Changing Diets in China's Cities: Empirical Fact or Urban Legend?

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Abstract

China's economic reforms, which began in 1978, resulted in remarkable income growth, and urban Chinese consumers have responded by dramatically increasing their consumption of meat, other livestock products, and fruits and by decreasing consumption of grain-based foods. Economic prosperity, a growing openness to international markets, and domestic policy reforms have changed the food marketing environment for Chinese consumers and may have contributed to shifts in consumer preferences. The objective of this paper is to uncover evidence of structural change in food consumption among urban residents in China. Both parametric and nonparametric methods are used to test for structural change in aggregate household data from 1981 to 2004. The tests provided a reasonably clear picture of changing food consumption over the study period.

Keywords: China, demand models, food consumption, nonparametric analysis, parametric tests, structural change.

DIETARY STRUCTURAL CHANGE IN CHINA'S CITIES: EMPIRICAL FACT OR URBAN LEGEND?

Over the last two decades, urban Chinese consumers have dramatically increased their consumption of meat, other livestock products, and fruits and have decreased consumption of grain-based foods. China's per capita grain consumption declined from 145.44 kg in 1981 to 78.18 kg in 2004 in urban areas, whereas the per capita consumption of meats, fruits, and aquatic products increased respectively from 20.52 kg, 23.04 kg, and 7.26 kg in 1981 to 29.22 kg, 56.45 kg, and 12.48 kg in 2004 (CNBS, *China Statistical Yearbook,* various). As shown in Figure 1, the share of annual per capita consumption of grain has decreased significantly, from 23% to 9%, in the last 20 years while there have been some increases in the shares of fish, fruits, and other foods. As these significant changes in food consumption patterns in urban china are noted, it is natural to ask, Are consumer responses to price changes and income growth entirely responsible for the transformation in food consumption in urban China, or are there structural changes in China's food demand?

In this study, we define structural change as a shift in the economic relationship in food demand, which ultimately presents as a shift in consumers' preferences. In economic analysis, this broadly defined food demand structural change shows through a changing set of parameters. A structural change or preference change in food demand can be induced by many factors, including changes in consumer tastes, health information, policy, and market structure. As China has gone through so many changes, such as policy reforms, economic prosperity (per capita GDP increased from 117.5 yuan in 1981 to 760 yuan in 2004, with 1978=100), development of infrastructure, and external influences with its increasing openness to the world, it is highly possible that these factors have exerted some impacts on consumers' food demand. Several authors have noted significant changes in food consumption patterns in urban China (Hsu et al., 2001; Shono et al., 2000; Guo et al., 2000; Huang and Bouis, 1996); however, most support their conjectures with evidence based on trends in the data or estimates of expenditure and income elasticities. While trends and income elasticities are informative, they do not provide convincing evidence of a shift in preferences or economic relationship in food demand. Indeed, it is possible that consumer responses to price changes and income growth under a stable set of preferences may be entirely responsible for the changes in food consumption in urban China.

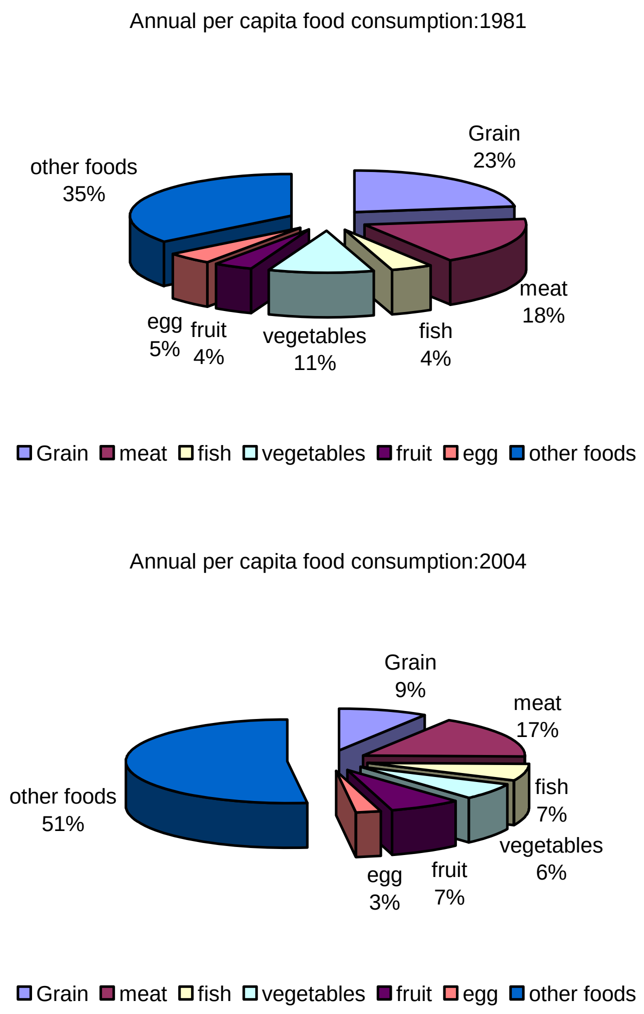

Figure 1. Structure of Food Consumption in Urban China, 1981 versus 2004

Testing structural changes in the demand for food in urban China is useful, not least because such empirical studies are lacking in the literature. This issue is of fundamental importance to policymakers and the food industry. In the presence of significant structural changes, price-based strategies or income policy would exert different effects on demand. Especially as China is becoming more and more important in world markets, even minor shifts in consumer preferences can exert significant effects on international agricultural markets. Therefore, it is necessary to examine the structural changes in urban Chinese food demand.

