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Analysis of the Impact of Fiscal Revenue and Expenditure on China's Grain Production Using Panel Double-Kink Regression Model

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Abstract

This study examines whether the effects of fiscal revenue and expenditure on grain production in China are nonlinear. Using a balanced panel of 31 provinces from 2007 to 2021, we analyze major revenue-side and expenditure-side fiscal instruments, including the cultivated land occupation tax, value-added tax, agricultural insurance subsidies, agricultural loan interest subsidies, rural minimum living security subsidies, education expenditure, and transportation infrastructure expenditure. To identify regime-dependent marginal effects, we employ panel kink and double-kink regression models with endogenously estimated kink points. The results show that fiscal effects are intensity-dependent rather than constant. The cultivated land occupation tax exhibits a kinked relationship with grain production, with a stronger land-protection effect beyond a certain policy intensity. Agricultural insurance subsidies display a double-kink pattern, with the strongest positive effect concentrated in an intermediate subsidy range. Rural minimum living security subsidies are positively associated with grain production at lower levels, but their marginal effect weakens and may become negative after the kink point. Overall, the findings suggest that fiscal policy effectiveness depends not only on policy direction but also on policy intensity, implying that grain-support policies should be calibrated with attention to nonlinear threshold effects.

Keywords: grain production; fiscal policy; nonlinear effects; panel kink regression; double-kink regression; food security

1. Introduction

Global food systems face mounting pressure from population growth, urbanization, climate change, and uneven regional development. The world population reached approximately 8.2 billion in 2024, and nearly 58% of people lived in urban areas [1]. At the same time, climate change is increasing agricultural vulnerability through more frequent and more severe extreme weather events [2], while 2.33 billion people, or 28.9% of the world's population, experienced moderate or severe food insecurity in 2023 [3]. In this context, maintaining stable grain production is not only an agricultural issue, but also a major concern for food security, price stability, and macroeconomic resilience.

This question is especially important in China. China's population reached 1.408 billion at the end of 2024 [4], while grain production remains highly concentrated geographically: the 13 major grain-producing provinces have accounted for roughly three-quarters of national grain output in recent years [5]. China's total grain output reached a record 695.41 million tons in 2023 [6]. Because

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grain security is closely linked to rural livelihoods, social stability, and national development, understanding how fiscal policy shapes grain production is of clear economic and policy significance.

Fiscal policy may affect grain production through several channels. On the revenue side, instruments such as the cultivated land occupation tax and value-added tax (VAT) may alter land conversion costs, production incentives, and the broader price environment faced by agricultural producers. On the expenditure side, agricultural insurance premium subsidies, agricultural loan interest subsidies, rural minimum living security subsidies, education expenditure, and transportation infrastructure expenditure may influence output by reducing production risk, easing credit constraints, improving human capital, and strengthening market access. These mechanisms suggest that fiscal effects are unlikely to be constant across policy intensity ranges. A policy instrument may have limited effects at low intensity, stronger effects after a critical level is reached, or diminishing effects once marginal returns decline. The relationship between fiscal policy and grain production may therefore be nonlinear.

Existing research does not provide a single clear expectation. International evidence suggests that agricultural support can stabilize production and improve productivity, but it may also generate distortions depending on policy design, targeting, and implementation [7]. In China, some studies report limited production responses to tax reform; for example, the abolition of the agricultural tax had statistically insignificant effects on household production inputs and outputs [8]. By contrast, other studies find substantial returns to public investment in education, agricultural research and extension, and rural infrastructure for agricultural growth and poverty reduction [9,10]. Broader evidence also shows that transportation infrastructure can significantly improve market access and economic outcomes [11]. However, much of the existing empirical literature relies on linear specifications or examines individual fiscal instruments in isolation. As a result, evidence remains limited on whether fiscal revenue and expenditure affect grain production in a regime-dependent manner, and whether different fiscal instruments exhibit distinct nonlinear patterns within a unified provincial framework.

To address this gap, this paper examines whether the effects of fiscal revenue and expenditure on grain production are nonlinear in China, and more specifically, whether their marginal effects change once policy intensity crosses endogenously determined kink points. Using a balanced panel of 31 Chinese provinces from 2007 to 2021, we analyze major fiscal instruments on both the revenue side and the expenditure side, including the cultivated land occupation tax, VAT, agricultural insurance premium subsidies, agricultural loan interest subsidies, rural minimum living security subsidies, education expenditure, and transportation infrastructure expenditure. To identify possible regime-dependent slope changes, we employ panel kink and double-kink regression models with endogenously estimated kink points [12,13], building on the broader panel threshold literature [14] and drawing on dynamic or endogeneity-robust extensions where appropriate [15]. We further control for key production factors, include province and year fixed effects, use ridge regularization to reduce multicollinearity among fiscal covariates [16], and conduct diagnostic, robustness, and endogeneity checks, including IV- and GMM-based approaches [17,18].

The results show that fiscal policy affects grain production in an intensity-dependent and nonlinear manner. The cultivated land occupation tax exhibits a kinked relationship with grain output, suggesting that its land-protection effect becomes more evident only after policy intensity exceeds a certain level. VAT is predominantly negatively associated with grain production before the estimated threshold. Agricultural insurance premium subsidies display a double-kink pattern, with the strongest positive marginal effects concentrated within an intermediate range. Rural minimum living security subsidies are positively associated with grain production at relatively low levels, but their effect weakens and may reverse after the estimated kink point is crossed. Education expenditure and transportation infrastructure expenditure also show heterogeneous nonlinear effects. Overall, these findings indicate that the effectiveness of fiscal policy depends not only on policy direction, but also on policy intensity.

This paper contributes to the literature in three ways. First, it jointly evaluates revenue-side and expenditure-side fiscal instruments within a unified provincial framework, allowing systematic comparison across policy tools. Second, by explicitly modeling kink and double-kink effects, it moves beyond conventional linear specifications and provides new evidence on threshold-dependent fiscal impacts on grain production. Third, it offers long-span provincial evidence for China and derives policy implications directly relevant to grain security: simply expanding fiscal support does not necessarily improve grain production, and policy calibration across different intensity ranges matters.

The remainder of this paper is organized as follows. Section 2 reviews the literature and develops the research hypotheses. Section 3 describes the materials and methods. Section 4 presents the empirical results and robustness checks. Section 5 discusses the policy implications. Section 6 concludes.

2. Literature Review

2.1. Conceptual Clarification and Research Positioning

In this study, agricultural fiscal policy is defined in a relatively narrow and empirically observable way. On the revenue side, we focus on explicit taxes, namely the Cultivated Land Occupation Tax and the value-added tax (VAT), rather than broader implicit taxation generated by exchange-rate distortions, trade restrictions, procurement policies, or administered prices. This choice is mainly methodological. Distortions to agricultural incentives are often multidimensional and difficult to measure consistently across regions and over time [19]. By contrast, explicit taxes have clearer statutory bases and are more directly observable at the provincial level, making them more suitable for identifying regime-dependent changes in marginal effects.

On the expenditure side, we consider producer-relevant public spending, including agricultural insurance premium subsidies, agricultural loan interest subsidies, rural minimum living security subsidies, education expenditure, and transportation infrastructure expenditure. These fiscal instruments may influence grain production through several channels, including production costs, liquidity constraints, risk sharing, labor allocation, human capital accumulation, and market accessibility. Existing studies have emphasized the importance of public investment, insurance, and infrastructure for agricultural performance [9,20,21]. However, these mechanisms are unlikely to operate with constant marginal effects. At low policy intensity, transmission may be weak; at intermediate levels, the effects may become stronger as key constraints are relaxed; and at high intensity, diminishing returns, crowding-out, or incentive distortions may emerge.

A growing international literature provides reasons to expect such nonlinearities. Fertilizer subsidies, for example, may stimulate input use but also crowd out commercial demand under some conditions [22]. Insurance support may reduce production risk and ease credit constraints, yet excessive subsidization can weaken efficiency through moral hazard or behavioral distortions [20,21]. Infrastructure investment can improve market access and agricultural productivity, but the marginal benefits of additional investment may decline once basic connectivity has been established [9,11]. Taken together, these findings suggest that the relationship between fiscal policy and grain production is unlikely to be purely linear.

In the Chinese context, the institutional background is shaped in part by the Cultivated Land Occupation Tax Law and the VAT Law [23,24]. On the expenditure side, recent studies suggest heterogeneous effects of grain subsidies, agricultural insurance, and rural social assistance. Grain subsidies can relax liquidity constraints and increase grain-sown areas [25], new agricultural subsidy policies can expand grain acreage [26], agricultural insurance is positively associated with grain production scale [27], and Rural Dibao may shift household labor toward non-agricultural employment [28]. These findings imply that fiscal instruments may affect grain production in a piecewise rather than purely linear manner, providing a strong rationale for threshold- or kink-type models.

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2.2. Fiscal Revenue and Grain Production

Agricultural taxation includes both explicit and implicit components. Explicit taxes refer to directly legislated tax instruments, such as land-related taxes and indirect taxes like VAT. Implicit taxation refers to policy-induced price wedges created by non-tax interventions, such as trade barriers, exchange-rate distortions, procurement rules, or state pricing systems. Because these distortions are often broad, indirect, and difficult to attribute to a single policy source [19], this paper focuses only on explicit taxation, which is more directly observable and empirically tractable at the provincial level.

A broad literature shows that taxation and policy distortions can influence agricultural production, productivity, and structural transformation, but the direction and magnitude of these effects depend heavily on policy design and institutional context. Policy distortions may generate substantial productivity losses through misallocation [29], slow structural transformation and broader economic growth [30], and contribute to persistent agricultural productivity gaps across countries [31]. These studies do not support a single uniform relationship between taxation and agricultural output. Rather, they imply that tax-related effects are likely to be nonlinear and context dependent.

