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.

CHANGING DIETS IN CHINA'S CITIES: EMPIRICAL FACT OR URBAN LEGEND?

Introduction

China's economic reforms, which began in 1978, resulted in remarkable growth in GDP, averaging 8% to 9% annually. Per capita nominal GDP increased from 379 yuan in 1978 to 10,561 yuan in 2004. In the meantime, 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, eggs, and aquatic products increased respectively from 20.52 kg, 5.22 kg, and 7.26 kg in 1981 to 29.22 kg, 10.35 kg, and 12.48 kg in 2004, (CNBS, various).

There is no doubt that household income and food prices strongly influence urban Chinese food consumption (Gould and Villarreal, 2006). However, economic prosperity, a growing openness to international markets, and domestic policy reforms have changed the food market environment for Chinese consumers. The removal of rationing, greater abundance and varieties of foods, changes in the marketing system, and changes in urban lifestyles may have contributed to shifts in consumer preferences. 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), but most support their conjectures with evidence based on trends in the data or estimates of expenditure elasticities, often from cross-sectional data sets. While trends and expenditure elasticities are informative, they do not provide convincing evidence of a shift in preferences. 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. Understanding whether or not observed shifts in food consumption in China are actual changes in preferences or simply the natural progression through a stable preference set requires empirical techniques that isolate preference changes.

The importance of preference change is highlighted in the growing number of articles that discuss globalization of food consumption in Asian countries (Mendez and Popkin, 2004; Pingali and Khwaja, 2004; Delgado, 2003; Lang, 1999). Lang defines dietary globalization as the general "transfer of diets, tastes, and health profiles from region to region." In particular, dietary patterns and foods from high-income countries tend to gain acceptance in developing countries as household incomes rise. Pingali and Khwaja (2004) note that, unlike the earlier phases of dietary evolution, in which income growth enables households to diversify their diets by purchasing a broader range of products within the local palette of traditional foods, households entering the globalization phase begin to "sever the link" between their purchases and local dietary habits. Clearly, dietary globalization represents a shift in preferences.

Given the size of China's population, the potential impacts on international agricultural markets of even minor shifts in consumer preferences can be significant. Ignorance of structural change can lead to faulty demand estimates that may provide misleading results from hypothesis tests, projections, and policy analysis (Moschini and Moro, 1996). Moreover, knowledge of the nature of parameter shifts aids in understanding the importance of the numerous policy and marketing changes in recent years in China and may improve projections of the path of future consumption changes. This paper extends our knowledge of Chinese food consumption by utilizing both parametric and nonparametric methods to investigate structural change. The parametric procedure described by Moschini and Meilke (1989) is used to estimate and test for shifts in parameter values over time. Changing parameter values can be viewed as evidence of structural change. To decrease the limitations of using a single functional form, both a dynamic AIDS (Almost Ideal Demand System) model and the Rotterdam model were used. In addition, to uncover evidence of structural change that is not dependent on the functional form chosen and to increase the robustness of our study, we employ nonparametric techniques for discovering preference changes.

The next section briefly describes the major policy and food market changes that have occurred in China over the last three decades. This information is extremely useful in interpreting the results of the empirical analysis. 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 the financial system have fueled the growth of China's economy and generated impressive economic development. Understanding the potential causes for structural change in urban Chinese food demand requires some knowledge of China's urban food rationing policy and economic reform.

Food rationing began in China in 1953 as a means of guaranteeing food security for urban residents. Rationed foods were obtained through a system of mandatory state procurement of agricultural products from farmers. The government was the sole seller of these rationed foods, and urban residents could only buy the rationed goods using rationing coupons. Before 1978, prices of most foods were administered by the Chinese government. 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. China's domestic trading entities were primarily organized as state-owned enterprises or cooperative enterprises, of which state-owned enterprises conducted the main activities of domestic trade. As economic restructuring progressed, the Chinese government began to lift its restrictions on trade in village fairs. A major shift in agricultural policy began in 1981, when the 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. In 1984, the state procurement of agricultural produce decreased dramatically. The number of foods and commodities subject to procurement dropped from 113 in 1981 to 60 in 1984. Price controls for 15 important non-staple foods, including pork, eggs, sugar, and vegetables, were lifted. However, the rationing system continued to dominate the free market for food in urban areas because rigid institutional constraints remained. Moreover, market prices were much higher than rationing prices during the period from 1978 to 1984. It was not until 1985 that the system of unified state procurement and sales of agricultural and sideline products was fully abolished for many non-staple foods. Within three years, rationing of the 15 non-staple foods in urban areas was totally eliminated. In addition, by 1987, a rapidly growing system of private and collectively owned food marketers coexisted with the state-owned system of commercial agencies and retail outlets. The emerging 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, which finally led to the abolition of the rationing system in 1993. At that time, staple foods such as grains and edible oils were no longer rationed.

In addition to changes in the food marketing system, other profound changes occurred in the economic and social aspects of life in urban China. Privatization of the housing market began in the late 1980s and was completed in 2001. Similarly, a growing share of medical and health care services was privatized in the late 1990s. Reforms of the education and employment markets gradually transformed the system of state provision from cradle to grave into a market-oriented approach that requires self-financed higher education and provides no guarantees of employment upon graduation. Finally, the increasing openness of China's domestic markets to foreign investment and imported goods has exposed China's consumers to a plethora of new food products offered in modern retail formats and has facilitated significant changes in consumer shopping behaviors (Veeck and Veeck, 2000; Hu et al., 2004). Each of these changes affects vital aspects of the daily lives of urban Chinese residents and has the potential to cause consumers to rethink their priorities and adjust their consumption preferences. The next section investigates the evidence for structural change in urban Chinese food demand using parametric methods.

Parametric Tests for Structural Change

Methodology

In demand analysis, structural change is often referred to as "changing tastes and preferences" (Moschini and Moro, 1996). These changes are reflected analytically as a change in the shape of individual utility functions. In this section, parametric methods are used to investigate structural changes in China's urban food demand by testing for parameter instability. As a precaution against the limitations of selecting a single functional representation of preferences, both the linear version of the AIDS model (LA/AIDS) and the Rotterdam model are estimated. Both models are flexible or equivalently flexible and are two of the most popular functional forms employed for demand analysis.