Understanding whether or not observed shifts in food consumption in China are actual changes in its demand structure or simply the natural progression through a stable preference set requires empirical techniques that isolate preference changes. This, then, becomes the objective of this study: to provide a test for shifts in preferences over time for urban Chinese food consumption. This paper extends our knowledge of Chinese food consumption by utilizing empirical methods to investigate dietary structural change. The parametric procedure described by Moschini and Meilke (1989) is used to estimate and test for shifts in demand parameter values over time. Changing parameter values can be viewed as evidence of preference change.

The next section briefly describes major policy and food market changes that have occurred in China over the last three decades. The background section is followed by a description of the methodology used to perform parametric tests of structural change and the results of those tests. Nonparametric tests are performed on the same data, and these test results are compared with the outcomes of parametric analysis. We conclude with a summary of our findings and suggestions for further research.

Background

Since 1978, China has gradually transformed its economy from a highly centralized, planned economy to a more market-oriented economy in a process that has been tightly controlled by the government. A series of reforms of the government's administrative system, agricultural policy, state-owned enterprises, investment regulations, fiscal and taxation policies, and financial system have fuelled the growth of China's economy and generated impressive economic development. Privatization of food production, procurement, and marketing dramatically increased the quantity and availability of food in urban China, creating new consumption opportunities. In 1978, the share of domestic trade under government price controls was 97%, 100%, and 92.6%, respectively, for total retail sales, sales of industrial goods, and total purchases of farm and sideline products. Moreover, China's domestic trade was mainly conducted by state-owned enterprises. A watershed shift in agricultural policy occurred in 1981, when China's government adopted a decentralized agricultural production system based on household units called the household responsibility system (HRS). Following the adoption of the HRS, China's agricultural production boomed, and the availability of agricultural produce and food greatly increased.

By 1984, the number of commodities subject to state procurement decreased dramatically from 113 in 1981 to 60, and price controls on several important non-staple foods—including pork, eggs, sugar, and vegetables—were lifted in late 1984. One year later, the government officially adopted a dual-track marketing system, allowing sale and distribution of food products by private firms in addition to the state-owned system. The emerging private food marketing chains provided a wider range of consumption choices for China's consumers. With increasing urban household income and abundant supply of farm produce, the free market soon became the dominant force in the dual-market system.

The opening of China's food processing and retailing sector to foreign direct investment (FDI) has facilitated the rapid modernization of China's food processing and distribution systems and created an environment that fosters food product innovation. FDI in China became possible in 1979, but interest from foreign investors was modest until China expressed interest in re-entering negotiations for membership in the World Trade Organization in 1994. From 1990 to 1997, annual FDI flows into China grew tenfold. In 1995, foreign joint ventures in the food processing sector accounted for 22% of the sector's total value of output and 31% of gross profits (Wei and Cacho, 2001). In addition to the food processing sector, substantial FDI has been targeted at the retail food sector, and the 1990s saw supermarkets rise to become a leading retail format for food products in urban areas (Hu et al., 2004). As foreign firms and products penetrated Chinese food markets, domestic firms responded by emulating and adapting foreign product designs, quality standards, and marketing strategies to better fit the tastes of Chinese consumers (Wei and Cacho, 2001). The result has been the advent of a plethora of new food products offered in modern retail formats, which has facilitated significant changes in consumer shopping behaviors (Veeck and Veeck, 2000).

Moreover, with higher education, influences from other countries, and changes in the labor market, urban Chinese people have changed their preference for food with consideration for food safety, health, convenience, or fashion. For example, according to the national population census, the rate of illiteracy of Chinese people has decreased from 22.81% in 1982 to 6.72% in 2000. And the percentage of people with a degree from a junior college and above increased from 0.615% in 1982 to 3.61% in 2000. With all of these significant changes, it is appealing to investigate whether there have been structural changes in consumers' consumption preferences during the last two decades in urban China.

Given the dynamic nature of China's food markets, assessing the relative importance of these major developments on consumer choices is not a trivial task. The remainder of the paper is devoted to examining the empirical evidence of changing structures in urban Chinese households' food demand and the impacts of these changes on diets.

Parametric Tests for Structural Change

Preference changes are reflected analytically as a change in the shape of individual utility functions, which may be detected empirically as parameter instability in the demand function. In the parametric tests for structural change, it is necessary to choose a functional form to approximate the real demand satisfactorily. During the last decade, the almost ideal demand system (AIDS) and the Rotterdam model have been adopted by agricultural economists as the demand systems in most applications. While these two models are similar in many respects, such as having identical data requirements and being (second-order) locally flexible, linear in parameters, and parsimonious with respect to numbers of parameters, they lead to different results in some applications (Alston and Chalfant, 1993). Therefore, it is necessary to make the right choice of functional form before we go to the test of the structural change. Alston and Chalfant initiated a test to distinguish econometrically between the AIDS and Rotterdam models. Their method, however, was criticized by LaFrance (1998) as being biased and inconsistent because of problems in the assumption of the error terms in the differenced form, the assumption of heteroskedasticity, simultaneity, and the use of the Stone price index in AIDS. As a result, we utilize the method suggested by LaFrance (1998) to overcome the deficiencies in Alston and Chalfant's method.