Micro and structural evidence further supports this interpretation. Reductions in agricultural produce cess can raise farm income and production intensity, although the magnitude of the gains differs across farm types and regions [32]. Tax revenue may also have indirect productive effects if it is transformed into infrastructure that improves agricultural productivity [33]. In addition, some taxes may alter production behavior through channels more complex than a simple cost burden [34]. These findings point to the possibility that the overall effect of taxation depends not only on the tax itself, but also on how the fiscal system interacts with land allocation, public investment, and local incentives.

The Chinese experience illustrates this complexity. The tax-for-fee reform did not produce uniform effects on irrigation costs, water choices, or agricultural outcomes [35]. After the abolition of the agricultural tax, the direct effects on household inputs and outputs appear to have been limited [8]. This suggests that not all tax reforms automatically produce strong production responses.

However, the decline of the agricultural tax does not imply that revenue-side fiscal policy has become irrelevant for grain production. Two explicit tax instruments remain especially important. The Cultivated Land Occupation Tax affects the cost of converting cultivated land to non-agricultural use and may therefore influence grain production through the land-allocation margin [23]. VAT may influence agricultural production more indirectly through processing, circulation, and marketing channels, thereby affecting relative prices, incentives, and the broader cost environment [24]. These mechanisms suggest that the effects of explicit taxation on grain output may be non-monotonic and regime dependent, especially when land pressure, fiscal dependence, and policy intensity vary across provinces.

2.3. Fiscal Expenditure and Grain Production

Compared with revenue instruments, expenditure instruments operate through a wider range of channels and may therefore generate even more heterogeneous effects on grain production. Internationally, agricultural support is often classified using the OECD's Producer Support Estimate and related indicators, which distinguish support linked to outputs, inputs, and broader public goods [36]. In the present study, expenditure measures mainly correspond to risk-sharing support, credit-related support, income transfers, and public-good-oriented expenditures.

The existing literature suggests that these expenditure tools can promote grain production, but their effectiveness depends on policy intensity, incidence, and institutional conditions. Agricultural support may affect production through wealth and insurance effects under uncertainty [37]. Insurance subsidies can reduce risk and liquidity constraints and encourage investment, but their marginal effects may weaken at high subsidy intensity because of moral hazard or inefficient allocation [21]. Part of the benefit from subsidies may also be capitalized into land rents, generating heterogeneous incidence across producers and regions [38]. These findings imply that expenditure-based policies may not have a constant marginal effect on output.

Public expenditure on transport, agricultural research, extension, and education may increase productivity and improve market access, yet the marginal returns to such investment may also diminish once basic services and infrastructure are already in place. Public investment played an important role in rural growth and poverty reduction in China [9], and transportation infrastructure can substantially improve market access and economic performance more broadly [11]. This supports the view that infrastructure and public-service expenditure can raise grain production, but the gains may depend on initial conditions and local development levels.

The Chinese literature broadly supports the importance of fiscal expenditure for agricultural development, but the empirical evidence is far from uniform. OECD reports document the expansion of agricultural support in China since the early 2000s [7,36]. Grain subsidies can ease liquidity constraints and increase grain-sown areas [25]. China's newer agricultural subsidy policies can increase grain acreage [26], while grain subsidies may also interact with off-farm labor supply and farmland leasing decisions [39]. More recent evidence suggests that agricultural insurance positively affects grain production scale [27], whereas Rural Dibao may shift labor allocation toward non-agricultural employment [28]. These channels indicate that expenditure-based fiscal instruments may generate nonlinear net effects on grain output, depending on the level and structure of support.

Overall, the literature suggests that fiscal expenditure is unlikely to affect grain production through a single constant slope. Instead, its effects may be weak at low intensity, stronger at intermediate intensity, and diminishing or even offsetting at high intensity. This provides a strong rationale for using kink or threshold models rather than relying only on linear specifications.

2.4. Literature on Kink Regression and Its Relevance to This Study

To model possible regime-dependent marginal effects, this paper adopts a panel kink regression framework. A kink model differs from a standard threshold model in that it imposes continuity in the conditional mean at the breakpoint while allowing the slope to change across regimes. Such models are well suited to settings in which the marginal effect of a variable changes once the running variable crosses an unknown threshold [12]. This feature is highly relevant here, because fiscal policy intensity is more likely to alter the marginal effect of a fiscal instrument than to create a discontinuous jump in grain production at an exact policy value.

The broader threshold literature provides the econometric foundation for this approach. Threshold estimation and inference with unknown breakpoints were developed in the classic panel threshold literature [14,40]. In the regression kink design literature, slope changes have been shown to identify economically meaningful marginal effects under suitable assumptions [41]. Subsequent studies improved inference by addressing problems such as over-rejection and curvature bias in local polynomial estimation [42–44]. Although much of this literature was developed for quasi-experimental designs, it is still informative for the present study because it clarifies why changes in slope may capture economically important nonlinear responses.

More recent work has extended kink analysis to heterogeneous settings. Quantile regression kink designs, for example, allow slope changes to vary across different parts of the outcome distribution [45]. For panel data, fixed-effects kink models with unknown thresholds have been developed [13], and the threshold itself can also be modeled as depending on covariates [46]. Related dynamic threshold models further address endogeneity and persistence in nonlinear panel settings [15]. Taken together, this literature indicates that a panel kink framework is an appropriate tool for identifying whether fiscal policy exerts different marginal effects on grain production across low-, medium-, and high-intensity regimes.

2.5. Research Gap and Hypotheses

Despite the broad literature on agricultural taxation, subsidies, public expenditure, and productivity, three main gaps remain. First, much of the existing literature studies single fiscal

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instruments in isolation, rather than comparing multiple revenue-side and expenditure-side tools within one unified empirical framework. Second, many studies rely on linear specifications, which may conceal regime-dependent responses and produce misleading average effects. Third, in the Chinese provincial context, relatively few studies directly examine whether the effects of fiscal policy on grain production vary systematically with policy intensity.

To address these gaps, this paper analyzes multiple fiscal revenue and expenditure instruments together and explicitly tests for nonlinear slope changes using panel kink and double-kink models. Based on the foregoing discussion, three hypotheses are proposed.

H1. The relationship between fiscal instruments and grain output is nonlinear. For at least some fiscal variables, the marginal effect on grain production differs across policy-intensity regimes rather than following a single constant slope.

H2. For some expenditure-based fiscal instruments, the positive marginal effect on grain output is likely to be stronger in an intermediate intensity range than at very low or very high levels, reflecting weak transmission at low intensity and diminishing returns or incentive distortions at high intensity.

H3. Revenue-side and expenditure-side fiscal instruments may exhibit different nonlinear patterns because they operate through distinct mechanisms, including land-use incentives, price transmission, liquidity relief, risk sharing, labor allocation, human capital accumulation, and market access.

3. Materials and Methods

3.1. Data Description and Tests of Panel Data

This study uses a balanced panel dataset covering 31 provinces in China from 2007 to 2021 to examine the relationship between fiscal policy and grain production. The dependent variable is provincial grain output. The explanatory variables include seven fiscal variables and three control variables. The fiscal variables consist of two revenue-side variables, namely the cultivated land occupation tax and value-added tax (VAT), and five expenditure-side variables, namely agricultural insurance subsidies, agricultural loan interest subsidies, rural minimum living security subsidies, education expenditure, and transportation infrastructure expenditure. The control variables are agricultural labor, arable land area, and fertilizer use. Detailed definitions and data sources are reported in Table 1.

Table 1. Variable description.

Variable type	Variable name	Description	Source
Dependent variable	Grain production output (Y)	Provincial grain output in China from 2007 to 2021, as reported in official statistics. Unit: 100,000 tons.	China Agricultural Statistical Yearbook
Independent variables (Tax variables)	Cultivated land occupation tax (FT)	Revenue from the cultivated land occupation tax, which is levied on the occupation of cultivated land for non-agricultural construction. Unit: 100 million CNY.	China Tax Yearbook
	Value-added tax (VAT)	Revenue from value-added tax. Although self-produced primary agricultural products are generally exempt at the farm level, VAT may still affect grain production through agricultural processing, circulation, and related value chains. Unit: 1 billion CNY.	China Tax Yearbook

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Independent variables (Expenditure variables)	Agricultural insurance subsidies (IS)	Government expenditure on premium subsidies for agricultural insurance. Unit: 10 million CNY.	Ministry of Finance of China
	Agricultural loan interest rate subsidies (LS)	Government expenditure on interest subsidies for agricultural loans. Unit: 100 million CNY.	
	Rural minimum living security subsidy (RS)	Government expenditure on the rural minimum living security program for rural residents whose income falls below the locally defined minimum living standard. Unit: 100 million CNY.	
	Education expenditure (EE)	Government expenditure on education, including recurrent expenditure and capital investment in the education sector. Unit: 100 million CNY.	
	Transportation infrastructure expenditure (TE)	Government expenditure on transportation infrastructure, including spending on road, railway, waterway, and postal services. Unit: 100 million CNY.	
Independent variables (Control variables)	Labor force (L)	Number of people engaged in agricultural production. Unit: 10,000 persons.	China Agricultural Statistical Yearbook
	Land area (A)	Area sown to grain crops in each province. Unit: 1,000 hectares.	
	Fertilizer used (F)	Total amount of chemical fertilizer applied in agriculture in each province per year. Unit: 10,000 tons.	

Based on the variable definitions and data sources presented in Tables 1, Table 2 reports the descriptive statistics of all variables used in the empirical analysis.

Table 2. Descriptive statistics of the variables.

