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Testing for Separability and Structural Change in Urban Chinese Food Demand

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Abstract: This paper estimates and tests separable demand specifications in food expenditures by urban consumers in China. Both parametric and nonparametric techniques are employed to obtain information about shifts in consumer preferences and appropriate commodity groupings. The results have implications for past and future research and policy analysis.

Key Words: Separability, Food Consumption, Structural Change, China, Nonparametric

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China's economic reforms, which began in 1978, resulted in remarkable growth in GDP, averaging 8-9 percent annually in the latter half of the 1990s. Per capita nominal GDP increased from 379 Yuan in 1978 to 8,184 Yuan in 2002. Economic prosperity and domestic policy reforms have changed the market environment for consumers in China. The removal of rationing, increased abundance and variety of foods, and changes in the marketing system have likely contributed to shifts in consumer preferences. 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 78.48 kg in 2002 in urban areas, whereas the per capita consumption of meats, eggs, and aquatic product increased from 20.52kg, 5.22kg, 7.26kg in 1981 to 32.52kg, 10.56kg, and 13.20kg in 2002, respectively (State statistical bureau of China).

Given the size of China's population and the potential impacts of even minor changes in consumption behavior on international agricultural markets, numerous demand studies have been conducted to estimate the impacts of the market reforms and income growth on consumption patterns. Many of these studies have utilized the idea of separability in their demand specification to limit the scope of their analysis to particular commodities or food aggregates. However, few demand studies that employ separability assumptions actually test for the validity of these assumptions, given their commodity groupings. Inappropriate aggregations can lead to faulty demand elasticity estimates that may provide misleading results from hypothesis tests, projections, and policy analysis (Moschini et al; Nicol).

This paper contributes to the growing literature on Chinese food consumption in two respects. First, we utilize the structure described by Moschini and Meilke to estimate and test for shifts in parameter values over time for important commodity groupings, such as grains, meats, vegetables, and aquatic products. Knowledge of the nature of parameter shifts in recent years will aid in projecting and accommodating future changes in policy analysis. Second, we exploit the structural change results to test separable demand specifications commonly used in the analysis of food expenditures by urban consumers in China.

The organization of the paper is as follows. First, we describe the methodology for testing structural change using both parametric and nonparametric approaches. Then we conduct the empirical tests of structural change for Chinese urban food demand. Following the structural change tests, we briefly discuss methods for empirically testing weak separability in demand. Based on the findings from structural change tests, we test a number of separable demand structures for Chinese food consumption. We conclude with a brief summary of our findings and suggestions for further research.

Methodology for Testing Structural Change

Test for Structural Change

In demand analysis, structural change is often referred to “changing tastes and preferences” (Moschini and Moro). These changes can be reflected by the change in the shape of individual utility functions. In this study, both nonparametric and parametric methods are used to investigate structural changes in Chinese urban food demand. Two of the most popular functional forms, the linear version of the Almost Ideal Demand System (LA/AIDS) and the Rotterdam model, are both estimated to check the robustness of our results.

The LA/AIDS model developed by Deaton and Muellbauer 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 is expressed as

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

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 a the translog price index below.

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

We linearize the model by replacing the translog price index above with the "corrected" Stone price index, $\log(P_s) = \sum_{i=1}^{n} w_i \log(\frac{p_i}{p_i^0})$, where p_i^0 is mean price (Moschini).

Adding up, homogeneity, and symmetry imply the following parameter restrictions.

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

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 general linear AIDS model can be reparameterized as follows.

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

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

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

where 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.

To capture the dynamic behavior of Chinese food demand, the first difference form is used for estimation; thus, the estimated model is given in equation (2)

$$\Delta w_{it} = \gamma_i \Delta h_t + \sum_{j=1}^{n} (\gamma_{ij} \Delta \log(p_{jt}) + a_{ij} \Delta (h_t \log(p_{jt})) + \beta_i \Delta \log(\frac{y_t}{P}) + b_i \Delta (h_t \log(\frac{y_t}{P})) \quad (2)$$

A testing the hypothesis that the time path parameters (γ_i, a_{ij}, and b_i) are equal to zero is same as a test of the hypothesis of no structural change.