Choice of Functional Form: AIDS versus Rotterdam Model

Based on the method of LaFrance, we develop a compound model to test the Rotterdam model against the first-differenced AIDS model. We choose the form for the Rotterdam model as the one used in Alston and Chalfant's study:

$$\begin{aligned}\overline{w}_{i,t}\Delta\ln q_{i,t} &= \sum_{i=1}^{n}\gamma_{ij}\Delta\ln p_{j,t} + \beta_i\sum_{i=1}^{n}\overline{w}_{i,t}\Delta\ln q_{i,t} \\ &= \sum_{i=1}^{n}\gamma_{ij}\Delta\ln p_{j,t} + \beta_i\sum_{i=1}^{n}\Delta\ln\frac{y}{P^*} \\ &= \sum_{i=1}^{n}\gamma_{ij}\Delta\ln p_{j,t} + \beta_i(\Delta\ln y - \sum_{j=1}^{n}\overline{w}_{j,t}\cdot\Delta\ln p_{j,t})\end{aligned} \tag{1}$$

where $w_{i,t}$ is the budget share of good *i*, $q_{i,t}$ is the quantity of good *i*, $\overline{w}_i$ is an average of $w_{i,t}$ and $w_{i,t-1}$, $p_{i,t}$ is the price of good *i*, y_t is the total expenditure, *n* is the number of goods in the system, and γ_{ij} and β_i are the parameters. Homogeneity, adding-up, and symmetry require the following parameter restrictions:

$$\sum_j \beta_j = 1,\ \sum_j \gamma_{ij} = \sum_i \gamma_{ij} = 0,\ \gamma_{ij} = \gamma_{ji}\,.$$

For the first-differenced AIDS model we use the following format, which is also used by Alston and Chalfant:

$$
\begin{aligned}
\Delta w_{i,t} &= \sum_{i=1}^{n} \gamma_{ij} \Delta \ln p_{j,t} + \beta_i \Delta \ln \frac{y}{P} \\
&= \sum_{i=1}^{n} \gamma_{ij} \Delta \ln p_{j,t} + \beta_i (\Delta \ln y - \Delta \ln P) \\
&= \sum_{i=1}^{n} \gamma_{ij} \Delta \ln p_{j,t} + \beta_i [\Delta \ln y - (\sum_{i=1}^{n} w_{i,t} \Delta \ln p_{i,t} + \sum_{i=1}^{n} \Delta w_{i,t} \cdot \ln p_{i,t} + \sum_{i=1}^{n} \Delta w_{i,t} \cdot \Delta \ln p_{i,t}) \\
&\approx \sum_{i=1}^{n} \gamma_{ij} \Delta \ln p_{j,t} + \beta_i (\Delta \ln y - \sum_{i=1}^{n} w_{i,t} \Delta \ln p_{i,t})
\end{aligned}
\tag{2}
$$

To account for the lack of invariance of Stone's price index to units of measurement, prices are scaled by their samples means in the tests (LaFrance; Moschini).

The compound model to test the Rotterdam model against the first-differenced AIDS model is written in the following implicit form (LaFrance):

$$
\begin{aligned}
&\varepsilon_{i,t} = (1-\lambda) \cdot \frac{1}{2} [(\frac{p_{i,t}}{y_t}) q_{i,t} + w_{i,t-1}] \cdot \log(\frac{q_{i,t}}{y_t}) + \lambda \cdot [(\frac{p_{i,t}}{y_t}) q_{i,t} - w_{i,t-1}] \\
&- \sum_{j=1}^{n-1} \gamma_{ij} \cdot \log(\frac{p_{j,t} \cdot p_{n,t-1}}{p_{j,t-1} \cdot p_{n,t}}) - \beta_i \{\log(\frac{y_t}{y_{t-1}}) - \frac{1}{2} \sum_{j=1}^{n} [(\frac{p_{j,t}}{y_t}) q_{j,t} + w_{j,t-1}] \\
&\cdot \log(\frac{p_{j,t} \cdot p_{n,t-1}}{p_{j,t-1} \cdot p_{n,t}}) - \frac{1}{2}(1 + w_{n,t-1}) \cdot \log(\frac{p_{n,t}}{p_{n,t-1}})\}
\end{aligned}
\tag{3}
$$

where ε_t is assumed to be i.i.d. N(0, Σ). The test of λ=0 is a hypothesis test of the Rotterdam model against the AIDS model. And similarly, the implicit form of the compound model for a test of the AIDS model against the Rotterdam model is as follows:

$$
\begin{aligned}
&\varepsilon_{i,t} = (1-\lambda) \cdot [(w_{i,t} - w_{i,t-1}] + \lambda \cdot \frac{1}{2} [(w_{i,t} + w_{i,t-1}] \cdot \log(\frac{q_{i,t}}{q_{i,t-1}}) \\
&- \sum_{j=1}^{n} \gamma_{ij} \cdot \log(\frac{p_{j,t} \cdot p_{n,t-1}}{p_{j,t-1} \cdot p_{n,t}}) - \beta_i \{\log(\frac{y_t}{y_{t-1}}) - \sum_{j=1}^{n} w_{j,t} \cdot \log \frac{p_{j,t}}{p_{n,t}} \\
&+ \sum_{j=1}^{n} w_{j,t-1} \cdot \log \frac{p_{j,t-1}}{p_{n,t-1}} - \log(\frac{p_{n,t}}{p_{n,t-1}})\}
\end{aligned}
\tag{4}
$$

Correspondingly, the test of λ=0 is a hypothesis test of the AIDS against the Rotterdam model.