Statistic	ln Y	ln FT	ln VAT	ln IS	ln LS	ln RS	ln EE	ln TE	ln L	ln A	ln F
Mean	4.702	2.897	4.411	3.670	3.436	3.106	6.256	5.202	4.202	7.653	3.635
Median	4.928	3.226	4.431	3.880	3.570	3.318	6.354	5.481	4.399	8.036	3.973
Maximum	6.668	5.643	6.735	6.538	5.794	5.279	8.242	7.592	7.370	9.585	5.821
Minimum	1.058	0	0.616	0.103	0.303	0.170	3.514	2.138	0.470	3.840	0.166
Std. Dev.	1.291	1.631	1.133	1.452	1.125	1.179	0.845	0.908	1.273	1.299	1.225
Skewness	-0.824	-1.141	-0.634	-0.862	-0.758	-0.623	-0.570	-0.970	-0.588	-0.896	-0.828
Kurtosis	-0.352	1.722	1.295	0.719	1.536	-0.137	0.278	0.644	0.018	-0.214	0.058

Note: All variables are reported in natural logarithms. Monetary variables are measured in nominal terms. To avoid losing zero observations, variables are transformed as $\ln(x + 1)$.

Before estimating the panel threshold models, we conduct diagnostic tests for multicollinearity and stationarity to ensure the appropriateness of the data and model specification. Specifically, multicollinearity among the explanatory variables is examined using pairwise correlation coefficients and variance inflation factors (VIFs).

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As reported in Table 3, several pairs of explanatory variables exhibit relatively high correlations, with some correlation coefficients exceeding 0.70 in absolute value. In addition, the VIF values for several variables are high, and the maximum VIF reaches 25.795, indicating substantial multicollinearity among the regressors. Therefore, this study employs a panel ridge double-kink model to mitigate multicollinearity and improve the stability of the estimates.

Table 3. Correlation analysis and VIF test.

	VIF	ln FT	ln VAT	ln IS	ln LS	ln RS	ln EE	ln TE	ln L	ln A	ln F
ln FT	3.830	1									
ln VAT	9.269	0.598	1								
ln IS	2.898	0.599	0.504	1							
ln LS	7.682	0.783	0.722	0.679	1						
ln RS	8.420	0.764	0.432	0.682	0.822	1					
ln EE	25.795	0.767	0.860	0.704	0.871	0.758	1				
ln TE	6.445	0.665	0.573	0.699	0.691	0.731	0.833	1			
ln L	4.553	0.689	0.594	0.465	0.804	0.708	0.740	0.530	1		
ln A	7.763	0.604	0.199	0.473	0.605	0.699	0.405	0.269	0.661	1	
ln F	8.878	0.697	0.363	0.505	0.687	0.707	0.530	0.322	0.745	0.908	1

Before conducting the panel unit root tests, it is necessary to determine whether cross-sectional dependence is present in the panel data, because this affects the choice of an appropriate unit root testing procedure. Therefore, this study applies several commonly used tests for cross-sectional dependence, including the LM test proposed by Breusch and Pagan [47], Pesaran's scaled LM test and CD test [48], and the bias-corrected scaled LM test developed by Baltagi, Feng, and Kao [49]. The results are reported in Table 4.

Table 4. Results of cross-section dependence tests.

Variable	CD	LM	Scaled LM	bias-corrected LM
ln FT	0.842	472.315	0.913	0.805
ln VAT	-0.517	458.274	-0.224	-0.336
ln IS	1.104	489.662	1.127	1.015
ln LS	-0.764	451.893	-0.487	-0.603
ln RS	0.693	476.528	0.741	0.629
ln EE	-1.216	444.307	-1.098	-1.205
ln TE	0.558	470.114	0.612	0.503
ln L	-0.935	459.781	-0.731	-0.846
ln A	1.287	495.206	1.342	1.227
ln F	-0.408	462.955	-0.183	-0.294

Note: No test statistic is statistically significant at the 10% level.

As shown in Table 4, the statistics of the CD test, scaled LM test, and bias-corrected LM test are generally insignificant for all variables. Thus, the null hypothesis of no cross-sectional dependence cannot be rejected. This suggests that there is no strong evidence of pervasive cross-sectional dependence in the sample. Under this condition, first-generation panel unit root tests are appropriate for the subsequent analysis. Accordingly, this study employs the Levin–Lin–Chu (LLC) test [50] and the Im–Pesaran–Shin (IPS) test [51].

Table 5 reports the results of the LLC and IPS panel unit root tests, both of which take the presence of a unit root as the null hypothesis. For all variables, the test statistics are negative and statistically significant at the 5% or 1% level, indicating rejection of the unit-root null in the log-level series. The evidence is particularly strong for ln EE, ln IS, ln TE, ln RS, and ln F, while ln L shows relatively weaker but still statistically significant rejection.

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Table 5. Results of panel unit root tests.

Variable	LLC	IPS
ln Y	-4.972***	-1.851**
ln FT	-10.600***	-10.030***
ln VAT	-3.847***	-2.834***
ln IS	-20.585***	-17.040***
ln LS	-9.179***	-1.985**
ln RS	-13.386***	-11.635***
ln EE	-23.886***	-18.686***
ln TE	-15.248***	-13.266***
ln L	-1.852**	-1.925**
ln A	-8.789***	-5.044***
ln F	-13.235***	-8.594***

Note: ***, **, and * represent statistical significance at the levels of 1%, 5%, and 10%, respectively.

Taken at face value, these results suggest that all variables are stationary in levels. Moreover, the cross-sectional dependence tests reported in Table 4 do not provide strong evidence of pervasive cross-sectional dependence across provinces. Therefore, the LLC and IPS results can be interpreted as supporting level stationarity in the sample. Nevertheless, because LLC and IPS are first-generation panel unit root tests that rely on the assumption of cross-sectional independence, the reported results should still be interpreted with some caution, and second-generation panel unit root tests may be used as a robustness check where appropriate.

3.2. Model Specification and Estimation

3.2.1. Model Specification

This study examines the nonlinear relationship between fiscal revenue and expenditure and grain production. To capture possible changes in marginal effects across different policy intensity ranges, both single-kink and double-kink panel models are considered. The corresponding specifications are given in Equations (1) to (3).

Single kink regression:

$$\ln Y_t = \sum_{q=1}^{p} \beta_q^{(1)} X_{qit} I\left(X_{qit} \le \gamma_q^{(1)}\right) + \sum_{q=1}^{p} \beta_q^{(2)} X_{qit} I\left(\gamma_q^{(1)} \le X_{qit}\right) + u_i + \varepsilon_{it} \quad (1)$$

Double kink regression:

$$\ln Y_t = \sum_{j=1}^{k} \beta_j^{(1)} X_{jit} I\left(X_{jit} \le \gamma_j^{(1)}\right) + \sum_{j=1}^{k} \beta_j^{(2)} X_{jit} I\left(\gamma_j^{(1)} < X_{jit} < \gamma_j^{(2)}\right) + \sum_{j=1}^{k} \beta_j^{(3)} X_{jit} I\left(\gamma_j^{(2)} \le X_{jit}\right) + u_i + \varepsilon_{it} \quad (2)$$

Full regression:

$$\begin{aligned}\ln Y_t = {} & \sum_{q=1}^{p} \beta_q^{(1)} X_{qit} I\left(X_{qit} \le \gamma_q^{(1)}\right) + \sum_{q=1}^{p} \beta_q^{(2)} X_{qit} I\left(\gamma_q^{(1)} \le X_{qit}\right) + \sum_{j=1}^{k} \beta_j^{(1)} X_{jit} I\left(X_{jit} \le \gamma_j^{(1)}\right) \\ & + \sum_{j=1}^{k} \beta_j^{(2)} X_{jit} I\left(\gamma_j^{(1)} < X_{jit} < \gamma_j^{(2)}\right) + \sum_{j=1}^{k} \beta_j^{(3)} X_{jit} I\left(\gamma_j^{(2)} \le X_{jit}\right) + \sum_{s=1}^{m} a_s Z_{sit} + u_i + \varepsilon_{it}\end{aligned} \quad (3)$$

where $i = 1, \ldots, N$ indexes provinces and $t = 1, \ldots, T$ indexes years. X_{qit} denotes the set of explanatory variables that are assumed to exhibit a single kink, while X_{jit} denotes the set of explanatory variables that are allowed to exhibit double-kink effects. Z_{sit} represents the set of exogenous variables whose effects are assumed to be regime-invariant. The coefficients $\beta_j^{(1)}$, $\beta_j^{(2)}$ and $\beta_j^{(3)}$ correspond to regimes 1, 2, and 3, respectively. The parameters $\gamma_j^{(1)}$ and $\gamma_j^{(2)}$ denote the kink points associated with variable $I(\cdot)$ is an indicator function determined by the kink value and the kink variable. u_i captures province fixed effects, and ε_{it} is the idiosyncratic error term.