The structural change version of the Rotterdam model is similar to the LA/AIDS model. By introducing a time path variable, the Rotterdam model is re-specified in equation (3)

$$w_{it} d \log(q_{it}) = (c_i + k_i h_t) \sum_{j=1}^{n} w_{jt} d \log(q_{jt}) + \sum_{j=1}^{n} (c_{ij} + k_{ij} h_t) d \log(p_{jt}) \quad (3)$$

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

$$\begin{aligned} &\textstyle\sum_j c_j = 1, \sum_j c_{ij} = \sum_i c_{ij} = 0, c_{ij} = c_{ji}, \\ &\textstyle\sum_j k_j = 0, \sum_j k_{ij} = \sum_i k_{ij} = 0, k_{ij} = k_{ji}. \end{aligned}$$

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

Nonparametric Tests of Stable Preferences

The parametric approach for testing structural change in consumer preferences is ultimately dependant on 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). Repeating the analysis using a variety of flexible functional forms is one means of assessing the robustness of evidence of preference shifts in a particular data set. However, nonparametric methods provide an alternative approach that does not require any assumptions regarding functional forms.

The 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, Houthakker, and Afriat, Varian (1982) demonstrated that 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 1 holds for the sequence of bundles $\{x_j, x_k, x_l, \ldots, x\}$.

$$p_j x_j \geq p_j x_k, p_k x_k \geq p_k x_l, \ldots, p_m x_m \geq p_m x \qquad (4)$$

In equation 4, 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 $x_j Rx$, 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).

Questions have been raised about the power of nonparametric tests, particularly when the real expenditures grow rapidly over time. Income effects may mask shifts in the underlying preferences by causing each successive consumption bundle to lie outside of the consumptions set of the previous observation, despite relative price changes. Real food expenditures for urban Chinese consumers increased an average of 3.3% annually since 1981. Consequently, income effects may be a potential problem in our data set.

Chalfant and Alston suggested using prior information about income elasticities to adjust the expenditure data as a means of removing the effects of income growth from the analysis, allowing the potential impacts of structural change to be observed in the residual data. Applying a similar concept, Sakong and Hayes argue that the impacts of shifts in consumer preferences could be isolated from income and price effects using the compensated demand curve. When all goods in the consumption set are normal goods and preferences are stable, changes in consumption from one period to the next can be described by movements along the compensated demand curve. If an optimal consumption bundle lies off the compensated demand curve, the distance from the demand curve to the observed bundle is a measure of taste changes. Sakong and Hayes proposed a linear programming system that combines the constraints implied by revealed preference theory with a priori information about income elasticities to solve for the optimal set of consumption bundles that minimize taste changes, given observed prices and expenditures.

Data

Annual data from 1981-2002 for per capita consumption, expenditures, and retail prices are obtained from *Chinese Urban Household Income and Expenditure Survey* and *China Statistical Yearbooks*. Ten food commodities (grain, pork, beef/mutton, poultry, eggs, fish, vegetables, fruit, milk, and other foods) were included in the nonparametric analysis. Pork is the most commonly consumed meat product with relatively low price. Beef currently only represents a small proportion of total meat products consumed and per capita beef and especially mutton consumption is much higher in the western pastoral provinces. Beef and mutton are always aggregated into one meat group in China's statistical system. Expenditures on other foods was calculated by deducting food expenditure on those nine commodity groups from total food expenditure. The consumer price index was used as the price of other foods and quantity was calculated from food expenditure by dividing by price index. In the parametric analysis, the limited number of observations and the relatively large number of parameters prevent inclusion of all ten commodities. Seven food groups (grain, meat, eggs, fish, vegetables, fruit, and others) are included in the parametric analysis. The Tornqvist quantity index is used for aggregating pork, beef/mutton, and poultry into meat. And aggregate meat price is recovered from meat expenditure divided by meat quantity. In addition, in order to compare the grouping effects on structural change and separability tests, an alternative grouping is also tested. A five-commodity group, which includes grain, meat and eggs, fish, vegetables/fruit, and other foods, is used to conduct the tests. Tornqvist quantity indices are also used for meat and egg aggregates, and vegetables and fruit aggregates. All prices and income were normalized by their sample mean.