Data

Annual data from 1981 to 2004 for per capita consumption, expenditures, and retail prices are obtained from the *Chinese Urban Household Income and Expenditure Survey* and various issues of the *China Statistical Yearbook* (CNBS, various). The data set contains seven food groups: grain, meat, eggs, fish, vegetables, fruit, and other foods. The consumer price index was used as the price of other foods. Expenditures on other foods were recalculated by deducting food expenditures on the other six commodity groups from total food expenditure. The aggregate quantities were calculated by dividing group expenditures by the price index. All prices and income were normalized by their sample mean.

Result of Model Selection

The results of the testing of the Rotterdam against AIDS model and AIDS against Rotterdam model are presented in Table 1. The test based on equation (3) shows that we cannot reject the null hypothesis of λ equal to zero, which suggests that the Rotterdam model is correct. The test based on equation (4) rejects the null hypothesis of λ equal to zero and thus rejects the null hypothesis that the AIDS model is correct. Based on these two results, we decide that the Rotterdam model in equation (1) is more appropriate for our analysis.[1]

Table 1. Hypothesis Tests of the Rotterdam versus AIDS Models

	parameter	standard error	t-ratio
Rotterdam versus AIDS			
LAM1	0.0383	0.0320	1.1958
AIDS versus Rotterdam			
LAM1	0.4799	0.1250	3.8376

[1] To be comparable, we still conducted all the tests on the Rotterdam model to the first difference AIDS model, and we report the results briefly in the footnotes.

Test of Structural Change

In food demand, an abrupt structural change will occur and can be captured by a binary variable. Sometimes, however, adjustment may take place gradually before the market settles to a new equilibrium (Peterson and Chen, 2005). Consequently, incorporating a continuous shift variable (Brester and Schroeder, 1995) or a time transition function (Ohtani and Katayama, 1986; Moschini and Meilke, 1989) into the econometric model is more plausible.

We utilize the method suggested by Moschini and Meilke (1989) into the Rotterdam model in equation (1), given the model selection results. Following Moschini and Meilke, structural change can be characterized by allowing the set of parameters of the demand system to change over time. With a common time path h_t, the Rotterdam model is re-parameterized in equation (5) to capture time-varying parameter shifts:

$$\overline{w}_{i,t}\Delta\ln q_{i,t} = \sum_{i=1}^{n}(\gamma_{ij} + a_{ij}h_t)\Delta\ln p_{j,t} + (\beta_i + b_i h_t)\sum_{i=1}^{n}\overline{w}_{i,t}\Delta\ln q_{i,t}\,. \tag{5}$$

Additional parametric restrictions in the structural change model associated with homogeneity, adding up, and symmetry are $\sum_{i=1}^{n}a_{ij} = \sum_{j=1}^{n}a_{ij} = 0,\ a_{ij} = a_{ji}$, and $\sum_{i=1}^{n}b_i = 0$. To approximate the actual shape of the time path, h_t is constructed as the piece-wise linear function defined in equation (5) (Ohtani and Katayama, 1986; Moschini and Meilke, 1989):

$$\begin{aligned} h_t &= 0, && \text{for } t = 1,\ldots.,\tau_1; \\ h_t &= (t-\tau_1)/(\tau_2-\tau_1), && \text{for } t = \tau_1+1,\ldots,\tau_2-1; \\ h_t &= 1, && \text{for } t = \tau_2,\ldots,T. \end{aligned} \tag{6}$$

The value τ_1 is the endpoint of the first regime and τ_2 is the starting point of the second regime ($\tau_1 < \tau_2$). The difference between τ_1 and τ_2 defines the transition path. If $\tau_2 = \tau_1 + 1$, the structural change is abrupt; otherwise, the change is gradual. The compensated and uncompensated price elasticities are computed from the model parameters and the value of the transition function for each period as

$$\varepsilon_{ijt} = \frac{\gamma_{ij} + a_{ij}h_t}{w_{it}}, \qquad (7)$$
$$\varepsilon^*_{ijt} = \varepsilon_{ijt} - w_{jt}\varepsilon_{iyt}$$

respectively, with income elasticities at time *t* computed as

$$\varepsilon_{iyt} = \frac{\beta_i + b_i h_t}{w_{it}}. \qquad (8)$$

In model (5), a test of the hypothesis of no structural change is equivalent to a test of the hypothesis that the time path parameters (a_{ij} and b_i) are all equal to zero.

Results from Structural Change Tests

The Rotterdam model in equation (5) was estimated using the maximum likelihood method. In estimating both models, the equation for other foods was omitted to avoid singularity problems.