3.2.2. Fixed Effect Estimation with Ridge Regression

Following Hansen [14], the province fixed effects are removed by applying the within transformation, whereby the province-specific mean is subtracted from each observation. Accordingly, Equations (1) and (2) can be expressed in the transformed forms reported in Equations (4) and (5).

$$\ln Y_{it}^{*} = \sum_{q=1}^{p} \beta_q^{\prime\,(1)} X_{qit}^{\prime *} I\left(X_{qit}^{*} \le \gamma_q^{(1)}\right) + \sum_{q=1}^{p} \beta_q^{\prime\,(2)} X_{qit}^{\prime *} I\left(\gamma_q^{(2)} \le X_{qit}^{*}\right) + \varepsilon_{it}^{*} \quad (4)$$

$$\ln Y_{it}^{*} = \sum_{j=1}^{k} \beta_j^{(1)} X_{jit}^{\prime *} I\left(X_{jit}^{*} \le \gamma_j^{(1)}\right) + \sum_{j=1}^{k} \beta_j^{(2)} X_{jit}^{\prime *} I\left(\gamma_j^{(1)} < X_{jit}^{*} < \gamma_j^{(2)}\right) + \sum_{j=1}^{k} \beta_j^{(3)} X_{jit}^{\prime *} I\left(\gamma_j^{(2)} \le X_{jit}^{*}\right) + \varepsilon_{it}^{*} \quad (5)$$

After stacking the transformed observations, the model can be expressed in matrix form as follows:

$$\ln Y^{*} = \beta^{\prime} X^{\prime}(\gamma) + \varepsilon^{*} \quad (6)$$

Because several explanatory variables are highly correlated, standard least-squares estimation may lead to unstable coefficient estimates. To mitigate this problem, this study adopts ridge regression within the fixed-effects kink framework. The ridge estimator is defined as follows:

$$\beta(\gamma) = \left(X^{*}(\gamma)^{\prime} X^{*}(\gamma) + \lambda I\right)^{-1} X^{*}(\gamma)^{\prime} \ln Y^{*} \quad (7)$$

The corresponding residual sum of squares is given by:

$$S_n(\beta,\gamma) = \left(\ln Y^{*} - \beta^{\prime} X^{\prime *}(\gamma)\right)^{2} + \lambda \sum \beta^{2} \quad (8)$$

$\lambda \ge 0$ is a non-negative regularization parameter. The additional term $\lambda \sum \beta^2$ is the ridge penalty function, which shrinks kink regression coefficients towards zero as λ increases. We already penalize only β and opt not to penalize γ since it is worth noting when γ is shrunk to zero. It is important to note that $\lambda \to \infty$ ridge penalty function approaches infinity and yields an empty ridge estimation whereas when $\lambda \to 0$, the ridge penalty function approaches zero, and the ridge kink regression estimator is reduced to least-square regression estimator. The role of the regularization parameter λ is essential in the estimation process. However, the loss function is quadratic in β but nonconvex in λ. Thus, it is computationally convenient to view this parameter as the tuning parameter for controlling the overall penalty level. Similar to the least-square estimator, given the candidate $\tilde{\lambda}$, ridge kink regression estimator $(\hat{\beta},\hat{\gamma},\lambda)$ is the joint minimizer of $S_n(\beta,\gamma,\tilde{\lambda})$:

$$\left(\hat{\beta},\hat{\gamma},\tilde{\lambda}\right) = \underset{\beta \in R^{p}, \gamma \in \Gamma}{\arg\min}\, S_n\left(\beta,\gamma,\tilde{\lambda}\right) \quad (9)$$

Then, the optimal $\tilde{\lambda}$ can be easily determined by minimization of the concentrated sum of the squared errors Equation (7). Once the optimal $\tilde{\lambda}$ is defined, we can estimate $\hat{\beta}$ and $\hat{\gamma}$ using standard penalized least-squares of Y^{*} on $X^{*}(\gamma)$ with the additional penalty function $\hat{\lambda} \sum \beta^2$.

3.2.3. Kink Effects Testing

To determine whether a statistically significant kink effect is present, this study tests the null hypothesis of no kink effect against the alternative hypothesis of a regime-dependent slope change. Following Hansen [14], the test is based on an F-statistic, and bootstrap p-values are computed to account for the nonstandard distribution of the test statistic. If the null hypothesis cannot be rejected, the linear panel specification is retained. Otherwise, the panel kink model is adopted to capture changes in the marginal effects of the explanatory variables across regimes. The test statistic is defined in Equation (10).

$$F(\gamma) = \frac{S_0(\beta) - S_1(\beta,\gamma)}{\hat{\sigma}^{2}} \quad (10)$$

where $S_0(\beta)$ is the sum squared residuals of the linear model, while $S_1(\beta,\gamma)$ is the sum squared residuals of the kink model. $\hat{\sigma}^2$ is the residual variance of the kink model.

Table 6 reports the results of the panel kink effect tests for the fiscal policy variables. The null hypothesis of no kink effect is rejected for $\ln$ FT, $\ln$ VAT, $\ln$ IS, $\ln$ LS, $\ln$ RS, and $\ln$ EE, indicating that these variables exhibit significant nonlinear effects on grain production. For $\ln$ TE, the p-value is 0.10,

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which provides only weak evidence of a kink effect at the 10% significance level. Therefore, the results suggest that most fiscal variables are associated with regime-dependent marginal effects on grain production. In the next step, fixed-effects specifications are compared using the Akaike information criterion (AIC) to assess whether a single-kink or double-kink model provides a better fit for each variable.

Table 6. Panel kink effect test of fiscal policy variable.

	ln FT	**ln VAT**	**ln IS**	**ln LS**	**ln RS**	**ln EE**	**ln TE**
F_k^*	14.79	49.41	12.62	19.54	17.32	7.93	5.20
p- value	0.00	0.00	0.00	0.00	0.00	0.00	0.1

Table 7 compares the goodness of fit of the single-kink and double-kink specifications using the Akaike information criterion (AIC). Since a lower AIC indicates a better model fit, the results suggest that the double-kink specification is preferred for agricultural insurance subsidies and education expenditure, whose AIC values are lower under the double-kink model than under the single-kink model. By contrast, for ln FT, ln VAT, ln LS, and ln RS, the single-kink specification provides a better fit. Therefore, only agricultural insurance subsidies and education expenditure are identified as exhibiting double-kink relationships with grain production. The estimated kink points are reported in Table 8

Table 7. Comparison of single-kink and double-kink models.

AIC	**ln FT**	**ln VAT**	**ln IS**	**ln LS**	**ln RS**	**ln EE**
Single kink	**-103.94**	**27.85**	-134.45	**-147.76**	**-377.07**	77.98
Double kink	-35.73	36.83	**-169.17**	-73.73	-200.20	**53.51**

Table 8. Kink point estimates.

	Ridge-Fixed Effect	Fixed Effect	Random Effect
lnFT kink point	0.874	0.855	0.854
lnVAT kink point	4.912	4.983	4.844
lnIS first kink point	3.496	3.461	3.485
lnIS second kink point	4.199	4.112	4.168
lnLS kink point	2.229	2.229	2.229
lnRS kink point	3.007	3.459	3.319
lnEE first kink point	5.965	5.400	5.432
lnEE second kink point	6.534	6.651	6.613

Note: Kink point estimates are reported under ridge fixed-effects, fixed-effects, and random-effects specifications.

4. Results

4.1. Panel Kink Regression Result

Table 9 reports the panel kink and double-kink regression estimates for 31 provinces in China under the ridge fixed-effects, fixed-effects, and random-effects specifications. According to the Akaike information criterion (AIC), the ridge fixed-effects model provides the best fit among the three specifications, with the lowest AIC value (-721.43). Therefore, the discussion below focuses primarily on the ridge fixed-effects results, while the fixed-effects and random-effects estimates are used as supplementary evidence on the robustness of the estimated patterns.

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Table 9. Panel kink and double-kink regression results for 31 provinces in China.

	Panel Kink Regression		
Variable	**Ridge-Fixed Effect**	**Fixed Effect**	**Random Effect**
Intercept	-	-	-0.066*** (0.012)
$\ln FT \le \gamma_1^{(1)}$	-0.021* (0.011)	-0.005 (0.005)	-0.002 (0.005)
$\ln FT > \gamma_1^{(1)}$	0.014 (0.445)	0.012*** (0.003)	0.011** (0.005)
$\ln VAT \le \gamma_2^{(1)}$	-0.019** (0.008)	-0.046 (0.029)	-0.054 (0.034)
$\ln VAT > \gamma_2^{(1)}$	-0.017 (0.063)	-0.006*** (0.001)	-0.007** (0.003)
$\ln IS \le \gamma_3^{(1)}$	0.067*** (0.026)	0.048*** (0.014)	0.046*** (0.014)
$\gamma_3^{(1)} < \ln IS < \gamma_3^{(2)}$	0.247*** (0.007)	0.142*** (0.030)	0.128** (0.059)
$\ln IS \ge \gamma_3^{(2)}$	0.191*** (0.017)	0.134*** (0.016)	0.089*** (0.016)
$\ln LS \le \gamma_4^{(1)}$	0.007 (0.141)	0.017 (0.016)	0.022* (0.013)
$\ln LS > \gamma_4^{(1)}$	0.050*** (0.019)	0.045*** (0.009)	0.050*** (0.009)
$\ln RS \le \gamma_5^{(1)}$	0.066*** (0.025)	0.003 (0.008)	0.001 (0.009)
$\ln RS > \gamma_5^{(1)}$	-0.035** (0.014)	-0.056** (0.026)	-0.044** (0.020)
$\ln EE \le \gamma_6^{(1)}$	0.118*** (0.015)	0.076* (0.045)	0.089* (0.051)
$\gamma_6^{(1)} < \ln EE < \gamma_6^{(2)}$	0.013 (0.008)	0.035 (0.049)	0.005 (0.037)
$\ln EE \ge \gamma_6^{(2)}$	-0.228*** (0.008)	-0.167*** (0.041)	-0.168*** (0.039)
$\ln TE$	0.020 (0.021)	0.031 (0.022)	0.033 (0.022)
$\ln L$	0.077*** (0.021)	0.043*** (0.012)	0.042*** (0.011)
$\ln A$	0.511*** (0.098)	0.762*** (0.013)	0.824*** (0.017)
$\ln F$	0.269*** (0.087)	0.102*** (0.017)	0.121*** (0.017)
AIC	**-721.43**	-575.26	-564.68

Note: () is the standard error, "*," "**," and "***" denote rejection of the null hypothesis at the 10%, 5%, and 1% significance level, respectively.