Empirical Results from Structural Change Tests

Parametric Tests

The dynamic linear AIDS model in (2) and the Rotterdam model in (3) were estimated using the maximum likelihood estimation procedure in TSP 4.5. Both two models have six equations with the equation for other foods omitted to avoid singularity problem. Homogeneity, symmetry, and adding-up restrictions were imposed on the model parameters. There are 210 possible combinations of $\boldsymbol{\tau}_1$ and $\boldsymbol{\tau}_2$, and the two demand systems were estimated for each combination assuming seven and five commodity groups. With the limitation of degree of freedom, not all sets of combinations can have estimable parameters of the system of equations. For AIDS model, all the parameters of system of equations are estimable for $1981 \leq \boldsymbol{\tau}_1 \leq 1994$ and$1989 \leq \boldsymbol{\tau}_2 \leq 2002$; and for Rotterdam, all the parameters are estimable for $1982 \leq \boldsymbol{\tau}_1 \leq 1995$ and $1989 \leq \boldsymbol{\tau}_2 \leq 2002$.

The maximum likelihood estimates of the structural change points ($\boldsymbol{\tau}_1$, $\boldsymbol{\tau}_2$) are shown in Table 1. These results indicate a rather abrupt change in the middle of the observation period. Based on log-likelihood ratio tests, the additional structural change points in Table 1 are combinations of ($\boldsymbol{\tau}_1$, $\boldsymbol{\tau}_2$) which cannot be rejected with five commodity groups in both AIDS and Rotterdam models.

Understanding the identified possible structural change points, requires some knowledge of the urban food rationing policy in China. The Chinese food rationing began in 1953 to guarantee the food security for urban residents. Rationed foods were obtained from 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 foods with rationing coupons. From 1978, Chinese central government started the economic reforms. During the first phase of reform, 1978-1984, the state monopoly or prescribe purchases of agricultural produce decreased dramatically, for example, from 113 kinds in 1981 to 60 kinds in 1984. Then, farmers were allowed to sell their surplus produce in free markets. However, since there were still rigid institutional constraints on the free markets and the market prices were much higher than rationing prices, during 1978-1984, the rationing system still dominated free market in urban food supply. Then in late 1984, price controls for 15 nonstaple foods, including pork, eggs, sugar, and vegetables, were lifted, but the government distribution of these goods was still maintained with ration prices well below free market prices. It was not until 1987-1988 that rationing of the 15 nonstaple foods was totally eliminated. After that, only grains and edible oils were subject to rationing. In the following years, with increasing urban household income and abundant supply of farm produce, the free market became more and more important in the dual market system and finally led to the abolition of the rationing system in 1993 and grain and edible oils became nonbinding.

Our empirical results of structural change tests are very consistent with those policy change points, although different groups capture different change points. With seven commodity groups, both AIDS model and Rotterdam capture the point of elimination of grain rationing, which occurred in 1993. With five commodity groups, Rotterdam model captures both points where eliminations of nonstaple foods and grains rationing occurred. AIDS model with five commodity groups has more extensive indications. It captures not only the exact policy change points but also the policy transforming periods.

From above results, we can see that policy change has dominating effects on Chinese urban food demand shift. In addition, grouping also affects the indication of structural change points by models. In this study, the more disaggregated model with seven food groups is only able to capture the point of elimination of grain rationing; and more aggregated model captures more policy change points. One explanation for this result may be that disaggregating the data disperses the rationing effects on food demand across the individual commodity groups, and the effects of the eliminating grain rationing may dominate the effects of other, somewhat more gradual policy transformations in specific commodities. On contrary, when commodities experiencing similar market transformation 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.