The Rotterdam model was estimated for each combination of τ_1 and τ_2 in the feasible ranges, and the likelihood function was checked to find the combination that yielded the highest function value. The structural change point (τ_1, τ_2) resulting in the maximum values for the likelihood function is τ_1=1983 and τ_2=1989, as shown in Table 2.[2] Based on likelihood ratio tests, several other combinations of τ_1 and τ_2 failed to reject the null hypothesis of structural change. And all these sub-optimal structural change points focus on the same period. The results indicate a gradual shift in preferences in j1980s that corresponds to the period of time when the dual-track marketing system was established and some agricultural policies (such as HRS) were reformed.

Table 2. Maximum Likelihood Structural Change Points

	Structural Change Periods
Optimal Points	(83,89)
Additional Points	[(81-85),89],[(81-85),91]

Note: The numbers in the first set of parentheses are possible values for τ1, and the numbers in the second set of parentheses are possible values for τ2. These combinations are structural change points that cannot be rejected at the 0.05 significance level.

[2] The structural change point found by the first-difference AIDS model is τ_1=1982 and τ_2=1990, which nearly perfectly overlapped the period of structural change found by the Rotterdam.

The results from the structural change test suggest that policy changes associated with the development of free markets have been the most important agents of structural change in Chinese diets.

The maximum likelihood parameter estimates for the Rotterdam model are listed in Table 3. The R^2 of each single equation indicates that the fit of the model is good. The Durbin-Watson statistics show no evidence of autocorrelation in the residuals. In all equations, there is at least one time variable parameter that is significant, indicating that there are structural changes in all kinds of foods.

To further investigate the nature and significance of the structural change, conditional on the optimal combination of (τ_1, τ_2), we conducted likelihood ratio tests for the hypothesis of constancy of the parameter vector over time, i.e., whether or not the coefficients for the time path variables are equal to zero. The results are reported in Table 4. The hypothesis of no structural change in the full set of parameters is rejected at the 5% significance level, suggesting that a constant set of parameters does not adequately characterize urban consumer behavior in China within the assumed model.[3] Some structural change over the period must be incorporated. Price, income, and intercept structural change parameters are also tested to shed light on the nature of the preference change. All of the tests reject the hypothesis of no structural change at the 5% significance level. Thus, joint tests suggest that shifts in preferences of urban Chinese households involve consumers' response to both price and volume changes.

The average Marshallian price and expenditure elasticities calculated at the mean shares for the Rotterdam model are reported in Table 5. Standard errors for the elasticities are computed using the delta method (Green et al., 1987). Most own-price elasticities were negative before and after the optimal structural change except the one for grain before the structural change, which was not significant. And after the structural change, all food demands became less elastic except for fruit. Moreover, meat, fish, vegetables, and eggs changed from price elastic

[3] The hypothesis of no structural change in the full set of parameters is also rejected at the 5% significance level in the first-difference AIDS model.

Table 3. Maximum Likelihood Parameter Estimates for Rotterdam Model with Optimal Structural Change Points for Seven Food Group at (1983, 1989)

	Grain	Meat	Fish	Veg.	Fruit	Eggs	Other
Volume Index	0.0209	0.0068	-0.0632	-0.0907	0.0246	0.2266*	0.8424*
β_i	(0.1021)	(0.1306)	(0.0721)	(0.0481)	(0.0381)	(0.0250)	(0.1684)
	-0.0156	0.2610	0.1856*	0.0794	0.0775	-0.2208*	-0.3266
b_i	(0.1067)	(0.1383)	(0.0813)	(0.0514)	(0.0464)	(0.0263)	(0.1877)
Grain	-0.2433						
γ_{ij}	(0.1310)						
	0.2445						
a_{ij}	(0.1315)						
Meat	0.1223	-0.7393*					
γ_{ij}	(0.0990)	(0.1636)					
	-0.1347	0.6215*					
a_{ij}	(0.1010)	(0.1695)					
Fish	0.0560	0.3759*	-0.3008*				
γ_{ij}	(0.0642)	(0.0711)	(0.0435)				
	-0.0790	-0.3520*	0.2611*				
a_{ij}	(0.0651)	(0.0740)	(0.0486)				
Vegetables	0.0238	-0.2615*	0.2626*	-0.2493*			
γ_{ij}	(0.0535)	(0.0428)	(0.0275)	(0.0270)			
	-0.0349	0.2513*	-0.2401*	0.2138*			
a_{ij}	(0.0537)	(0.0440)	(0.0286)	(0.0274)			
Fruit	-0.0003	-0.0883*	0.1276*	-0.0487*	-0.0439*		
γ_{ij}	(0.0246)	(0.0306)	(0.0171)	(0.0131)	(0.0125)		
	-0.0031	0.1117*	-0.1557*	0.0707*	-0.0171		
a_{ij}	(0.0258)	(0.0328)	(0.0211)	(0.0146)	(0.0167)		
Eggs	-0.2214*	0.2895*	-0.0557*	0.0034	0.0267*	-0.0856*	
γ_{ij}	(0.0309)	(0.0228)	(0.0150)	(0.0152)	(0.0064)	(0.0207)	
	0.2262*	-0.2932*	0.0629*	0.0051	-0.0234*	0.0594*	
a_{ij}	(0.0309)	(0.0234)	(0.0154)	(0.0153)	(0.0072)	(0.0209)	
Others	0.2633	0.2831	-0.4411*	0.2738*	0.0255	0.0273	-0.4320
γ_{ij}	(0.1529)	(0.1837)	(0.0928)	(0.0703)	(0.0435)	(0.0421)	(0.2865)
	-0.2210	-0.1864	0.4786*	-0.2709*	0.0191	-0.0215	-0.2020
a_{ij}	(0.1560)	(0.1925)	(0.0977)	(0.0720)	(0.0484)	(0.0432)	(0.3010)
R^2	0.2038	0.8991	0.0986	0.4130	0.6451	0.9236	
DW	1.4665	1.1146	1.8272	1.7681	1.8451	1.7919	

Note: Asymptotic standard errors are reported in the parenthesis. The second row of parameters is those of time variables. The parameters for the seventh equation are recovered using delta method. * indicates significance at 5% level.