The LA/AIDS model developed by Deaton and Muellbauer (1980) is extensively used in demand estimation because of its consistency with the axioms of choice, aggregation properties, and flexibility in approximating arbitrary demand functions. Starting from a price independent generalized logarithmic (PIGLOG) cost function, the AIDS demand functions in budget share form are expressed as

$$w_{it} = \alpha_i + \sum_{j=1}^{n} \gamma_{ij} \log(p_{jt}) + \beta_i \log(\frac{y_t}{P}) \,, \tag{1}$$

where w_i is the budget share of good i; p_j is the price of good j; y is the total expenditure; and P is the translog price index in equation (2):

$$\log(P) = \alpha_0 + \sum_{k=1}^{n} \alpha_k \log(p_k) + \frac{1}{2} \sum_{j=1}^{n} \sum_{k=1}^{n} \gamma_{kj} \log(p_k) \log(p_j) \tag{2}$$

Since the stone price index suggested by Deaton and Muellbauer is not invariant to changes in the units of measurement of prices, we linearize the model by replacing the translog price index above with the Tornqvist price index.

$$\log(P_T) = \frac{1}{2} \sum_{i=1}^{n} (w_{it} + w_{i,t-1}) \log(\frac{p_{it}}{p_{i,t-1}}) \,.$$

The Tornqvist index is invariant to changes in units of measurement, and Diewert (1976) demonstrates that it is an exact approximation of the translog price index. We impose adding up, homogeneity, and symmetry properties of demand on the model, which imply the parameter restrictions in equation (3):

$$\sum_{i=1}^{n}\alpha_i = 1, \sum_{i=1}^{n}\beta_i = 0, \sum_{i=1}^{n}\gamma_{ij} = 0, \sum_{j=1}^{n}\gamma_{ij} = 0, \text{ and } \gamma_{ij} = \gamma_{ji} \tag{3}$$

Following Moschini and Meilke (1989), 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 general linear AIDS model is re-parameterized in equation (4) to capture time-varying parameter shifts:

$$w_{it} = \alpha_i + \gamma_i h_t + \sum_{j=1}^{n}(\gamma_{ij} + a_{ij}h_t)\log(p_{jt}) + (\beta_i + b_i h_t)\log(\frac{y_t}{P_T}) \tag{4}$$

Additional parametric restrictions in the structural change model associated with homogeneity, adding up, and symmetry are $\sum_{i=1}^{n}\gamma_i = 0$, $\sum_{i=1}^{n}b_i = 0$, $a_{ij} = a_{ji}$, and $\sum_{i=1}^{n}a_{ij} = 0$. 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):

$$\begin{aligned} h_t &= 0, && \textit{for } t = 1,....,\tau_1; \\ h_t &= (t-\tau_1)/(\tau_2-\tau_1), && \textit{for } t = \tau_1+1,...,\tau_2-1; \\ h_t &= 1, && \textit{for } t = \tau_2,...,T. \end{aligned} \tag{5}$$

The value τ_1 is the end point 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.

Preliminary testing suggested that the dynamic version of the LA/AIDS model provides a better representation of the dynamics of Chinese food demand. The first-difference form of the estimated model is given in equation (6):

$$\Delta w_{it} = \gamma_i \Delta h_t + \sum_{j=1}^{n}(\gamma_{ij}\Delta\log(p_{jt}) + a_{ij}\Delta(h_t \log(p_{jt})) + \beta_i \Delta\log(\frac{y_t}{P}) + b_i \Delta(h_t \log(\frac{y_t}{P})) \tag{6}$$

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

The structural change version of the Rotterdam model is similar to the LA/AIDS model. By introducing the time path variable, the structural change version of the Rotterdam model is specified in equation (7):

$$w_{it} d\log(q_{it}) = (c_i + k_i h_t)\sum_{j=1}^{n} w_{jt} d\log(q_{jt}) + \sum_{j=1}^{n}(c_{ij} + k_{ij} h_t) d\log(p_{jt}) \tag{7}$$

Homogeneity, adding-up, and symmetry require the following parameter restrictions:

$$\sum\nolimits_j c_j = 1, \sum\nolimits_j c_{ij} = \sum\nolimits_i c_{ij} = 0, c_{ij} = c_{ji},$$
$$\sum\nolimits_j k_j = 0, \sum\nolimits_j k_{ij} = \sum\nolimits_i k_{ij} = 0, k_{ij} = k_{ji}.$$

Unlike the LA/AIDS model, which approximates the demand function in the variable space, the Rotterdam model approximates the demand function in the parameter space. Although it cannot be considered as an exact representation of preferences without imposing strong constraints on the model, the Rotterdam model is still very useful as a flexible function form for approximating a demand system.

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 10 food groups: grain, pork, beef/mutton, poultry, eggs, fish, vegetables, fruit, milk, and other foods. Pork is the most commonly consumed meat product, and pork prices are low relative to other meats. Beef currently represents only a small proportion of total meat products consumed, but its share of meat expenditures has increased substantially over the study period. Per capita beef consumption is highest in China's western pastoral provinces. Given the limited number of annual observations and the relatively large number of parameters in the structural change model, attempts to estimate the models for all 10 commodities individually did not converge consistently. Consequently, it was necessary to reduce the number of commodities estimated by aggregating some food groups together for the parametric tests.

Two different aggregations were estimated. First, the three meats were combined into a meat group, and milk and other food expenditures were combined. The resulting seven food groups were grain, meat, eggs, fish, vegetables, fruit, and other foods. The consumer price index was used as the price of other foods. The Tornqvist price index was used to aggregate pork, beef, and poultry prices into a meat price index and to aggregate the milk price with the price for 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. In addition, in order to compare the grouping effects on structural change results, a second set of estimates were produced by further aggregating meat and eggs into a single group and by combining fruits and vegetables. The resulting five food groups include grain, meat and eggs, fish, vegetables and fruit, and other foods. The Tornqvist price index was used to aggregate pork, beef, poultry, and egg prices into the meat and egg group price index and to aggregate vegetables and fruit prices into a single index. Aggregate quantities are recovered by dividing group expenditures by the appropriate price indices. All prices and income were normalized by their sample mean.

To be comparable with the parametric analysis, the nonparametric methods were also applied to the 7- and 5-commodity group aggregations, as well as to the original 10 commodity groups. The discussion in the results section for both parametric and nonparametric methods focuses on the analysis with 7 commodity groups, but important differences in the outcomes using 5 commodity groups and additional findings from the 10 commodity groups are also noted.

Results from Structural Change Tests

The dynamic linear AIDS model in equation (6) and the Rotterdam model in equation (7) were estimated using the maximum likelihood estimation procedure in TSP 4.5. In estimating both models, the equation for other foods was omitted to avoid singularity problems. Homogeneity, symmetry, and adding-up restrictions were imposed on the model parameters. There are 210 possible combinations of τ_1 and τ_2. With the limitation on degree of freedom, not all sets of combinations can be estimated. For the dynamic linear AIDS model, a system of equations are estimable for the periods $1981 \leq \tau_1 \leq 1996$ and $1989 \leq \tau_2 \leq 2004$ for the model with seven groups and $1981 \leq \tau_1 \leq 1998$ and $1987 \leq \tau_2 \leq 2004$ for the model with five groups. The corresponding ranges for the Rotterdam model are $1981 \leq \tau_1 \leq 1997$ and $1988 \leq \tau_2 \leq 2004$ for the model with seven groups and $1981 \leq \tau_1 \leq 1999$ and $1987 \leq \tau_2 \leq 2004$ for the model with five groups.

The AIDS and Rotterdam models were 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 points (τ_1, τ_2) resulting in the maximum values for the likelihood function are shown in Table 1. Based on likelihood ratio tests, several other combinations of τ_1 and τ_2 failed to reject the null hypothesis of structural change. These additional structural change points are also displayed in Table 1.