To investigate the significance of the structural change, conditional on the optimal combination of ($\boldsymbol{\tau}_1$, $\boldsymbol{\tau}_2$), we conducted likelihood ratio tests for the hypothesis of constancy of the parameter vector over time, i.e. coefficients for time path variables equal to zero. The results were reported in Table 2 for both AIDS and Rotterdam models with seven and five commodity groups, respectively. The hypothesis of no structural change in the full set of parameters is rejected at 0.05 significance level, suggesting that a constant set of parameters cannot be postulated to characterize Chinese urban consumer behavior within the assumed models and that 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 structural changes. We do not reject the hypothesis of no structural change for price parameters at 0.05 significance level in Rotterdam model with five commodity groups; however, intercept and income indicate significant structural change. All tests in other models and under other aggregation assumptions rejected the hypothesis of no structural change, which suggest that those changes not only result from price and expenditure change, but also from intercept change.

The average marshallian price elasticities for AIDS model with seven and five commodity groups before and after the optimal structural change points are reported in Table 3 and Table 4, respectively. With seven commodity groups, except grains, all own price elasticities are negative. The positive own price elasticity for grain mainly results from the transformation of policy and market development. With rationing and meager supply of foods as well as scarce variety of food supply, Chinese urban food demand was skewed in favor of grains. With rationing reform and increasing food supply and varieties, there was an adjustment in urban food demand. Therefore, during this period and hereafter, even though grain prices went down, the demand for grains also went down. This could be confirmed by income elasticities, which show that grains change from necessities to inferior goods. After the structural change, except grains and meat, all other food demands become more elastic. Moreover, fish, vegetables, and other foods change from price inelastic to price elastic. Grains change from necessities to inferior goods and vegetables change from inferior to luxury goods. Meat and fruits change from necessities to luxuries, and oppositely, eggs change from luxuries to necessities. In both regimes, meat shows substitution relationship with both eggs and vegetables, while meat shows complementary relationship with both grains and fish.

For comparison, the model with five commodity groups was evaluated at a similar structural change point. We computed elasticities at the (1991, 1993) structural change point. The results were reported in Table 4. Before structural change, fish's own price elasticity has positive sign, and the same occurs for grain after the structural change. Same interpretation of the results for grains as with seven commodity groups could be applied here. The change in the own-price elasticity for fish may be due to the diversity and quality variation of the fish products included in the aggregate. Along with the deepened reform and market development, aquatic products quality and variety were increased greatly. However, the data cannot show quality change. Although prices increased with average quality of aquaculture products, demand also increased, potentially generating the positive own price elasticity before the structural change. Except grain, all other foods became more price elastic after structural change. Like the analysis with seven commodity groups, grains changed from necessities to inferior goods, whereas fish changed from inferior to luxury goods. The aggregate of meat and egg changed from income inelastic to income elastic. Grains are complementary to meat/egg, while meat/egg is substitution to vegetable/fruit in both regimes.

Comparing the elasticity estimates from both grouping, we can conclude that different grouping do have some degree of effects on elasticity estimations. It not only affects values but also sometimes signs. To further analyze exclusive structural change effects on quantity demand and also grouping effects, we estimated the bias of structural change holding prices and expenditure constant for both seven and five commodity groups for AIDS model.

Let B_i denote the change of share before and after the structural change, $B_i=w_i^a-w_i^b$, where w_i^a is the share after the structural change and w_i^b is the same share before the structural change. If B_i >0, then the structural change favors the ith good; and if B_i <0, then the structural change is against the ith good. Evaluating the shares at the sample mean of the exogenous variables, from equation (1), B_i could be expressed as $B_i = \boldsymbol{\gamma}_i$. The estimated bias of structural change and its standard errors were reported in Table 5. The results show that the structural change is significantly biased against grain and in favor of fish in both seven and five groups. In seven groups, the structural change is neutral for meat and fruit, and significantly against eggs and vegetables. In five groups, the structural change is neutral for both meat/egg, and vegetable/fruit. Compared with mean shares before structure change, it shows that there was about 28% decrease in grain share after the structural change for seven and five commodity groups, respectively, while there was about as high as 85% increase for fish with seven commodity groups and about 23% increase with five commodity groups.