The results indicate that the relationship between fiscal policy and grain production is nonlinear and regime dependent. Agricultural insurance subsidies (IS) are positive in all three regimes, and the coefficient is largest in the middle regime, suggesting that their strongest marginal contribution to grain production occurs at an intermediate subsidy intensity. Education expenditure (EE) is significantly positive in the low-intensity regime, insignificant in the middle regime, and significantly negative in the high-intensity regime, implying that its effect on grain production may weaken and eventually reverse once expenditure exceeds a certain level. Agricultural loan interest subsidies (LS) are insignificant below the kink point but become significantly positive above it, indicating that stronger credit support may help ease production constraints. Rural minimum living security subsidies (RS) are significantly positive in the low regime but significantly negative in the high regime, suggesting that their effect may shift from support to disincentive as subsidy intensity increases.

For the revenue-side instruments, VAT has a negative coefficient in both regimes in the benchmark ridge fixed-effects model, although only the lower-regime coefficient is statistically significant. In the fixed-effects and random-effects models, the upper-regime coefficient is also significantly negative, which provides additional support for a negative association between VAT and grain production. The cultivated land occupation tax (FT) is weakly negative below the kink point in the benchmark model and turns positive above it. This pattern is consistent with the fixed-effects and random-effects estimates, in which the upper-regime coefficient is also positive and statistically significant, suggesting that the land-protection effect of FT becomes more evident once policy intensity exceeds the estimated kink point.

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Among the control variables, labor (L), land area (A), and fertilizer use (F) are consistently positive and statistically significant across specifications, whereas transportation infrastructure expenditure (TE) is not statistically significant. Overall, the results suggest that fiscal policy affects grain production through intensity-dependent nonlinear mechanisms, although the strength and statistical significance of these effects vary across fiscal instruments and model specifications.

4.2. Robustness Tests

To further assess the reliability of the baseline results, this study conducts a series of robustness checks. These checks are designed to examine whether the estimated kink points and regime-specific coefficients remain stable under alternative model specifications. Specifically, the robustness analysis includes two parts. The first part examines sensitivity to omitted controls, and the second part evaluates whether the main conclusions remain broadly unchanged under alternative estimators that account for potential endogeneity and dynamic adjustment.

4.2.1. Sensitivity to Omitted Controls

To assess whether the baseline results are sensitive to omitted-variable bias, the models are re-estimated without including the control variables. The purpose is to examine whether the estimated kink points and the nonlinear effects of fiscal revenue and fiscal expenditure on grain production in China's 31 provinces remain broadly stable under a more parsimonious specification.

Table 10 reports the robustness checks for the fiscal revenue variables under alternative specifications, including models with and without control variables as well as different estimators. Overall, the estimated kink points remain relatively stable across specifications, suggesting that the identified nonlinear patterns are robust. For $\ln \mathrm{FT}$, the coefficient below the kink point is negative or insignificant, whereas the coefficient above the kink point is generally positive and statistically significant, indicating that the positive effect mainly emerges after the kink point is crossed. For $\ln \mathrm{VAT}$, the estimated coefficients remain negative in both regimes, and the upper-regime effect is generally more adverse. These results are consistent with the baseline findings and suggest that the main conclusions for the fiscal revenue variables are broadly robust to alternative control specifications.

Table 10. Robustness test for fiscal revenue variables.

Variable	Ridge-fixed Effect	Fixed Effect	Random Effect	Difference GMM with IV	System GMM with IV
			$\ln FT$ With control variables		
Kink point γ	0.893	0.868	0.867	0.858	0.862
Intercept	-	-	-0.019* (0.011)	-	-
$\ln \mathrm{FT} \leq \gamma$	-0.004* (0.002)	-0.003 (0.003)	-0.005 (0.005)	-0.001 (0.004)	-0.002(0.002)
$\ln \mathrm{FT} > \gamma$	0.017** (0.008)	0.011** (0.005)	0.011** (0.004)	0.007*** (0.002)	0.003* (0.002)
$\ln L$	0.068** (0.030)	0.029** (0.013)	0.042*** (0.010)	0.007* (0.004)	0.009** (0.004)
$\ln A$	0.625*** (0.061)	0.863*** (0.014)	0.852*** (0.015)	0.726*** (0.091)	0.086*** (0.025)
$\ln F$	0.277*** (0.058)	0.093*** (0.018)	0.112*** (0.018)	0.025 (0.023)	0.001 (0.007)
			$\ln FT$ without control variables		
Kink point γ	0.891	0.868	0.896	0.860	0.861
Intercept	-	-	-0.327*** (0.059)	-	-
$\ln \mathrm{FT} \leq \gamma$	-0.117* (0.065)	-0.131*** (0.020)	-0.122*** (0.019)	-0.006 (0.007)	-0.002* (0.001)

$\ln \mathrm{FT} > \gamma$	0.243*** (0.019)	0.238*** (0.020)	0.255*** (0.019)	0.008*** (0.003)	0.004** (0.002)
$\ln \mathrm{VAT}$ with control variables					
Kink point γ	4.912	4.121	4.510	4.454	4.608
Intercept	-	-	-0.030*** (0.009)	-	-
$\ln \mathrm{VAT} \le \gamma$	-0.032* (0.018)	-0.039*** (0.005)	-0.051*** (0.005)	-0.022** (0.010)	-0.006*** (0.002)
$\ln \mathrm{VAT} > \gamma$	-0.003** (0.001)	-0.006*** (0.002)	-0.008*** (0.003)	-0.008 (0.007)	-0.001 (0.002)
$\ln L$	0.078*** (0.025)	0.022** (0.010)	0.027** (0.010)	0.006* (0.003)	0.011** (0.005)
$\ln A$	0.627*** (0.059)	0.882*** (0.015)	0.878*** (0.015)	0.711*** (0.007)	0.095*** (0.026)
$\ln F$	0.285*** (0.055)	0.101*** (0.017)	0.118*** (0.017)	0.015 (0.014)	0.001 (0.007)
$\ln \mathrm{VAT}$ without control variables					
Kink point γ	4.240	4.290	4.252	4.454	4.736
Intercept	-	-	0.322*** (0.065)	-	-
$\ln \mathrm{VAT} \le \gamma$	-0.314*** (0.101)	-0.332*** (0.028)	-0.329*** (0.033)	-0.090*** (0.019)	-0.006** (0.003)
$\ln \mathrm{VAT} > \gamma$	-0.244*** (0.074)	-0.322*** (0.050)	-0.271*** (0.050)	-0.019** (0.008)	-0.002 (0.004)

Note: () is the standard error, "*," "**," and "***" denote rejection of the null hypothesis at the 10%, 5%, and 1% significance level, respectively.

Table 11 reports the robustness checks for agricultural insurance subsidies and agricultural loan interest subsidies. For $\ln \mathrm{IS}$, the coefficients remain positive across regimes, and the middle regime generally shows the strongest effect, supporting the baseline finding of an inverted-U type relationship. For $\ln \mathrm{LS}$, the coefficients are positive in both regimes, and the upper-regime coefficient is generally larger, indicating that stronger interest-subsidy support is associated with greater grain production. Overall, the estimated kink points remain reasonably stable across specifications, which further supports the robustness of the identified nonlinear patterns.

Table 11. Robustness tests for fiscal expenditure variables.

Variable	**Ridge-fixed Effect**	**Fixed Effect**	**Random Effect**	**Difference GMM with IV**	**System GMM with IV**
$\ln \mathrm{IS}$ With control variables					
First kink point $\boldsymbol{\gamma}^{(1)}$	2.979	3.628	2.979	3.500	3.219
Second kink point $\boldsymbol{\gamma}^{(2)}$	4.242	4.316	3.829	4.237	4.373
Intercept	-	-	-0.020* (0.011)	-	-
$\ln \mathrm{IS} \le \gamma^{(1)}$	0.069*** (0.023)	0.056*** (0.014)	0.048*** (0.013)	0.001 (0.012)	0.017* (0.010)
$\gamma^{(1)} < \ln \mathrm{IS} < \boldsymbol{\gamma}^{(2)}$	0.243*** (0.010)	0.142*** (0.033)	0.136*** (0.032)	0.021** (0.010)	0.019 (0.013)
$\ln \mathrm{IS} \ge \boldsymbol{\gamma}^{(2)}$	0.176*** (0.024)	0.141*** (0.016)	0.107*** (0.016)	0.013** (0.006)	0.011** (0.005)
$\ln L$	0.098*** (0.033)	0.051*** (0.009)	0.063*** (0.009)	0.005* (0.003)	0.004 (0.005)
$\ln A$	0.565*** (0.053)	0.814*** (0.014)	0.804*** (0.015)	0.703*** (0.095)	0.138** (0.063)
$\ln F$	0.262*** (0.050)	0.089*** (0.016)	0.111*** (0.016)	0.025* (0.014)	0.009 (0.012)
$\ln \mathrm{IS}$ without control variables					
First kink point $\gamma^{(1)}$	2.973	3.634	3.481	3.499	3.680

Second kink point $\boldsymbol{\gamma}^{(2)}$	4.189	4.278	3.891	4.027	4.145
Intercept	-	-	-0.071* (0.041)	-	-
$\ln \mathrm{IS} \leq \gamma^{(1)}$	0.741*** (0.040)	0.744*** (0.071)	0.757*** (0.069)	0.004 (0.022)	0.017* (0.010)
$\gamma^{(1)} < \ln \mathrm{IS} < \boldsymbol{\gamma}^{(2)}$	1.219*** (0.132)	1.390*** (0.192)	1.341*** (0.185)	0.080** (0.036)	0.031** (0.014)
$\ln \mathrm{IS} \geq \boldsymbol{\gamma}^{(2)}$	1.007*** (0.035)	1.057*** (0.081)	1.054*** (0.079)	0.046** (0.022)	0.012 (0.015)
			$\ln \mathrm{LS}$ with control variables		
Kink point γ	2.229	2.900	2.426	2.738	2.370
Intercept	-	-	-0.033*** (0.009)	-	-
$\ln \mathrm{LS} \leq \gamma$	0.008* (0.005)	0.010* (0.006)	0.013** (0.006)	0.005* (0.003)	0.002 (0.002)
$\ln \mathrm{LS} > \gamma$	0.033** (0.015)	0.039*** (0.008)	0.048*** (0.008)	0.018* (0.010)	0.009** (0.004)
$\ln L$	0.062** (0.028)	0.018* (0.010)	0.023** (0.010)	0.005 (0.007)	0.007 (0.006)
$\ln A$	0.632*** (0.059)	0.869*** (0.014)	0.860*** (0.014)	0.718*** (0.095)	0.174** (0.068)
$\ln F$	0.285*** (0.055)	0.104*** (0.017)	0.120*** (0.017)	0.019 (0.027)	0.015 (0.016)
			$\ln \mathrm{LS}$ without control variables		
Kink point γ	2.724	2.778	2.712	2.738	2.515
Intercept	-	-	-0.091* (0.052)	-	-
$\ln \mathrm{LS} \leq \gamma$	0.253*** (0.084)	0.243*** (0.081)	0.271*** (0.076)	0.003* (0.002)	0.005** (0.002)
$\ln \mathrm{LS} > \gamma$	0.288*** (0.081)	0.296*** (0.050)	0.307*** (0.047)	0.023* (0.013)	0.010** (0.005)

Note: () is the standard error, "*," "**," and "***" denote rejection of the null hypothesis at the 10%, 5%, and 1% significance level, respectively.

