowledgements
This study is financially supported by grant 2012BS22 from Jilin Provincial Social Science Fund and grant 2013GXS4D112 from National Soft Science Project of China.

REFERENCES

[1] J Huang; J Yang; S Rozelle. *Agric. Econ*. **2010**, 41 (s1): 47–55.
[2] AD Brauw; S Rozelle. *China. Econ. Rev*. **2008**, 19 (2): 320-335.
[3] A Valdés; W Foster. *World. Dev*. **2010**, 38 (10): 1362-1374.
[4] K Cao; JA Birchenall. J. Dev. Econ. **2013**, 104: 165-180.
[5] L Shan; C Cui; L Wu. *J. Food. Agric. Eviron*. **2011**, 9 (3-4): 923-925.
[6] S Gao; X Tong; L Wu. *J. Food. Agric. Eviron*. **2012**, 10 (1): 919-922.
[7] J Huang; Y Jun; Z Xu; S Rozelle; N Li. *China. Econ. Rev*. **2007**, 18 (3): 244-265.
[8] W Yu; HG Jensen. *J. Agric. Econ*. **2010**, 61 (2): 343–368.
[9] L Meng. *China. Econ. Rev*. **2012**, 23 (3): 729-741.
[10] AI Sanjuán-López; PJ Dawson. *J. Agric. Econ*. **2010**, 61 (3): 565–583.

[11] F Hu; Z Xu; Y Chen. *China. Econ. Rev*. **2011**, 22 (1): 64-74.
[12] S Rozelle; J Huang; L Zhang. *China. Econ. Rev*. **2011**, 13 (4): 345-353.
[13] CWJ Granger. *Econometrica*. **1969**, 37 (3): 424-438.
[14] B Balassa. *Manchester. School. Econ. Social. Stud*. **1965**, 33 (2): 99–123.
[15] Y Hong; H Su. *AASRI Procedia*. **2012**, 2: 20-24.
[16] S Haggblade; P Hazell; T Reardon. *World Dev*. **2010**, 38 (10): 1429-1441.
[17] C Li; J Gibson. *World. Dev*. **2013**, 47: 16-29.