Nonparametric Tests of Structural Change

The algorithms suggested by Varian (1982) for checking direct revealed preference and transitive closure were applied to the Chinese urban consumption data using three different aggregations of the 10 primary food categories. GARP was initially tested on all 10 food groups, and no violations were found. As in the parametric analysis, beef, pork, and poultry were aggregated into a meat group using a Tornqvist quantity index. Repeating the GARP test on the set of seven commodities revealed a violation in the relationship between observations in 1981 and 1985. Finally, eggs were added to the meat group, and fruits and vegetables were combined into a single aggregate. The resulting 5 food groups were tested for consistency with GARP and a violation was identified in the relationship between the observations in 1996 and 1997.

The test results may provide some evidence of structural change in 1985 and in 1997; however, it is also possible that the violations indicate a rejection of our separability assumptions rather than preference stability. As Chalfant and Alston note, it is not possible to distinguish whether aggregation bias, omitted goods, or preference shifts generate violations of GARP; however, failure to find inconsistencies in the data suggest that one's separability assumptions do not create significant problems. In the present analysis, the fact that the group of 10 commodities did not violate GARP, while other aggregations of the same data did generate violations, points toward aggregation bias and not structural change as the likely cause. The consistency of the more disaggregated data with GARP does not rule out the possibility of structural change

We applied the Sakong and Hayes model to the Chinese urban consumption data using the same three aggregation assumptions applied above. Like Cortez and Senauer, we incorporate the adjustments to the model suggested by Chalfant and Zhang to avoid dependence of the test results on scaling and price deflator choices. Our expenditure elasticity assumptions were derived from a brief survey of the literature on urban food consumption in China. In particular, the model allows the analyst to select a range over which the expenditure elasticity for each may vary. We attempted to select bounds that would include the majority of the estimates found in the literature. Some sensitivity analysis was conducted to determine whether the broadening the selected ranges would significantly alter the results. While the magnitudes of some taste changes did vary, particularly grain, the qualitative result did not change substantially. Table 7 displays the ranges selected for this study.

Figures 1-3 display the cumulative taste changes for the three simulations. It is interesting to note that the cumulative taste changes in all three aggregations display a negative shift in grain consumption around 1985 and a positive shift in demand for fruits. This time period roughly corresponds to the period of market liberalization for vegetables, fruits, aquatic products, and livestock products. In the mid-1980s, the two-track price system was introduced for grains, and procurement quotas for most non-grain products were eliminated, giving farmers more flexibility and incentives to increase production of fruits, vegetables, and livestock products (Tuan and Ke). Consequently, availability of these products increased in urban areas, perhaps leading to the shifts on consumption displayed in Figures 1-3.

Preferences appear to be fairly stable from the mid 1980s until the mid 1990s, but milk, beef, and fish consumption all show increases in the later half of the 1990s. During the same period, egg and grain consumption experience negative changes. Again, these shifts coincide with substantial changes in China's food marketing system. Urban grain rationing ended in 1993 in most urban areas, and retail food marketing in the larger coastal cities began to rapidly move toward supermarkets as a primary channel for food delivery in the late 1990s. Both changes increased the variety of food products available to urban consumers, as well as relative prices. The results of this analysis are consistent with the observation that market reforms and income growth in China have prompted consumers to move away from staple foods, such as grains and eggs, toward more diverse diets including greater quantities of fruit, fish, and dairy products. Somewhat surprising is the fact that pork and poultry did not display any taste change. Price and income effects appear to be sufficient to explain changes in consumption of these two meats.