Table 12 reports the robustness checks for the remaining fiscal expenditure variables. For $\ln \mathrm{RS}$, the coefficient is positive or weak in the lower regime but turns negative in the upper regime, suggesting that excessive transfer intensity may weaken grain production incentives. For $\ln \mathrm{EE}$, the estimates indicate a double-kink structure, with positive effects in the lower and middle regimes but negative effects once expenditure exceeds the second kink point. For $\ln \mathrm{TE}$, the coefficient is generally positive, although its statistical significance is less stable across specifications. Taken together, the main conclusions for the fiscal expenditure variables remain broadly robust under alternative control specifications.

Table 12. Robustness tests of fiscal expenditure variables (continued).

Variable	**Ridge-fixed Effect**	**Fixed Effect**	**Random Effect**	**Difference GMM with IV**	**System GMM with IV**
			$\ln \mathrm{RS}$ With control variables		
Kink point γ	3.319	2.939	3.098	3.518	3.582
Intercept	-	-	0.004 (0.010)	-	-
$\ln \mathrm{RS} \leq \gamma$	0.035** (0.015)	0.021** (0.009)	0.026*** (0.007)	0.011 (0.008)	0.007 (0.006)
$\ln \mathrm{RS} > \gamma$	-0.009* (0.005)	-0.072*** (0.019)	-0.068*** (0.017)	-0.024** (0.010)	-0.025* (0.014)
ln L	0.069*** (0.021)	0.049*** (0.010)	0.058*** (0.010)	0.007 (0.008)	0.003 (0.008)
ln A	0.610*** (0.059)	0.892*** (0.016)	0.884*** (0.016)	0.719*** (0.096)	0.208*** (0.073)
ln F	0.275*** (0.055)	0.092*** (0.017)	0.110*** (0.017)	0.016 (0.034)	0.015 (0.017)
			$\ln \mathrm{RS}$ without control variables		
Kink point γ	3.374	3.569	3.300	3.459	3.594

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Intercept	-	-	0.175** (0.076)	-	-
$\ln RS \leq \gamma$	0.510*** (0.113)	0.534*** (0.024)	0.539*** (0.025)	0.039** (0.017)	0.002 (0.004)
$\ln RS > \gamma$	-0.465*** (0.018)	-0.693*** (0.084)	-0.586*** (0.082)	-0.089 (0.060)	-0.012*** (0.003)
$\ln EE$ with control variables					
First kink point $\boldsymbol{\gamma}^{(1)}$	5.970	5.381	5.377	5.438	5.276
Second kink point $\boldsymbol{\gamma}^{(2)}$	6.597	6.195	6.358	6.266	6.754
Intercept	-	-	0.001 (0.009)	-	-
$\ln EE \leq \gamma^{(1)}$	0.115*** (0.014)	0.086** (0.037)	0.084** (0.036)	0.043* (0.024)	0.009* (0.005)
$\gamma^{(1)} < \ln EE < \boldsymbol{\gamma}^{(2)}$	0.146*** (0.007)	0.108** (0.053)	0.134** (0.058)	0.016 (0.047)	0.008 (0.021)
$\ln EE \geq \boldsymbol{\gamma}^{(2)}$	-0.025* (0.014)	-0.049* (0.028)	-0.066** (0.028)	-0.160** (0.070)	-0.003* (0.002)
lnL	0.063*** (0.019)	0.014 (0.012)	0.022* (0.012)	0.003 (0.009)	0.010** (0.004)
lnA	0.622*** (0.053)	0.866*** (0.015)	0.858*** (0.015)	0.840*** (0.130)	0.085*** (0.026)
lnF	0.283*** (0.050)	0.094*** (0.017)	0.110*** (0.017)	0.055* (0.031)	0.002 (0.008)
$\ln EE$ without control variables					
First kink point $\boldsymbol{\gamma}^{(1)}$	5.896	5.381	5.381	5.239	5.299
Second kink point $\boldsymbol{\gamma}^{(2)}$	6.609	6.657	6.010	6.266	6.770
Intercept	-	-	0.099* (0.055)	-	-
$\ln EE \leq \gamma^{(1)}$	1.681*** (0.202)	1.739*** (0.158)	1.757*** (0.151)	0.114* (0.063)	0.031** (0.013)
$\gamma^{(1)} < \ln EE < \boldsymbol{\gamma}^{(2)}$	1.616*** (0.100)	1.509*** (0.335)	1.689*** (0.290)	0.082* (0.045)	0.040* (0.021)
$\ln EE \geq \boldsymbol{\gamma}^{(2)}$	-0.656*** (0.203)	-0.529*** (0.155)	-0.682*** (0.148)	-0.014 (0.044)	-0.035*** (0.011)
$\ln TE$ with control variables					
Intercept	-	-	0.114*(0.063)	-	-
$\ln TE$	0.096*** (0.027)	0.055** (0.019)	0.048** (0.020)	0.009 (0.014)	0.002 (0.010)
lnL	0.051** (0.022)	0.017 (0.011)	0.031*** (0.009)	0.007* (0.004)	0.004* (0.002)
lnA	0.621*** (0.065)	0.855*** (0.015)	0.846*** (0.015)	0.721*** (0.096)	0.165*** (0.050)
lnF	0.291*** (0.061)	0.111*** (0.018)	0.126*** (0.018)	0.019 (0.031)	0.014* (0.008)
$\ln TE$ without control variables					
Intercept	-	-	0.139*** (0.021)	-	-
$\ln TE$	0.837*** (0.054)	1.053*** (0.100)	1.135*** (0.097)	0.015* (0.008)	0.028*** (0.005)

Note: () is the standard error, "*," "**," and "***" denote rejection of the null hypothesis at the 10%, 5%, and 1% significance level, respectively.

Overall, the results reported in Tables 10 to 12 show that removing the control variables does not materially alter the estimated kink points, coefficient signs, or general nonlinear patterns. This suggests that the baseline conclusions are not driven by a specific set of control variables and are reasonably robust to omitted-variable concerns.

4.2.2. Alternative Estimators and Endogeneity Checks

To address potential endogeneity in the panel kink regressions, this study employs two alternative estimation strategies. The first is an instrumental-variable-based panel kink framework, which follows the logic of threshold models with endogenous regressors developed by Caner and Hansen [52] and Fouquau et al. [53]. The second is a dynamic panel specification that incorporates lagged dependent variables and is estimated using GMM-based methods, following Kremer et al. [54], Seo and Shin [15], Arellano and Bover [17], and Blundell and Bond [18].

Lagged values of the explanatory variables, up to five periods, are used as instrumental variables. These instruments are demeaned and allowed to vary across regimes through the indicator function $I(\cdot)$. Following Fouquau et al. [53], the regime-dependent coefficients are estimated as:

$$\begin{aligned}\hat{\beta}_{IV}(\hat{\gamma}) = \left(X^*(\hat{\gamma})'\eta^*_{it}(\hat{\gamma}) \cdot \left(\eta^*_{it}(\hat{\gamma})'\eta^*_{it}(\hat{\gamma})\right)^{-1} \cdot \eta^*_{it}(\hat{\gamma})'X^*(\hat{\gamma})\right)^{-1} X^*(\hat{\gamma})'\eta^*_{it}(\hat{\gamma}) \\ \cdot \left(\eta^*_{it}(\hat{\gamma})'\eta^*_{it}(\hat{\gamma})\right)^{-1} \cdot \eta^*_{it}(\hat{\gamma})' \ln Y^*\end{aligned} \tag{11}$$

where $\eta^*_{it}(\hat{\gamma}) = Z_{it}(\hat{\gamma}) - \bar{Z}_{it}(\hat{\gamma})$ is regime dependent instrumental variables. To assess the validity of these instruments, we employ the overidentifying restrictions test, as proposed by Hansen [55].

As a further check, this study also estimates a dynamic panel kink specification. Compared with the static panel model, the dynamic specification allows the inclusion of lagged dependent variables and helps account for persistence in grain production. In addition, it provides a way to mitigate potential endogeneity and reverse causality involving the fiscal variables. Under this framework, Equations (1) and (2) can be extended to the dynamic single-kink and dynamic double-kink models, as shown in Equations (12) and (13).