Table 1.Maximum Likelihood Structural Change Points

	Dynamic AIDS Model	Rotterdam Model
Seven Commodity Groups		
Optimal Points	(1982, 1990)	(83,89)
Additional Points		[(81-85),89],[(81-85),91]
Five Commodity Groups		
Optimal Points	(1985, 1993)	(1993, 1994)
Additional Points	[(85,86,87),88] [(84, 85, 86), (92,93)], (91, 93)	[(85,86,87),(94,99,00)], (95,96), [85,(89,90,98)],[(85,86),(87,88)], [(81,82,83),02], [(84-88),(01-04)], [(84,88,89,92), 94]

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.

With seven commodity groups, the optimal structure change points are τ_1=1982, τ_2=1990 for the dynamic AIDS model and τ_1=1983, τ_2=1989 for the Rotterdam model. Both models identify a gradual shift in preferences that corresponds to the period of time when the dual-track marketing system was established and rationing of non-staple foods was eliminated. The additional structural change points identified by the Rotterdam model confirm the second half of the 1980s as a period of preference change in urban China. Interestingly, the AIDS model with five commodity groups also identifies the late 1980s as a period of structural change for food demand. The optimal structural change points are τ_1=1985, τ_2=1993. This range nearly perfectly overlaps the dual-track marketing period, which ended with the complete abolition of rationing. The Rotterdam model with five commodities finds abrupt structural change in the year when rationing ended (τ_1=1993, τ_2=1994). However, it also identifies (τ_1=1984, τ_2=1994) as possible structural change points, which matches the results from the other models. Both models with five commodity groups identify a number of additional structural change points that include ranges in the

1980s and early 1990s. The Rotterdam model with five commodities also provides evidence of abrupt structural change in 1995–1996 and more gradual structural change throughout the 1990s. The late 1990s was a period marked by the growing importance of supermarkets as a retail format in urban areas and the influx of foreign food products (Hu et al., 2004). These results may provide some support for the hypothesis that globalization of diets is occurring in China.

The results from the structural change test demonstrate that policy changes and market transformations have prompted significant shifts in urban Chinese food demand, especially during the 1980s when food rationing was abolished and free markets were developed. This result is robust across models and commodity aggregations. Furthermore, the more aggregated models with five food groups capture more policy change points, including the elimination of grain rationing, while the more disaggregated model with seven food groups only captures the gradual structural change during policy reform periods in the 1980s. One explanation for this result may be that when commodities experiencing similar market transformations are aggregated into a single group, such as meats and eggs, the effects of policy changes in the individual markets reinforce one another, allowing the model to detect additional change points. On the contrary, disaggregating the data disperses the effects of rationing on food demand across the individual commodity groups, and the elimination of rationing for non-staple foods, which occurred first, may appear to be the larger effect. And after the establishment of free markets for non-staples, the effect of the elimination of grain rationing appears less prominent.

The maximum likelihood parameter estimates for the dynamic AIDS model and the Rotterdam model with seven commodity groups are listed in Tables 2 and 3, respectively. 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. For the dynamic AIDS model, all time variable parameters in the grain equation are insignificant, indicating that there was no significant structural change for grain demand during the optimal structural change period, (1982–1990). One reason for the insignificance is that the effect of eliminating grain rationing is less prominent because of the earlier elimination of rationing for non-staple foods. As most of the time path parameters for intercepts are insignificant,

Table 2. Maximum Likelihood Parameter Estimates for Dynamic AIDS Model with Optimal Structural Change Points for Seven-Commodity Group at (1982, 1990)

	Grain	Meat	Fish	Veg.	Fruit	Eggs	Other
	0.0783	0.0148	-0.1141	0.0126	-0.0780	0.0095	0.0767
Intercept	(0.1127)	(0.1282)	(0.0768)	(0.0700)	(0.0396)	(0.0359)	(0.2297)
Grain							
γ_{ij}	0.0493						
	(0.0802)						
α_{ij}	0.0460						
	(0.0834)						
Meat							
γ_{ij}	0.0491	-0.6530					
	(0.0872)	(0.1871)					
α_{ij}	-0.0824	0.6947					
	(0.0904)	(0.1927)					
Fish							
γ_{ij}	-0.0300	0.4950	-0.3227				
	(0.0506)	(0.0866)	(0.0505)				
α_{ij}	-0.0125	-0.4993	0.3362				
	(0.0523)	(0.0888)	(0.0535)				
Vegetables							
γ_{ij}	0.0020	-0.4692	0.3906	-0.2258			
	(0.0451)	(0.0531)	(0.0319)	(0.0337)			
α_{ij}	-0.0211	0.4423	-0.3546	0.2470			
	(0.0472)	(0.0553)	(0.0338)	(0.0357)			
Fruit							
γ_{ij}	-0.0154	-0.1796	0.1497	-0.0962	-0.0209		
	(0.0215)	(0.0305)	(0.0178)	(0.0155)	(0.0115)		
α_{ij}	-0.0015	0.1897	-0.1739	0.1149	0.0371		
	(0.0234)	(0.0326)	(0.0208)	(0.0176)	(0.0153)		
Eggs							
γ_{ij}	-0.0545	0.2892	-0.0897	0.0440	0.0458	-0.0105	
	(0.0258)	(0.0371)	(0.0198)	(0.0205)	(0.0083)	(0.022)	
α_{ij}	0.0423	-0.3005	0.0999	-0.0436	-0.0404	0.0246	
	(0.0268)	(0.0382)	(0.0210)	(0.0216)	(0.0098)	(0.0223)	
Others	-0.0004	0.4684	-0.5929	0.3546	0.1166	-0.2244	-0.1219
γ_{ij}	(0.1431)	(0.2006)	(0.1064)	(0.0821)	(0.0445)	(0.0462)	(0.3108)
	0.0292	-0.4444	0.6401	-0.3850	-0.1259	0.2177	0.1042
α_{ij}	(0.1526)	(0.2116)	(0.1131)	(0.0882)	(0.0512)	(0.0544)	(0.3407)
Expenditure							
β_i	-0.0713	-0.0631	-0.0270	-0.1458	0.0131	0.0506	0.2435
	(0.0522)	(0.0844)	(0.0451)	(0.0368)	(0.0209)	(0.0271)	(0.1052)
b_i	0.0424	0.1055	0.067	0.0934	-0.0098	-0.0583	-0.2402
	(0.0619)	(0.0895)	(0.0543)	(0.0435)	(0.0285)	(0.0315)	(0.1319)
$\mathbf{R^2}$	0.7021	0.6743	0.3652	0.6102	0.7682	0.8016	
DW	1.1112	1.5474	1.6046	1.2268	1.5339	1.2073	

Note: 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.