Tests for Separability

Parametric Tests

To limit the scope of analysis to particular commodities or food aggregates, or to use aggregate data when disaggregated data are unavailable or of poor quality, many studies have utilized the idea of separability in demand specification. With separability of preferences, the commodities can be partitioned into groups so that choices within groups can be determined are only impacted by other goods through group expenditures. Consequently, the decision behavior can be explained through a much smaller number of variables. Separability opens up the possibility of multi-stage budgeting and decision making. Given the convenience of the assumption of separability, many studies of Chinese food consumption use this idea without testing its validity (e.g., Wang et al, Gao, Wailes, and Cramer, Wang and Chern, Wu, Li, and

Samuel, Lewis and Andrews, and Wu and Wu). To authors' knowledge, only Fan and Chern briefly mention the separability test using Varian's method in their study of Chinese urban food consumption patterns for the period of 1985-1990.

Subsequently, we use tests for weak separability that are based on a combination of results from Goldman and Uzawa, and Blackorby, Davidson, and Schworm. For detailed developments the reader is referred to Moschini, Moro, and Green. Let σ_{ij} be Allen-Uzawa elasticity of substitution between goods i and j; and ε_i be income elasticities. For any two goods (i, j) $\in I_g$ (group g) and a good k $\in I_s$ (group s) (g≠s), the necessary and sufficient conditions for goods i and j weakly separable from k can be expressed in elasticity terms as

$$\frac{\sigma_{ik}}{\sigma_{jk}} = \frac{\varepsilon_i}{\varepsilon_j}$$

The Allen-Uzawa elasticity $\sigma_{ij}=e_{ij}/w_j$, where e_{ij} is the compensated price elasticity; and ε_i is income elasticity.

From AIDS model (2), Marshallian demand elasticities that reflect the effects of structural change are

$$\begin{aligned} e_{iit} &= (\gamma_{ii} + b_i h_t)/w_{it} - (\beta_i + b_i h_t) - 1 \\ e_{ijt} &= (\gamma_{ij} + b_i h_t)/w_{it} - (\beta_i + b_i h_t) w_{jt} / w_{it} \end{aligned}$$

And income elasticity that reflects the structural change is

$$\varepsilon_i = (\beta_i + b_i h_t)/w_{it} + 1$$

From slutsky equation, the Hicksian demand elasticities are

$$\begin{aligned} e_{iit} &= (\gamma_{ii} + b_i h_t)/w_i + w_j - 1 \\ e_{ijt} &= (\gamma_{ij} + b_i h_t)/w_i + w_j \end{aligned}$$

Following Moschini, Moro, and Green, the separability restrictions based on information of structural change for the linear AIDS model (2) can be expressed as

$$\frac{\gamma_{ik} + a_{ik} h_t + w_i w_k}{\gamma_{jk} + a_{jk} h_t + w_j w_k} = \frac{w_i + \beta_i + b_i h_t}{w_j + \beta_j + b_j h_t}$$

or

$$\gamma_{ik} = \frac{(w_i + \beta_i + b_i h_t)}{(w_j + \beta_j + b_j h_t)} (\gamma_{jk} + a_{jk} h_t + w_j w_k) - (a_{ik} h_t + w_i w_k)$$

Due to the characteristics of AIDS model that it is separability-inflexible, it is of interest to consider the separability restrictions at the mean point, which we chose the mean shares. In addition, we treat parameters a_{ij} and b_i, and time path variable h_t as known based on the information from structural change estimation in order to make the estimation converge easily.

Unlike AIDS model, Rotterdam model is separability-flexible for the purpose of modeling weak separability (Moschini, Moro, and Green). Therefore, the separability restrictions will hold not only locally, but also globally without any further restrictions. The separability restrictions for Rotterdam model only depend upon coefficients and are much simpler than those for AIDS model. The separability restrictions for Rotterdam model (3) are:

$$\frac{c_{im} + k_{im} h_t}{c_{jm} + k_{jm} h_t} = \frac{c_i + k_i h_t}{c_j + k_j h_t}$$

or

$$c_{im} = \frac{c_i + k_i h_t}{c_j + k_j h_t} (c_{jm} + k_{jm} h_t) - k_{im} h_t$$

Similarly, time path variable and its parameters are treated as known based on the findings from structural change to make the estimation converge easily.