Table 4. Likelihood Ratios for Structural Change Tests for Rotterdam Model

Hypothesis	**Restrictions**	**Likelihood Ratio**	$\chi^2_{0.05}$
No Structural Change in:			
All parameters	27	137.8345	40.1133
Price parameters	21	120.9985	32.6706
Volume Index parameters	6	43.1565	12.5916

Table 5. Average Marshallian Price and Income Elasticities for the Rotterdam Model

	Grain	Meat	Fish	Vegetables	Fruit	Eggs	Other Foods
				Before Structural Change			
Grain	-1.1652	0.5574	0.2590	0.1017	-0.0070	-1.0464*	1.2020
	(0.6844)	(0.4184)	(0.3095)	(0.2568)	(0.1243)	(0.1494)	(0.7013)
Meat	0.6645	-4.0706*	2.0646*	-1.4413*	-0.4876*	1.5890*	1.6438
	(0.5749)	(0.8078)	(0.4125)	(0.2410)	(0.1738)	(0.1295)	(1.0964)
Fish	1.5828	8.8314*	-6.7946*	6.1377*	2.9880*	-1.1937*	-10.1108*
	(1.5972)	(1.4641)	(1.0295)	(0.6334)	(0.4162)	(0.3551)	(2.1687)
Vegetables	0.4103	-2.3308*	2.5362*	-2.2810*	-0.4164*	0.0775	2.8674*
	(0.5678)	(0.3778)	(0.2655)	(0.2736)	(0.1360)	(0.1432)	(0.6353)
Fruit	-0.1021	-1.6984*	2.3156*	-0.9390*	-0.8287*	0.4645*	0.3376
	(0.4737)	(0.5289)	(0.3242)	(0.2490)	(0.2470)	(0.1220)	(0.8390)
Egg	-5.1687*	4.7585*	-1.2576*	-0.3913	0.2738*	-1.8666*	-0.6921
	(0.6459)	(0.4290)	(0.2889)	(0.3232)	(0.1370)	(0.3884)	(0.7740)
Other Foods	0.2199	0.4070	-1.4418*	0.5084*	-0.0594	-0.0072	-1.5378*
	(0.4759)	(0.4723)	(0.2661)	(0.2203)	(0.1359)	(0.1233)	(0.2346)
Expenditure	0.0985	0.0375	-1.4409	-0.8632	0.4506	4.3440*	2.5027*
	(0.4802)	(0.7178)	(1.6426)	(0.4576)	(0.6964)	(0.4788)	(0.4720)
				After Structural Change			
Grain	0.0051	-0.1153	-0.2024*	-0.1009	-0.0328	0.0396	0.3608*
	(0.1132)	(0.1115)	(0.1008)	(0.0719)	(0.0723)	(0.0243)	(0.1349)
Meat	-0.2442*	-0.9338*	0.0322	-0.2033*	0.0271	-0.0773*	-0.1146
	(0.0798)	(0.1413)	(0.0875)	(0.0509)	(0.0584)	(0.0209)	(0.1633)
Fish	-0.5464*	0.0327	-0.7080*	0.1589	-0.5380*	0.0390	-0.2424
	(0.1814)	(0.2105)	(0.2751)	(0.1258)	(0.1689)	(0.0552)	(0.2546)
Vegetables	-0.1021	-0.0854	0.2426*	-0.3578*	0.2368*	0.0926*	0.0911
	(0.0832)	(0.0783)	(0.0858)	(0.0662)	(0.0700)	(0.0251)	(0.0993)
Fruit	-0.2194	0.0759	-0.5052*	0.1757	-0.9841*	-0.0084	-0.0105
	(0.1287)	(0.1377)	(0.1629)	(0.0992)	(0.1657)	(0.0487)	(0.1782)
Egg	0.1097	-0.1264	0.1828	0.2124*	0.0758	-0.7061*	0.0962
	(0.0759)	(0.0832)	(0.0948)	(0.0660)	(0.0930)	(0.0494)	(0.1060)
Other Foods	-0.0334	0.0161	0.0060	-0.1029*	0.0201	-0.0293*	-1.0377
	(0.0509)	(0.0778)	(0.0508)	(0.0298)	(0.0387)	(0.0138)	(0.0973)
Expenditure	0.0459	1.5138*	1.8041*	-0.1178	1.4761*	0.1557	1.1612*
	(0.2069)	(0.1954)	(0.4018)	(0.1522)	(0.2891)	(0.1829)	(0.1261)

Note: Structural change point is (1983, 1989). Asymptotic standard errors are reported in parentheses.
* indicates significance at 5% level.

to price inelastic. Before the structural change, only eggs and other foods had significant expenditure elasticities and showed to be luxury goods. After the structural change, eggs changed from a luxury good to a necessity. And conversely, meat, fish, and fruit changed from necessities to luxuries. Grains, vegetables, and eggs expenditure elasticities were not significant after the structural change.