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

	Grain	Meat	Fish	Veg.	Fruit	Eggs	Other
Volume Index							
c_i	-0.0270	0.0308	-0.0918	-0.0832	0.0290	0.2324	0.9097
	(0.1025)	(0.1321)	(0.0728)	(0.0484)	(0.0387)	(0.0243)	(0.1660)
k_i	0.0341	0.2385	0.2132	0.0679	0.0733	-0.2280	-0.3991
	(0.1070)	(0.1398)	(0.0825)	(0.0517)	(0.0471)	(0.0256)	(0.1845)
Grain							
c_{ij}	-0.1949						
	(0.1326)						
k_{ij}	0.1950						
	(0.1331)						
Meat							
c_{ij}	0.0779	-0.7154					
	(0.0981)	(0.1647)					
k_{ij}	-0.0901	0.5989					
	(0.1000)	(0.1707)					
Fish							
c_{ij}	0.0849	0.3724	-0.3072				
	(0.0635)	(0.0712)	(0.0434)				
kij	-0.1071	-0.3491	0.2679				
	(0.0645)	(0.0742)	(0.0489)				
Vegetables							
c_{ij}	0.0124	-0.2648	0.2693	-0.2517			
	(0.0537)	(0.0426)	(0.0272)	(0.0270)			
k_{ij}	-0.0229	0.2536	-0.2473	0.2169			
	(0.0540)	(0.0437)	(0.0285)	(0.0273)			
Fruit							
c_{ij}	-0.0049	-0.0889	0.1355	-0.0529	-0.0451		
	(0.0249)	(0.0308)	(0.0173)	(0.0132)	(0.0127)		
k_{ij}	0.0009	0.1134	-0.1639	0.0767	-0.0168		
	(0.0261)	(0.0331)	(0.0214)	(0.0148)	(0.0169)		
Eggs							
c_{ij}	-0.2223	0.2960	-0.0559	-0.0006	0.0276	-0.0865	
	(0.0313)	(0.0229)	(0.0144)	(0.0151)	(0.0062)	(0.0203)	
k_{ij}	0.2268	-0.3000	0.0632	0.0090	-0.0240	0.0612	
	(0.0313)	(0.0235)	(0.0148)	(0.0153)	(0.0071)	(0.0206)	
Others							
c_{ij}	0.2468	0.3227	-0.4991	0.2882	0.0286	0.0417	-0.4290
	(0.1604)	(0.1762)	(0.0912)	(0.0709)	(0.0438)	(0.0422)	(0.2880)
k_{ij}	-0.2027	-0.2267	0.5361	-0.2861	0.0138	-0.0362	0.2017
	(0.1632)	(0.1848)	(0.0961)	(0.0724)	(0.0486)	(0.0432)	(0.3017)
$\mathbf{R^2}$	0.1805	0.8969	0.0889	0.3908	0.6670	0.9290	
DW	1.5151	1.1248	1.8393	1.7876	1.8576	1.7610	

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.

we can conclude that there is no significant trend change in urban Chinese food demand. The exception is fruit, which shows a significant negative change in the trend. On the other hand, price and income changes have positive effects on fruit consumption, which offset, to some degree, the negative effects from the trend change. While most time path parameters for prices are significant, most time path parameters for expenditures are insignificant. This indicates that urban consumers' responses to price movements changed significantly after the structural change period while the response to income growth did not. Similarly, in the Rotterdam model, all time path parameters in the grain equation, except the coefficient for the egg price, are insignificant. And most time path parameters for price are significant.

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 for both the AIDS and Rotterdam models. 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 models. 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, except for the test on expenditure parameters in the AIDS model with five commodity groups. Thus, despite the insignificance of individual parameters, joint tests suggest that shifts in preferences of urban Chinese households involve consumers' response to price and income changes, as well as a general shift in consumption trends.

The average Marshallian price and income elasticities calculated at the mean shares for both the AIDS and Rotterdam models with seven commodity groups are reported in Tables 5 and 6, respectively. Standard errors for the elasticities are computed using the delta method (Green et al., 1987). For the dynamic AIDS model, all own-price elasticities were negative except that of grains before the optimal structural change. With rationing, meager supplies, and scant varieties of foods, Chinese urban food demand was skewed in favor of grains. After the structural change, all food demands became less elastic. Moreover, meat, fish, vegetables, fruits, and eggs changed from price elastic to price inelastic. Grains and

Table 4. Likelihood Ratios for Structural Change Tests for Dynamic AIDS and Rotterdam Models

Hypothesis	Restrictions	Likelihood Ratio	$\chi^2_{0.05}$
AIDS Model			
No Structural Change in:			
All parameters			
-Seven groups	33	162.8220	43.77
-Five groups	18	73.3431	28.8693
Intercept parameters			
-Seven groups	6	30.6360	12.5916
-Five groups	4	18.5251	9.4877
Price parameters			
-Seven groups	21	126.1280	32.6706
-Five groups	10	40.0631	18.3070
Expenditure parameters			
-Seven groups	6	15.1240	12.5916
-Five groups	4	5.9031*	9.4877
Price and expenditure parameters			
-Seven groups	27	143.1060	40.1133
-Five groups	14	62.1931	23.6848
Rotterdam Model			
No Structural Change in:			
All parameters			
-Seven groups	27	137.8345	40.1133
-Five groups	14	37.7529	23.6848
Price parameters			
-Seven groups	21	120.9985	32.6706
-Five groups	10	37.0489	18.3070
Volume Index parameters			
-Seven groups	6	43.1565	12.5916
-Five groups	4	11.2389	9.4877

* indicates cannot be rejected at 0.05 significance level.

vegetables changed from inferior goods to necessities, but their income elasticities are not significant before the structural change. Meat and fish changed from necessities to luxuries. Conversely, eggs changed from a luxury to a necessity. In addition, eggs and fish have a substitution relationship with vegetables, while meats are complementary goods with vegetables.

Table 5. Average Marshallian Price and Income Elasticities for the AIDS Model with Seven-Commodity Group*