Single-dynamic regression:

$$\begin{aligned}\ln Y^*_{it} = a^{(1)} \ln Y_{it-1} I\left(X_{qit} \le \gamma_q^{(1)}\right) + \sum_{q=1}^{p} \beta_q'^{(1)} X'_{qit} I\left(X_{qit} \le \gamma_q^{(1)}\right) + a^{(2)} \ln Y_{it-1} I\left(\gamma_q^{(2)} < X_{qit}\right) \\ + \sum_{q=1}^{p} \beta_q'^{(2)} X'_{qit} I\left(\gamma_q^{(2)} < X_{qit}\right) + u_i + \varepsilon^*_{it}\end{aligned} \tag{12}$$

Double-dynamic regression:

$$\begin{aligned}\ln Y^*_{it} = a^{(1)} \ln Y_{it-1} I\left(X_{jit} \le \gamma_j^{(1)}\right) + \sum_{j=1}^{k} \beta_j'^{(1)} X'_{jit} I\left(X_{jit} \le \gamma_j^{(1)}\right) \\ + a^{(2)} \ln Y_{it-1} I\left(\gamma_j^{(1)} < X_{jit} \le \gamma_j^{(2)}\right) + \sum_{j=1}^{k} \beta_j'^{(2)} X'_{jit} I\left(\gamma_j^{(1)} < X_{jit} \le \gamma_j^{(2)}\right) \\ + a^{(3)} \ln Y_{it-1} I\left(\gamma_j^{(2)} < X_{jit}\right) + \sum_{j=1}^{k} \beta_j'^{(3)} X'_{jit} I\left(\gamma_j^{(2)} < X_{jit}\right) + u_i + \varepsilon^*_{it}\end{aligned} \tag{13}$$

5. Policy Recommendation

Based on the empirical results, fiscal policy affects grain production in a nonlinear and regime-dependent manner. Therefore, policy design should focus not only on the direction of fiscal intervention, but also on its intensity. A key implication of this study is that both insufficient and excessive intervention may weaken policy effectiveness. The following recommendations are proposed.

1. Improve the coordination between land protection and tax policy, and reduce the indirect tax burden on grain production

The results show that the cultivated land occupation tax only begins to exert a more favorable effect on grain production after the kink point is exceeded, while value-added tax is negatively associated with grain output in both regimes. This suggests that, in China, tax policy should simultaneously serve two goals: protecting cultivated land and avoiding excessive production burdens on farmers. On the one hand, in provinces facing strong pressure from urban expansion, industrial land demand, and local fiscal dependence on land conversion, the cultivated land occupation tax should be strengthened so that it can more effectively restrain the non-agricultural conversion of cultivated land. On the other hand, VAT-related burdens that are indirectly passed on through agricultural inputs, processing, transport, and circulation should be reduced as much as possible, especially for small-scale producers who are less able to absorb rising costs. In practice, this means that tax policy should not simply focus on revenue collection but should be better aligned with the national objective of stabilizing grain production capacity.

2. Keep production-oriented support within an effective policy range, rather than simply increasing subsidies across the board

The empirical results indicate that agricultural insurance subsidies and agricultural loan interest subsidies are both helpful for grain production, but their effects are clearly regime dependent. Agricultural insurance subsidies have the strongest effect in the middle regime, while agricultural loan interest subsidies become much more effective only after the kink point is crossed. This implies that China's production-oriented support policies should focus on "appropriate intensity" rather than "maximum intensity." For agricultural insurance, the priority should be to keep subsidies at a level that is high enough to meaningfully reduce risk, but not so high that marginal gains become weaker. For agricultural credit support, more attention should be given to farmers and regions that remain credit constrained, especially in major grain-producing areas where production depends heavily on seasonal financing for seeds, fertilizer, machinery, and other inputs. In this sense, the policy goal should be to improve the precision and efficiency of subsidies, rather than merely expanding fiscal expenditure.

3. Coordinate social support and human-capital expenditure with grain-production goals to avoid unintended disincentive effects

The results show that rural minimum living security subsidies turn negative in the upper regime, and education expenditure also becomes negative after the second kink point. This suggests that, although these expenditures may be beneficial at lower levels, excessively high intensity may weaken agricultural labor supply or reduce production incentives. In the Chinese context, this is especially important because many grain-producing regions still face labor outflow, aging in agriculture, and rising opportunity costs of farming. Therefore, social assistance and education expenditure should not be designed in isolation from agricultural policy. For rural minimum living security subsidies, the policy focus should remain on guaranteeing basic livelihoods, but excessive dependence on transfers should be avoided. For education expenditure, greater emphasis should be placed on rural vocational education, agricultural technology training, and the cultivation of skills directly related to modern farming, so that education investment strengthens agricultural productivity rather than only accelerating movement away from grain production. In short, these expenditures should be better coordinated with the broader objective of rural productive development.

4. Adopt differentiated, region-specific, and intensity-sensitive fiscal policy design

A general conclusion of this study is that different fiscal instruments follow different nonlinear patterns, and even the same instrument may produce opposite effects at different intensity levels. This means that China's grain-supporting fiscal policy should move away from a uniform expansion approach and instead adopt differentiated strategies across instruments and regions. For example, in provinces where land conversion pressure is high, the cultivated land occupation tax should play a stronger protective role; in major grain-producing provinces with high exposure to weather and market risks, agricultural insurance and credit support should be kept within their most effective ranges; and in regions experiencing rapid labor outflow, social and education expenditure should be more carefully coordinated with agricultural development goals. More broadly, policymakers should pay closer attention to whether a fiscal instrument is currently below, near, or above its effective kink point. Only by matching policy intensity to local conditions can fiscal resources be used more efficiently and grain security be supported in a more sustainable way.

5. Recommendations for future research

Although this study provides provincial-level evidence on the nonlinear effects of fiscal policy on grain production, future research should further examine the underlying transmission mechanisms using more disaggregated data, such as county-level or household-level datasets. Greater attention could be given to how fiscal policy affects grain production through land allocation, labor supply, credit access, risk management, and technology adoption, as well as to whether these effects vary across regions with different resource endowments, market conditions, and stages of rural transformation. Further work could also incorporate spatial spillovers and dynamic adjustment processes to provide a more complete understanding of how fiscal policy supports grain security in the long run.

6. Conclusions

The overall results suggest that the effects of fiscal revenue and expenditure on grain production in China are nonlinear rather than linear and are better characterized by single-kink or double-kink relationships. The results from the fixed-effects model and the panel ridge kink model are broadly consistent, confirming that several fiscal variables exert significant nonlinear effects on grain production.

Three main findings emerge from the analysis. First, the cultivated land occupation tax appears to play a role in protecting cultivated land and supporting grain production, but this effect is regime dependent. In the lower regime, its effect is negative or weak, suggesting that a relatively low tax intensity may be insufficient to effectively restrain the conversion of cultivated land. In the higher regime, however, the estimated effect becomes positive, indicating that the tax begins to exert a stronger land-protection function.

Second, agricultural insurance subsidies are positively associated with grain production in all regimes, but their effect is strongest in the middle regime. This implies that agricultural insurance subsidies are most effective when they remain within an intermediate intensity range, where they can better reduce farmers' risk exposure and ease production constraints. Once the subsidy level moves beyond this range, the positive marginal effect begins to weaken.

Third, rural minimum living security subsidies also exhibit a nonlinear effect on grain production. At lower levels, these subsidies may help relieve basic consumption pressure and indirectly support agricultural production. However, in the upper regime, the estimated effect turns negative, suggesting that excessively high transfer intensity may weaken production incentives and reduce grain output.

Overall, these findings indicate that the effectiveness of fiscal policy depends not only on the direction of intervention, but also on its intensity. Therefore, fiscal policy aimed at supporting grain production should pay greater attention to regime-specific effects and avoid a uniform expansion approach

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References

1. United Nations, Department of Economic and Social Affairs. 2024. Available online: https://www.un.org/development/desa/pd/sites/www.un.org.development.desa.pd/files/undesa_pd_2024_e_cn.9_2024_2_advance-unedited.pdf
2. IPCC. Urgent climate action can secure a liveable future for all. 2023. Available online: https://www.ipcc.ch/2023/03/20/press-release-ar6-synthesis-report/
3. FAO; IFAD; UNICEF; WFP; WHO. The State of Food Security and Nutrition in the World 2024—Financing to End Hunger, Food Insecurity and Malnutrition in All Its Forms; Rome, 2024. https://doi.org/10.4060/cd1254en
4. National Bureau of Statistics of China. 2025. Available online: https://www.stats.gov.cn/english/PressRelease/202502/t20250228_1958822.html
5. Wang, S.; Wu, H.; Li, J.; Xiao, Q.; Li, J. Assessment of the effect of the main grain-producing areas policy on China's food security. Foods 2024, 13, 654. https://doi.org/10.3390/foods13050654
6. National Bureau of Statistics of China. 2023. Available online: https://www.stats.gov.cn/WZWSREL2VuZ2xpc2gvUHJlc3NSZWxlYXNlLy8yMDIzMTIvdDIwMjMxMjIxXzE5NDU3MDguaHRtbA==
7. OECD. Agricultural Policy Monitoring and Evaluation 2024: Innovation for Sustainable Productivity Growth; OECD Publishing: Paris, 2024. https://doi.org/10.1787/74da57ed-en
8. Wang, X.; Shen, Y. The effect of China's agricultural tax abolition on rural families' incomes and production. China Economic Review 2014, 29, 185–199. https://doi.org/10.1016/j.chieco.2014.04.010
9. Fan, S.; Zhang, L.; Zhang, X. Growth and Poverty in Rural China: The Role of Public Investments. 2000. https://doi.org/10.22004/ag.econ.16115
10. Deng, H.; Jin, Y.; Pray, C.; Hu, R.; Xia, E.; Meng, H. Impact of public research and development and extension on agricultural productivity in China from 1990 to 2013. China Economic Review 2021, 70, 101699. https://doi.org/10.1016/j.chieco.2021.101699
11. Donaldson, D. Railroads of the Raj: Estimating the impact of transportation infrastructure. American Economic Review 2018, 108, 899–934. DOI: 10.1257/aer.20101199
12. Hansen, B.E. Regression kink with an unknown threshold. Journal of Business & Economic Statistics 2017, 35, 228–240. https://doi.org/10.1080/07350015.2015.1073595
13. Zhang, Y.; Zhou, Q.; Jiang, L. Panel kink regression with an unknown threshold. Economics Letters 2017, 157, 116–121. https://doi.org/10.1016/j.econlet.2017.05.033
14. Hansen, B.E. Threshold effects in non-dynamic panels: Estimation, testing, and inference. Journal of Econometrics 1999, 93, 345–368. https://doi.org/10.1016/S0304-4076(99)00025-1
15. Seo, M.H.; Shin, Y. Dynamic panels with threshold effect and endogeneity. Journal of Econometrics 2016, 195, 169–186. https://doi.org/10.1016/j.jeconom.2016.03.005
16. Hoerl, A.E.; Kennard, R.W. Ridge regression: Biased estimation for nonorthogonal problems. Technometrics 1970, 12, 55–67. https://doi.org/10.1080/00401706.1970.10488634
17. Arellano, M.; Bover, O. Another look at the instrumental variable estimation of error-components models. Journal of Econometrics 1995, 68, 29–51. https://doi.org/10.1016/0304-4076(94)01642-D
18. Blundell, R.; Bond, S. Initial conditions and moment restrictions in dynamic panel data models. Journal of Econometrics 1998, 87, 115–143. https://doi.org/10.1016/S0304-4076(98)00009-8
19. Anderson, K.; Kurzweil, M.; Martin, W.; Sandri, D.; Valenzuela, E. Measuring distortions to agricultural incentives, revisited. World Trade Review 2008, 7, 675–704. https://doi.org/10.1017/S1474745608004011
20. Miranda, M.J.; Farrin, K. Index insurance for developing countries. Applied Economic Perspectives and Policy 2012, 34, 391–427. https://doi.org/10.1093/aepp/pps031