The expenditure elasticity estimates from both models fall within the range of estimates from other studies (Yen et al., 2004; Zhang and Wang, 2003; Gould, 2002; Liu and Chern, 2003; Gould and Dong, 2004; Wu et al., 1995). For comparison, we also estimated the elasticities for the Rotterdam models without considering the structural change. The elasticities of the models without structural change are generally less price and income elastic than the corresponding models with structural change (see appendix), which underscores the importance of testing and adjusting for structural change in empirical and applied analysis.[4]

Additional Test: Nonparametric Method

Because a functional form is assumed in the parametric approach, parametric tests are ultimately joint tests of the functional form used to perform the analysis. Rejection of the hypothesis of stable preferences is conditioned on the assumption that the test results are insensitive to the functional form chosen (Alston and Chalfant, 1991). This fact motivated the selection of the dynamic AIDS and the Rotterdam models in the tests above. As a further check on the robustness of the results described in the last section, we utilize the theory of revealed preference to conduct nonparametric tests for stable preferences, which do not depend on assumptions regarding the functional representation of preferences.

Nonparametric analysis of structural change is derived from the idea that a vector of prices and a corresponding vector of consumption bundles generated by consumers with stable preferences will satisfy the necessary and sufficient conditions for the data to be rationalized by a utility function. Building on the work of Samuelson (1948), Houthakker (1950), and Afriat (1967), Varian (1982) demonstrated that the generalized axiom of revealed preference (GARP) is a sufficient condition for utility maximization. GARP states that if a consumption bundle, x_j, is revealed preferred to another bundle, x, then x_j cannot cost less than x evaluated at the price vector associated with bundle x; otherwise, the data is not consistent

[4] Although showing some differences in values compared to those from Rotterdam model, the elasticities calculated from the first-difference AIDS model showed the same characteristics: all own-price elasticities were negative; food demands became less price elastic after the structural change; and without considering the structural change, the elasticities were generally less price and income elastic than those obtained from the corresponding models with structural change.

with utility-maximizing behavior (Varian, 1982). Finding one observation that violates GARP is technically sufficient to reject consistency of the data with utility maximization under stable preferences.

We apply the algorithm described by Varian (1982) for testing GARP to the data used for the parametric tests described in the previous section. The data set satisfies GARP for all observations. Thus, it would appear that there is no evidence of structural change. However, questions have been raised about the power of the GARP test, particularly when the real expenditures grow rapidly over time. Real food expenditures for urban Chinese consumers have increased an average of 3.3% annually since 1981. Income effects may mask shifts in the underlying preferences by causing each successive budget line to lie outside of the consumption set of the previous observation, despite relative price changes. In other words, the budget lines associated with two observed consumption bundles do not cross, making it impossible to identify a violation of stable preferences.

Income-Adjusted Tests

To improve the power of revealed preference tests, Chalfant and Alston (1988) suggest using prior information about income elasticities to adjust the expenditure data as a means of removing the effects of income growth from the analysis. By filtering out the income effects, the potential impacts of structural change may be observed in the residual data. Applying a similar concept, Sakong and Hayes (1993) argue that the impacts of shifts in consumer preferences could be isolated from income and price effects using the compensated demand curve. Given some reasonable range of income elasticities, any change in the quantity purchased from one time period to the next that cannot be explained by the price and income changes can be attributed to a change in tastes. They compute estimates of taste changes by solving a linear programming problem for the set of consumption bundles that minimizes cumulative taste changes and that satisfy convexity of preferences. The convexity condition is similar to the revealed preference condition discussed above. Income elasticities are endogenous in this model, but they are constrained by the Engel aggregation condition and by assumed upper and lower bounds.

We applied the Sakong and Hayes model to the seven-commodity aggregations. We incorporate the adjustments to the model suggested by Chalfant and Zhang (1997) to avoid dependence of the test results on scaling and price deflator choices. We attempted to select bounds that would include the majority of the estimates found in a brief survey of studies analyzing urban household consumption in China. Some sensitivity analysis was conducted to determine whether broadening the selected ranges would significantly alter the results. While the magnitudes of some taste changes did vary, the qualitative result did not change substantially. Table 6 displays the expenditure elasticity ranges selected for this study.

Table 6. Expenditure elasticity ranges by commodity

Commodity	Maximum	Minimum
Grain	1.3	0.0
Eggs	1.0	0.4
Fish	1.5	0.8
Meat	1.3	0.7
Fruits	1.5	0.6
Vegetables	1.2	0.6
Other	1.6	0.9

In addition to the bounds placed on expenditure elasticities, we placed bounds on the year-to-year change in expenditure elasticity values. Using the results from the Rotterdam models, we computed the average change in the expenditure elasticity values on an annual basis. For most commodities, expenditure elasticities changed by less than 0.15 in absolute value from one year to the next 95% of the time. All commodities changed by less than 0.2 in absolute value 95% of the time. Thus, we computed taste changes for each commodity set using three different assumptions regarding year-to-year expenditure elasticity changes: limited to 0.15, limited to 0.2, and no limit. Allowing expenditure elasticities to change by larger amounts on a year-to-year basis changed only a small number of structural change points.