	Grain	Meat	Fish	Vegetables	Fruit	Eggs	Other Foods
				Before Structural Change			
Grain	0.0218	1.1831	-0.3355	0.1587	-0.0912	-1.0077	0.4456
	(1.5616)	(1.5968)	(1.0472)	(0.8865)	(0.4710)	(0.5061)	(2.7661)
Meat	0.3055	-4.7452	2.9525	-2.7044	-0.9922	1.6987	2.8532
	(0.5100)	(1.0260)	(0.5685)	(0.3067)	(0.1833)	(0.2234)	(1.2458)
Fish	-0.1613	2.8145	-2.7909	2.2134	0.8661	-0.5004	-3.2892
	(0.2845)	(0.4602)	(0.3140)	(0.1765)	(0.1100)	(0.1140)	(0.6282)
Vegetables	0.1096	-5.0894	4.7717	-3.4418	-0.8525	0.5581	4.6149
	(0.5196)	(0.5777)	(0.4011)	(0.3927)	(0.2057)	(0.2335)	(0.9518)
Fruit	-0.1074	-1.2149	0.9848	-0.6507	-1.1527	0.3033	0.7499
	(0.1440)	(0.1952)	(0.1325)	(0.1021)	(0.0849)	(0.0568)	(0.3023)
Egg	-1.7855	8.7670	-3.0839	1.2384	1.1948	-1.3800	-7.5330
	(0.8212)	(1.0932)	(0.7061)	(0.6802)	(0.3217)	(0.6775)	(1.4499)
Other Foods	-0.0395	1.2919	-1.9261	1.0096	0.2426	-0.7029	-1.2972
	(0.4344)	(0.5809)	(0.3519)	(0.2486)	(0.1478)	(0.1428)	(0.2792)
Expenditure	-0.3749	0.6319	0.8478	-0.6707	1.0878	2.5821	1.7372
	(1.0071)	(0.4921)	(0.2538)	(0.4218)	(0.1397)	(0.8474)	(0.3187)
				After Structural Change			
Grain	-0.2389	-0.2148	-0.3125	-0.1246	-0.1137	-0.0850	0.3115
	(0.1354)	(0.1181)	(0.0896)	(0.0931)	(0.0578)	(0.0438)	(0.1744)
Meat	-0.2086	-0.8185	-0.0372	-0.1672	0.0378	-0.0699	0.0360
	(0.0864)	(0.1255)	(0.0785)	(0.0634)	(0.0488)	(0.0355)	(0.1507)
Fish	-0.7516	-0.1842	-0.8277	0.5047	-0.4264	0.1340	-0.0792
	(0.2048)	(0.2261)	(0.3077)	(0.1831)	(0.1817)	(0.1199)	(0.3225)
Vegetables	-0.1229	-0.1715	0.3939	-0.7356	0.2240	0.0267	-0.0901
	(0.1220)	(0.1156)	(0.1186)	(0.1083)	(0.0845)	(0.0532)	(0.1649)
Fruit	-0.2413	0.1319	-0.3405	0.2551	-0.7775	0.0741	-0.1487
	(0.1164)	(0.1233)	(0.1576)	(0.1139)	(0.1402)	(0.0708)	(0.1963)
Egg	-0.2711	-0.2365	0.2566	0.0303	0.1446	-0.6547	-0.0839
	(0.1459)	(0.1495)	(0.1786)	(0.1203)	(0.1233)	(0.0982)	(0.2341)
Other Foods	0.0697	0.0574	0.0272	-0.0753	-0.0234	-0.0166	-1.0468
	(0.0881)	(0.0957)	(0.0682)	(0.0573)	(0.0484)	(0.0355)	(0.1461)
Expenditure	0.7779	1.2276	1.6305	0.4754	1.0470	0.8146	1.0080
	(0.1897)	(0.1367)	(0.3361)	(0.1673)	(0.1929)	(0.2528)	(0.1409)

*structural change point is (1982, 1990). Asymptotic standard errors are reported in parentheses.

Table 6. Average Marshallian Price and Income Elasticities for Rotterdam Model with Seven-Commodity Group*

	Grain	Meat	Fish	Vegetable	Fruit	Egg	Other Foods
				Before Structural Change			
Grain	-0.8892	0.3896	0.4050	0.0718	-0.0162	-1.0389	1.2049
	(0.6928)	(0.4153)	(0.3052)	(0.2606)	(0.1265)	(0.1519)	(0.7130)
Meat	0.3924	-3.9628	2.0392	-1.4729	-0.4980	1.6182	1.7145
	(0.5712)	(0.8127)	(0.4129)	(0.2418)	(0.1753)	(0.1305)	(1.0938)
Fish	2.3813	8.8688	-6.9097	6.3595	3.2037	-1.1651	-10.6454
	(1.5809)	(1.4669)	(1.0273)	(0.6289)	(0.4222)	(0.3416)	(2.1643)
Vegetable	0.2865	-2.3747	2.5970	-2.3112	-0.4597	0.0352	3.0182
	(0.5707)	(0.3764)	(0.2622)	(0.2741)	(0.1367)	(0.1427)	(0.6359)
Fruit	-0.2031	-1.7237	2.4566	-1.0233	-0.8533	0.4778	0.3378
	(0.4812)	(0.5322)	(0.3280)	(0.2501)	(0.2496)	(0.1205)	(0.8395)
Egg	-5.2097	4.8639	-1.2671	-0.4807	0.2860	-1.8907	-0.7574
	(0.6539)	(0.4304)	(0.2750)	(0.3207)	(0.1347)	(0.3836)	(0.7607)
Other Foods	0.1526	0.4497	-1.5418	0.5508	-0.0604	-0.0163	-1.5597
	(0.4892)	(0.4703)	(0.2652)	(0.2223)	(0.1365)	(0.1225)	(0.2341)
Expenditure	-0.1269	0.1693	-2.0932	-0.7914	0.5312	4.4557	2.6022
	(0.4819)	(0.7259)	(1.6605)	(0.4601)	(0.7074)	(0.4658)	(0.4748)
				After Structural Change			
Grain	-0.0062	-0.1163	-0.1961	-0.0967	-0.0391	0.0364	0.3558
	(0.1150)	(0.1114)	(0.1033)	(0.0731)	(0.0733)	(0.0238)	(0.1349)
Meat	-0.2439	-0.9275	0.0285	-0.2091	0.0328	-0.0794	-0.1231
	(0.0804)	(0.1427)	(0.0898)	(0.0514)	(0.0594)	(0.0205)	(0.1650)
Fish	-0.5320	0.0271	-0.6993	0.1525	-0.5418	0.0415	-0.2368
	(0.1865)	(0.2154)	(0.2859)	(0.1305)	(0.1737)	(0.0548)	(0.2607)
Vegetable	-0.0904	-0.0875	0.2398	-0.3461	0.2584	0.0931	0.0914
	(0.0848)	(0.0797)	(0.0900)	(0.0673)	(0.0722)	(0.0250)	(0.1012)
Fruit	-0.2283	0.0916	-0.5101	0.2015	-0.9964	-0.0030	-0.0334
	(0.1309)	(0.1397)	(0.1674)	(0.1014)	(0.1685)	(0.0480)	(0.1804)
Egg	0.1057	-0.1286	0.1888	0.2133	0.0884	-0.6839	0.0963
	(0.0743)	(0.0809)	(0.0935)	(0.0653)	(0.0915)	(0.0473)	(0.1025)
Other Foods	-0.0336	0.0128	0.0054	-0.1074	0.0162	-0.0309	-1.0303
	(0.0506)	(0.0775)	(0.0516)	(0.0296)	(0.0389)	(0.0133)	(0.0970)
Expenditure	0.0622	1.5218	1.7889	-0.1587	1.4779	0.1199	1.1678
	(0.2087)	(0.1971)	(0.4116)	(0.1542)	(0.2938)	(0.1790)	(0.1253)

*structural change point is (1983, 1989). Asymptotic standard errors are reported in parentheses.