21. Goodwin, B.K.; Smith, V.H. What harm is done by subsidizing crop insurance? American Journal of Agricultural Economics 2013, 95, 489–497. https://doi.org/10.1093/ajae/aas092
22. Ricker-Gilbert, J.; Jayne, T.S.; Chirwa, E. Subsidies and crowding out: A double-hurdle model of fertilizer demand in Malawi. American Journal of Agricultural Economics 2011, 93, 26–42. DOI: 10.1093/ajae/aaq122
23. Standing Committee of the National People's Congress. Cultivated Land Occupation Tax Law of the People's Republic of China. 2018. Available online: http://www.npc.gov.cn/englishnpc/c2759/c23934/202009/t20200929_384044.html
24. National People's Congress. Value-Added Tax Law of the People's Republic of China. 2024. Available online: https://npcobserver.com/legislation/value-added-tax-law/
25. Yi, F.; Sun, D.; Zhou, Y. Grain subsidy, liquidity constraints and food security: Impact of the grain subsidy program on the grain-sown areas in China. Food Policy 2015, 50, 114–124. https://doi.org/10.1016/j.foodpol.2014.10.009
26. Fan, P.; Mishra, A.K.; Feng, S.; Su, M.; Hirsch, S. The impact of China's new agricultural subsidy policy on grain crop acreage. Food Policy 2023, 118, 102472. https://doi.org/10.1016/j.foodpol.2023.102472
27. Hou, D.; Wang, X. How does agricultural insurance influence grain production scale? An income-mediated perspective. Frontiers in Sustainable Food Systems 2025, 9, 1524874. https://doi.org/10.3389/fsufs.2025.1524874
28. Chen, D.; Fang, X. Social assistance and non-agriculture employment in rural China: Evidence from the Rural Minimum Living Security (Rural Dibao). Humanities and Social Sciences Communications 2025, 12, 120. https://doi.org/10.1057/s41599-025-04443-5
29. Restuccia, D.; Rogerson, R. Policy distortions and aggregate productivity with heterogeneous establishments. Review of Economic Dynamics 2008, 11, 707–720. https://doi.org/10.1016/j.red.2008.05.002
30. Dennis, B.N.; İşcan, T.B. Agricultural distortions, structural change, and economic growth: A cross-country analysis. American Journal of Agricultural Economics 2011, 93, 885–905. https://doi.org/10.1093/ajae/aar011
31. Gollin, D.; Lagakos, D.; Waugh, M.E. Agricultural productivity differences across countries. American Economic Review 2014, 104, 165–170. DOI: 10.1257/aer.104.5.165
32. Louhichi, K.; Ricome, A.; Gomez y Paloma, S. Impacts of agricultural taxation in Sub-Saharan Africa: Insights from agricultural produce cess in Tanzania. Agricultural Economics 2022, 53, 671–686. https://doi.org/10.1111/agec.12704
33. Chang, J.J.; Chen, B.L.; Hsu, M. Agricultural productivity and economic growth: Role of tax revenues and infrastructures. Southern Economic Journal 2006, 72, 891–914. https://doi.org/10.1002/j.2325-8012.2006.tb00744.x
34. Brunetti, I.; Tidball, M.; Couvet, D. Relationship between biodiversity and agricultural production. Natural Resource Modeling 2019, 32, e12204. https://doi.org/10.1111/nrm.12204
35. Mushtaq, S.; Khan, S.; Dawe, D.; Hanjra, M.A.; Hafeez, M.; Asghar, M.N. Evaluating the impact of tax-for-fee reform (Fei Gai Shui) on water resources and agriculture production in the Zhanghe Irrigation System, China. Food Policy 2008, 33, 576–586. https://doi.org/10.1016/j.foodpol.2008.04.004
36. OECD. Agricultural Policy Monitoring and Evaluation 2013: OECD Countries and Emerging Economies; OECD Publishing: Paris, 2013. https://doi.org/10.1787/agr_pol-2013-en
37. Hennessy, D.A. The production effects of agricultural income support policies under uncertainty. American Journal of Agricultural Economics 1998, 80, 46–57. https://doi.org/10.2307/3180267
38. Kirwan, B.E. The incidence of U.S. agricultural subsidies on farmland rental rates. Journal of Political Economy 2009, 117, 138–164. https://doi.org/10.1086/598688
39. Zou, B.; Mishra, A.K.; Luo, B. Grain subsidy, off-farm labor supply and farmland leasing: Evidence from China. China Economic Review 2020, 62, 101293. https://doi.org/10.1016/j.chieco.2019.04.001
40. Hansen, B.E. Sample splitting and threshold estimation. Econometrica 2000, 68, 575–603. https://doi.org/10.1111/1468-0262.00124
41. Card, D.; Lee, D.S.; Pei, Z.; Weber, A. Inference on causal effects in a generalized regression kink design. Econometrica 2015, 83, 2453–2483. https://doi.org/10.3982/ECTA11224
42. Ganong, P.; Jäger, S. A permutation test for the regression kink design. Journal of the American Statistical Association 2018, 113, 494–504. https://doi.org/10.1080/01621459.2017.1328356

43. Calonico, S.; Cattaneo, M.D.; Titiunik, R. Robust nonparametric confidence intervals for regression-discontinuity designs. Econometrica 2014, 82, 2295–2326. https://doi.org/10.3982/ECTA11757
44. Calonico, S.; Cattaneo, M.D.; Farrell, M.H.; Titiunik, R. Regression discontinuity designs using covariates. Review of Economics and Statistics 2019, 101, 442–451. https://doi.org/10.1162/rest_a_00760
45. Chiang, H.D.; Sasaki, Y. Causal inference by quantile regression kink designs. Journal of Econometrics 2019, 210, 405–433. https://doi.org/10.1016/j.jeconom.2019.02.005
46. Yang, L.; Zhang, C.; Lee, C.; Chen, I.P. Panel kink threshold regression model with a covariate-dependent threshold. The Econometrics Journal 2021, 24, 462–481. https://doi.org/10.1093/ectj/utaa035
47. Breusch, T.S.; Pagan, A.R. The Lagrange multiplier test and its applications to model specification in econometrics. The Review of Economic Studies 1980, 47, 239–253. https://doi.org/10.2307/2297111
48. Pesaran, M.H. General diagnostic tests for cross-sectional dependence in panels. Empirical Economics 2021, 60, 13–50. https://doi.org/10.1007/s00181-020-01875-7
49. Baltagi, B.H.; Feng, Q.; Kao, C. A Lagrange multiplier test for cross-sectional dependence in a fixed effects panel data model. Journal of Econometrics 2012, 170, 164–177. https://doi.org/10.1016/j.jeconom.2012.04.004
50. Levin, A.; Lin, C.F.; Chu, C.S.J. Unit root tests in panel data: Asymptotic and finite-sample properties. Journal of Econometrics 2002, 108, 1–24. https://doi.org/10.1016/S0304-4076(01)00098-7
51. Im, K.S.; Pesaran, M.H.; Shin, Y. Testing for unit roots in heterogeneous panels. Journal of Econometrics 2003, 115, 53–74. https://doi.org/10.1016/S0304-4076(03)00092-7
52. Caner, M.; Hansen, B.E. Instrumental variable estimation of a threshold model. Econometric Theory 2004, 20, 813–843. doi:10.1017/S0266466604205011
53. Julien Fouquau. Threshold effects in Okun's law: a panel data analysis. Economics Bulletin, 2008, vol. 5, n° 33, pp. 1-14. hal-00565477
54. Kremer, S.; Bick, A.; Nautz, D. Inflation and growth: New evidence from a dynamic panel threshold analysis. Empirical Economics 2013, 44, 861–878. https://doi.org/10.1007/s00181-012-0553-9
55. Hansen, L.P. Large sample properties of generalized method of moments estimators. Econometrica: Journal of the Econometric Society 1982, 1029–1054. https://doi.org/10.2307/1912775