The results from the nonparametric analysis support the findings from the parametric estimation. In particular, both approaches consistently identify structural change in the 1980s.[5] The cumulative taste changes measured in kilograms per person showed that the shifts in preferences increased consumption of fish and fruits in the 1980s at the expense of consumption of other foods. Changes in grain, meat, and vegetable consumption during the study period are entirely explained by income and price effects.

Conclusions

The objective of this paper was to uncover evidence of structural change in food consumption among urban residents in China. The battery of tests applied to data for the period from 1981 to 2004 provided a reasonably clear picture of changing food consumption. First, both parametric and nonparametric tests indicated that the 1980s was a period of structural change in food consumption in urban China. The introduction of the HRS and the dual-track marketing system greatly increased the availability of nonstaple foods in urban areas. The nonparametric results suggest that during this same period, consumer preferences shifted in favor of fruits and aquaculture products, increasing per capita consumption of each product by roughly 2 kg.

Second, foods that have long played a major role in urban Chinese diets did not show strong evidence of structural change. In particular, changes in grain, meat, and vegetable consumption can be largely explained by normal price and income effects. In contrast, fruits and fish, while not absent from traditional Chinese diets but having played a less important role in daily food consumption, were frequently identified in the tests as showing evidence of structural change. In terms of Pingali and Khwaja's (2004) stages of dietary development, the decline in grain consumption and the growth in meat and vegetable expenditures are consistent with the dietary diversification that comes with income growth. The increasing consumption of fish may be evidence of an expansion of consumer food purchases to include

[5] However, the nonparametric procedure finds more evidence of structural change in the latter half of the 1990s than does the parametric approach. Although the nonparametric methodology is not able to distinguish between structural change in income and price responses, it does provide a measure of quantity and expenditure change due to preference change because taste changes are computed in quantity terms.

goods that are part of the national diet but may not have been prominent in local or regional diets.

Finally, the parametric analysis indicates that the greatest changes in preferences occurred in consumers' responses to price changes. In particular, consumer demands became less price elastic. As incomes have risen, food choices have increased, and consumers' food preparation and shopping behaviors have changed, product attributes other than prices may be playing a greater role in consumption decisions. All of these findings have important implications for the analysis and forecast of urban Chinese food demand, which plays a critical role in the world market. With structural change in urban Chinese food demand, researchers who use standard price and income elasticities will fail to predict accurately changes in consumption over time.

This study was limited by the number of observations and degree of aggregation in our data set. Future research of this type would be best conducted using a panel of household data. Using a single cross-section, however, is not adequate to address the question of change over time. Cross-sectional data is useful for identifying the types of dietary change associated with the first stage of dietary diversification, which is driven by income growth. A single cross-section is conditioned on the marketing infrastructure, consumer information channels, and the array of products available at the time when the data are collected. If researchers desire empirical evidence of globalization or other drivers of preference change, time-series or panel data should be used to capture the impacts of consumers' changing market environment on purchasing decisions.

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Appendix

Table A1. Average Marshallian Price and Income Elasticities for Rotterdam Model with Seven Commodity Groups without Structural Change

	Grain	Meat	Fish	Vegetables	Fruit	Eggs	Other Foods
Grain	-0.0490	-0.0998	-0.1177	-0.1100	-0.0451	-0.0422	0.3281
	(0.0086)	(0.0253)	(0.0326)	(0.0280)	(0.0122)	(0.0101)	(0.0801)
Meat	-0.2175	-0.7485	0.0516	-0.1154	0.0068	0.0243	-0.2618
	(0.0449)	(0.0650)	(0.0104)	(0.0118)	(0.0122)	(0.0160)	(0.1148)
Fish	-0.3330	0.2428	-0.5853	0.2211	0.0308	0.1064	-0.4838
	(0.0925)	(0.0404)	(0.0997)	(0.0374)	(0.0114)	(0.0147)	(0.0632)
Vegetables	-0.1443	-0.0128	0.1781	-0.4105	0.0691	0.0693	0.0863
	(0.0215)	(0.0025)	(0.0302)	(0.0686)	(0.0124)	(0.0137)	(0.0065)
Fruit	-0.1563	0.1241	0.0342	0.0461	-0.4608	0.0110	-0.2816
	(0.0444)	(0.0298)	(0.0099)	(0.0127)	(0.0615)	(0.0075)	(0.0656)
Egg	-0.1778	0.2652	0.1886	0.1432	0.0358	-0.5118	-0.4286
	(0.0362)	(0.0850)	(0.0506)	(0.0488)	(0.0116)	(0.1411)	(0.1576)
Other Foods	-0.0963	-0.1769	-0.1227	-0.1230	-0.1125	-0.0908	-0.8538
	(0.0727)	(0.0677)	(0.0160)	(0.0415)	(0.0311)	(0.0326)	(0.0399)
Expenditure	0.1357	1.2605	0.8012	0.1648	0.6834	0.4854	1.5760
	(0.0366)	(0.1576)	(0.1490)	(0.0287)	(0.1016)	(0.1390)	(0.2722)