All own-price elasticities from the Rotterdam model were negative. And similar to the AIDS model, food demands became generally less elastic after the structural change. Meat, fish, vegetables, and eggs changed from price elastic to price inelastic, as with the AIDS model, but fruit became slightly more price elastic. As with the AIDS model, grain changed from an inferior good to a necessity, but its income elasticity is not significant. Besides differences in magnitude, the elasticity estimates from these two models showed differences in signs, especially in the income elasticity for vegetables.

Generally, the income 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 both the AIDS and 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 Tables 1A and 2A in the Appendix), which underscores the importance of testing and adjusting for structural change in empirical and applied analysis.

Nonparametric Tests of Structural Change

Because a functional form is assumed in the parametric approach for testing preference changes, 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 use of both the 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 several nonparametric tests for stable preferences. The nonparametric approach has the advantage that no assumptions regarding the functional representation of preferences are needed.

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. Consequently, a simple test for stable preferences checks the data for compliance with GARP.

GARP states that if a consumption bundle, x_j, is revealed preferred to another bundle, x, then x cannot be strictly directly revealed preferred to x_j. The bundle x_j is revealed preferred to bundle x (written $x_j R x$) when the relationship in equation (8) holds for the sequence of bundles $\{x_j, x_k, x_l, \ldots, x\}$.

$$p_j x_j \geq p_j x_k, p_k x_k \geq p_k x_l, \ldots, p_m x_m \geq p_m x \tag{8}$$

In equation (8), the bundle x_j is directly revealed preferred to x_k (written $x_j R^0 x$) because the cost of purchasing x_k at prices p_j is less than or equal to the cost of purchasing x_j. In other words, if a consumer purchases x_j when x_k is affordable, then the consumer must prefer x_j. Revealed preference establishes a transitive closure for a sequence of bundles that are connected through the directly revealed preferred relationship. GARP stipulates that if $x_j R x$, then x_j cannot cost less than x evaluated at the price vector associated with bundle x; otherwise, the data is not consistent with utility-maximizing behavior (Varian, 1982). Finding one observation that violates GARP is technically sufficient to reject consistency of the data with utility maximization.

We apply the algorithm described by Varian (1982, 1983) for testing GARP to the data used for the parametric tests described in the previous section. Because degrees of freedom are not an issue, we are able to apply the test to the disaggregated data as well as to the data aggregated into five and seven commodity groups. All three data sets satisfy 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 nonparametric tests, 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 consumption bundle 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. They employ the Slutsky equation to separate the change in consumption from one time period to the next into a pure price component (a movement along an indifference curve) and an income component (corresponding to a shift to a new indifference curve as the budget set changes). This relationship is summarized in equation (9), where f_i is the Marshallian demand function for good i, q_i is the consumption of good i, y is income, ε_{iy} is the income elasticity of good i, and ctc_i is the change in tastes. The variable a is the change in expenditures that can be attributed to the change in the budget set and is defined as $a = \Delta y - \sum_{j=1}^{n} q_j \Delta p_j$.

$$\Delta q_i = \sum_{j=1}^{n} \frac{\partial f_i}{\partial p_j}\bigg|_{U_0} \Delta p_j + a\frac{q_i}{y}\varepsilon_{iy} + ctc_i \tag{9}$$

Given some reasonable value or 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 terms can be attributed to a change in tastes. The difficulty in making an empirical assessment of equation (9) is that the change in consumption due to price changes (the movement along the indifference curve) cannot be observed. Consequently, equation (9) is rearranged using the fact that $q_{i,t-1} + \sum_{j=1}^{n} \frac{\partial f_i}{\partial p_j}\Big|_{U_0} \Delta p_j = q_{i,t}^*$, where q_i^* is the compensated demand. Thus, equation (10) is a restatement of the Slutsky relationship; however, the change in consumption due to price changes has been subsumed into the Hicksian demand quantity, isolating the income and taste change (tc) components.

$$q_{i,t}^* = q_{i,t} + \sum_{j=2}^{t} a_j \frac{q_{ij}}{y_j}\varepsilon_{iy}^j + \sum_{j=2}^{t} tc_{ij} \tag{10}$$

Equation (10) defines the quantity consumed at a particular time *t*, which is computed by summing the income, price, and taste change effects from some arbitrary base period until time *t*. Consequently, both the income and the taste change components of equation (10) represent the cumulative changes in their respective quantities up to time *t*. Sakong and Hayes (1993) use the convexity of preferences to define a relationship between two compensated bundles, which is similar to the revealed preference relation described above. In particular, expenditures on the bundle of compensated demands at time *t* prices must exceed the expenditures on the optimal bundle of compensated demands at time *s* and time *t* prices, for all combinations of *t* and *s*.

$$p_t x_t^* > p_t x_s^* \quad \forall t, s \in T \tag{11}$$

If the relationship in equation (11) does not hold, then there has been a violation of the null hypothesis of stable preferences.

Based on the relationships in equations (10) and (11), we can define a linear programming problem that solves for the minimum change in tastes that are consistent with the observed prices and consumption quantities and an assumed range of income elasticity values. Income elasticities are endogenous in this model, but they are constrained by the Engel aggregation condition and the assumed upper and lower bounds. We applied the Sakong and Hayes model to the disaggregated consumption data and the seven and five 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. The model allows the analyst to select a range over which the expenditure elasticity for each commodity may vary. 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 7 displays the expenditure elasticity ranges selected for this study.

In addition to the bounds placed on income elasticities, we placed bounds on the year-to-year change in income elasticity values (Sakong and Hayes, 1993). Using the results from the AIDS and Rotterdam models as a guide, we computed the average change in the income elasticity values on an annual basis. For most commodities, income 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 income elasticity changes: limited to 0.15, limited to 0.2, and no limit. Regardless of the year-to-year change restriction, elasticities were bounded by the values in Table 7.

Table 7. Expenditure elasticity ranges by commodity

Commodity	Maximum	Minimum
Grain	1.3	0.0
Beef	1.4	0.8
Pork	1.3	0.7
Poultry	1.3	0.5
Eggs	1.0	0.4
Fish	1.5	0.8
Meat	1.3	0.7
Dairy	2.2	0.9
Fruits	1.5	0.6
Vegetables	1.2	0.6
Other	1.4	0.7
Meat & Eggs	1.2	0.7
Fruit & Vegetables	1.3	0.5
Milk & Other	1.6	0.9

Table 8 displays the structural change points identified when income elasticity changes were restricted to 0.15 per year. Allowing income 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 early 1980s and in the mid- to late 1990s; 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. Figure 1 shows the cumulative taste changes measured in kilograms per person for the seven-group analysis. Since the consumption of other foods cannot be

Table 8. Structural change points identified by the nonparametric approach

Commodity	10 Groups	7 Groups	5 Groups
Grain	–	2001	1996–1997, 1999, 2002–2003
Beef	1995, 2002	–	–
Pork	–	–	–
Poultry	1996, 2000	–	–
Eggs	1983, 1985–1986, 1990–1991, 1996 1998, 2000	1998–1999	–
Fish	1997–1999	1985, 1996–1997, 2001	1985, 1997–1999
Dairy	–	–	–
Fruits	1982–1983, 1985–1986, 1994	1982–1983, 1999, 2002–2003	–
Vegetables	–	–	–
Other	–	–	–
Meat	–	–	–
Meat & Eggs	–	–	1981–1985, 1998
Fruit & Vegetables	–	–	1986
Milk & Other	–	1982, 1985–1986, 1998–1999	1985, 1998

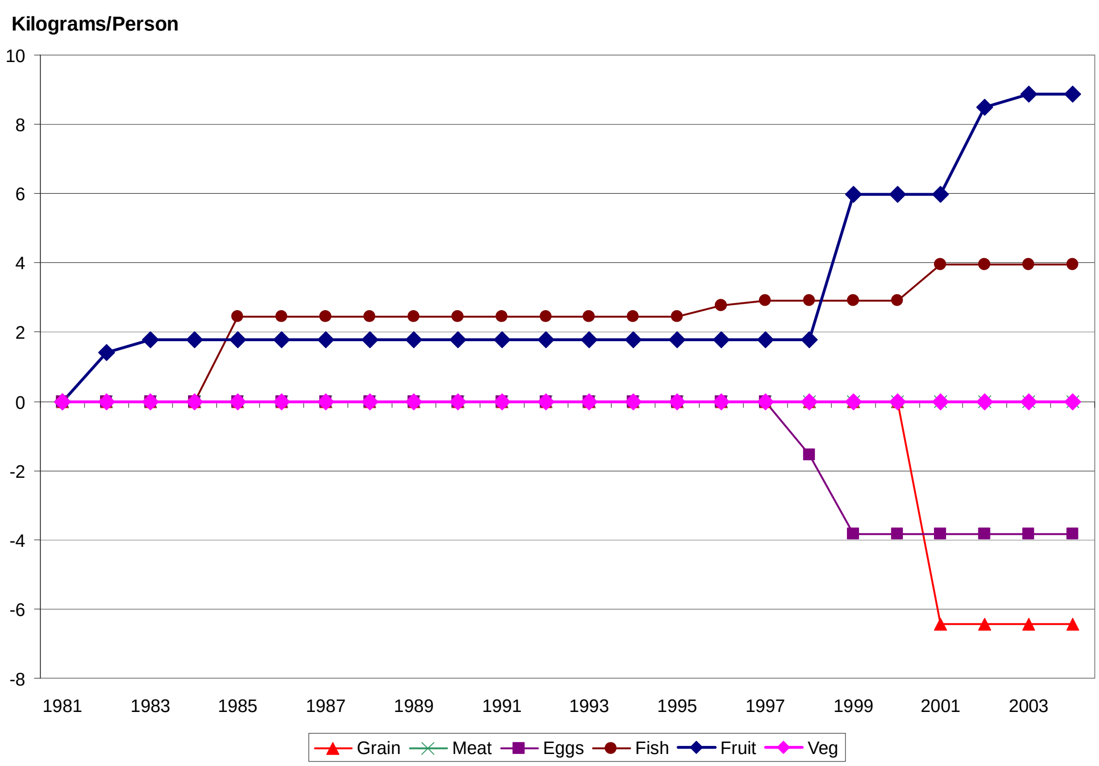

Figure 1. Cumulative taste changes in quantities: seven-commodity analysis

quantified, only six commodities are shown in Figure 1. Shifts in preferences were targeted toward increased consumption of aquatic products and fruits in the early 1980s. Consumption of these commodities also increased with preference shifts in the 1990s, but grain and egg consumption declined with preference changes in the late 1990s. In the case of grains, per capita consumption levels actually declined by 19 kg from 1995 to 2004, roughly 20% of the 1995 consumption level. In contrast, egg consumption after 1995 was higher than the 1995 level for eight of nine of the sample years. Thus, the negative taste change shown in Figure 1 suggests that egg consumption could have increased substantially more without the shift, roughly 36% above the 1995 level.

Milk and other food consumption also showed evidence of structural change. As a result of aggregation, it is not possible to display changes in quantities for milk and other foods, but Figure 2 displays the changes in real expenditures implied by the taste changes computed in the model. The growth in fruit and aquaculture expenditures in the early 1980s appears to have occurred at the expense of expenditures on foods in the milk and other category. Inflation eroded the growth in real expenditures on fruit and fish through

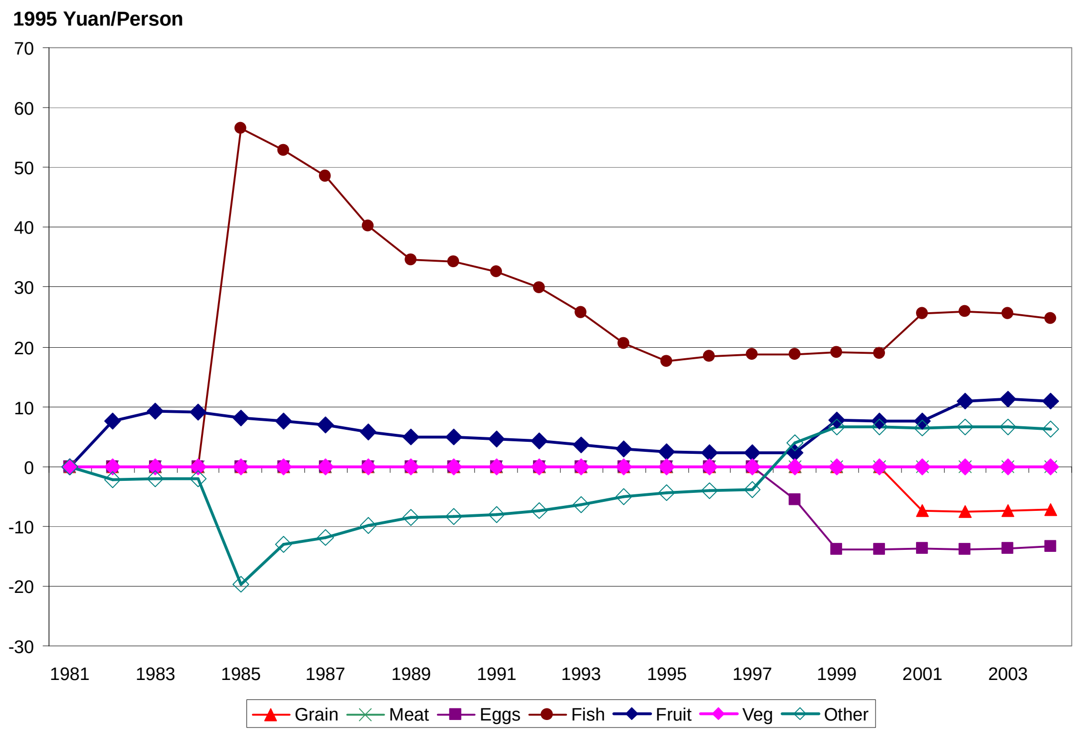

Figure 2. Cumulative taste changes in real expenditures: seven-commodity analysis

out the late 1980s and early 1990s, but taste changes in the late 1990s prompted real growth in aquaculture and fruit. Following the initial downward shift, real expenditures for milk and other food increased, particularly during the 1997–1999 period.

Given the income elasticity bounds in Table 7, changes in both meat and vegetable consumption during the study period can be entirely explained by income and price effects in the seven-group analysis. By contrast, the analysis of the disaggregated data shows some evidence of structural change for beef and poultry. In the seven-commodity analysis, the positive change in consumer preferences for beef and poultry were swamped by the large contribution of pork to the aggregated meat group. Figure 3 displays the estimated taste changes in quantity terms for 9 of the 10 commodity groups. In addition to the changes in individual meat types, the 10-group analysis finds changes in egg consumption throughout the study period, but no changes in preferences for grain, pork, vegetables, milk, and other food consumption. The taste changes in fruit and aquaculture products are similar to the seven-commodity results, but the early change in aquaculture consumption is not present in the disaggregated analysis.

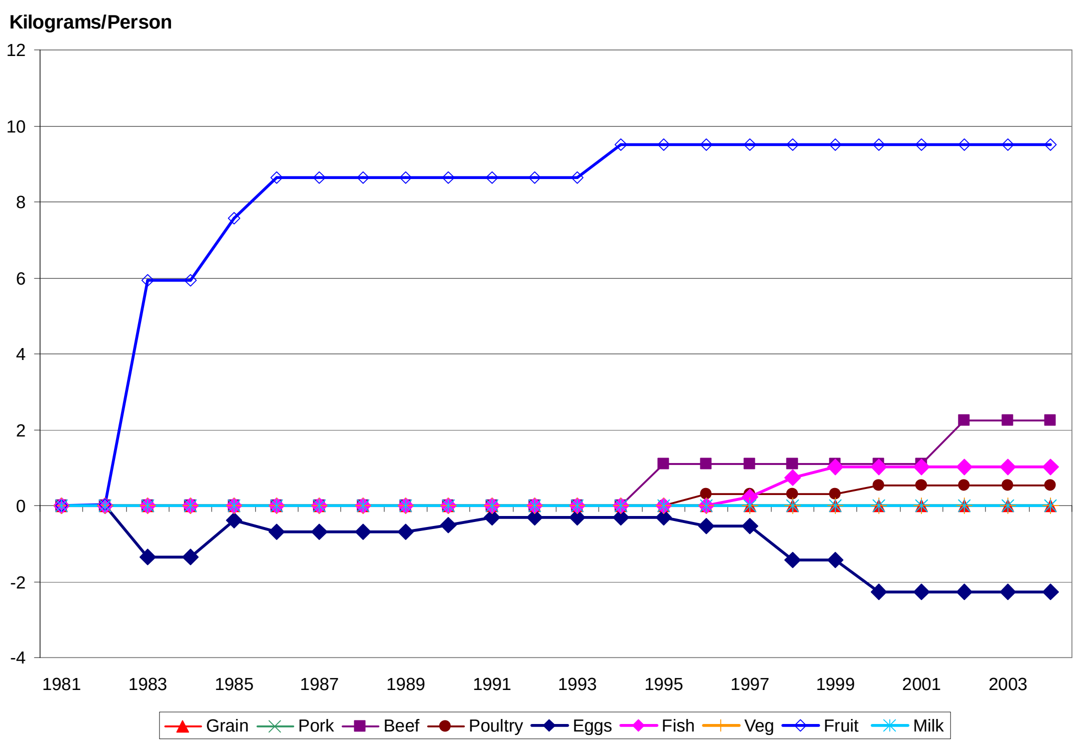

Figure 3. Cumulative taste changes in quantities: 10-commodity analysis

Conclusions

The objective of this paper is to uncover evidence of structural change in food consumption among urban residents in China. The battery of tests applied to data from the period from 1981 to 2004 provided a reasonably clear picture of changing food consumption. First, both parametric and nonparametric tests indicated that the early 1980s and mid- to late 1990s were likely periods 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. From 1980 to 1985, the output of fruits and freshwater aquaculture products in China increased by 71% and 130%, respectively. 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, pork, and vegetable consumption can be largely explained by normal price and income effects. In contrast, fruits, fish, beef, and poultry products, while not absent from traditional Chinese diets, have played a less important role in daily food consumption, particularly on a regional basis. These less-prominent foods 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 beef, fish, and, to some extent, poultry may not represent a true globalization of diets but may be evidence of an expansion of consumer food purchases to include goods that are part of the national diet but may not be included in local or regional diets. An important observation supporting this notion is the fact that structural change associated with these products occurs in the latter half of the 1990s. This period coincides with the rapid development of private retail food chains and the creation of more regional and national food markets.

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. This paper provides some empirical evidence that structural change has occurred in urban Chinese food demand, but further research with richer data sets is needed.

Endnote

1. Beef and mutton are always aggregated into one meat group in China's statistical system, so references to beef throughout the remainder of the paper include both beef and mutton.

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Appendix Tables

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

	Grain	Meat	Fish	Vegetables	Fruit	Eggs	Other Foods
	-0.1063	-0.1698	-0.1624	-0.1529	-0.1642	-0.0516	0.2184
Grain	(0.2270)	(0.0520)	(0.0411)	(0.0467)	(0.0443)	(0.0181)	(0.0861)
	-0.1721	-0.6765	0.0792	-0.0698	0.0699	0.0623	-0.3289
Meat	(0.0210)	(0.0413)	(0.0097)	(0.0083)	(0.0089)	(0.0081)	(0.0432)
	-0.3797	0.2494	-0.7851	0.4305	-0.1170	0.1121	-0.4727
Fish	(0.0693)	(0.0468)	(0.0395)	(0.0803)	(0.0218)	(0.0211)	(0.0897)
	-0.1916	-0.0482	0.2924	-0.6193	0.0217	0.0798	-0.1231
Vegetables	(0.0456)	(0.0167)	(0.0531)	(0.0592)	(0.0047)	(0.0101)	(0.0141)
	-0.3693	0.1615	-0.1157	-0.0260	-0.6218	0.0910	-0.4820
Fruit	(0.0581)	(0.0262)	(0.0170)	(0.0041)	(0.0578)	(0.0137)	(0.0738)
	-0.1393	0.4058	0.2119	0.2170	0.1131	-0.4018	-0.7969
Egg	(0.0606)	(0.1055)	(0.0653)	(0.0528)	(0.0320)	(0.1645)	(0.1849)
	-0.0208	-0.1919	-0.0909	-0.0989	-0.0452	-0.1155	-0.6901
Other Foods	(0.0124)	(0.0387)	(0.0141)	(0.0202)	(0.0093)	(0.0226)	(0.0708)
	0.5888	1.0359	0.9625	0.5882	1.1582	0.3903	1.2533
Expenditure	(0.1108)	(0.0045)	(0.0070)	(0.0717)	(0.0235)	(0.1746)	(0.0437)

Table A2. 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)