Nonparametric Tests

Varian's 1983 article developed nonparametric testing procedures for weak separability. Suppose we assume that preferences over a set of n goods can be represented by the separable utility function in equation 5.

$$u(x_1, x_2, \ldots, x_k, v(x_l, \ldots, x_n)) \tag{5}$$

Varian showed that any weakly separable utility function can be supported by the data if the goods in the separable group satisfy the following three conditions.

1. Goods $x_l \ldots x_n$ must satisfy GARP as a group.
2. Price and quantity indices $\left({}^{1}\!/_{\mu^i}, V^i \right)$ that satisfy the Afriat Inequality $\left(V^i \leq V^j + \mu^j p^j \left(x^i - x^j\right)\right)$ must exist.
3. The set of goods $\{x_1, \ldots, x_k, V\}$ and the corresponding prices $\{p_1, \ldots, p_k, 1/\mu\}$ satisfy GARP.

The greatest difficulty that must be overcome in applying this nonparametric test of weak separability is identifying appropriate price and quantity indices that satisfy the Afriat inequality. The method proposed by Varian is computationally intensive and has had difficulties identifying separable structures in monte carlo studies (Barnett and Choi). Fleissig and Whitney proposed an alternative method for computing appropriate price and quantity indices using a linear programming system that minimizes the errors applied to a superlative index subject to the Afriat Inequality. Diewert has shown that a superlative index, such as the chained Tornqvist index, provides a second-order approximation to an unknown aggregator function. Fleissig and Whitney use the Tornqvist index augmented with positive and negative error terms to construct price and quantity indices that satisfy the Afriat Inequality. The linear programming solution provides the error-minimizing indices, which are used to test consistency of the assumed separable structure with GARP.

Empirical Tests of Separability

Parametric Results

Since Wald test for separability lacks invariance when nonlinear restrictions are imposed, the size-correcting likelihood ratio test was conducted to test separability (Dagenais and Dufour; and Moschini, Moro, and Green). The tests results for different types of weakly separable structures with seven and five commodity groups, conditional on the optimal ($\boldsymbol{\tau}_1$, $\boldsymbol{\tau}_2$), were reported in Table 6. With seven commodity groups, the results show that all of the nine separability structures are rejected. With five commodity groups, conditional on the optimal ($\boldsymbol{\tau}_1$, $\boldsymbol{\tau}_2$), except the separable model that fish is separable from all other groups cannot be rejected at 0.025 significance level, all other separable models are rejected.

Nonparametric Results

We applied Fleissig and Whitney's approach to the urban consumption data. As in the parametric tests, we checked for asymmetric separability for each commodity. We also tested separability of meat, meat and eggs, fruits and vegetables, and milk and other foods. The test results are displayed in Table 8.

The test for asymmetric separability of individual commodities is essentially a test of whether or not the remaining 9 commodities can be considered a separable group. In most cases, asymmetric separability was not supported. However, the tests of fish and milk could not reject the null hypothesis of asymmetric separability. Eggs and other foods passed the test of condition 3, but the remaining 9 food groups violated condition 1. The separability tests for meat, meat and egg, fruit and vegetable, and milk and other groups failed in all cases, except meat and eggs. Fruits and vegetables had only one violation of condition 3, as did the meat group. Given the low number of violations for the meat and fruit and vegetable groups, separability of these groups may not be ruled out.

Conclusions

Several parametric and nonparametric tests were performed on aggregate food consumption data for urban Chinese residents. Although the results were quite mixed, a number of patterns did emerge from the data. First, both parametric and nonparametric tests revealed evidence of taste changes in the mid-1980s and early to mid 1990s. Both time periods coincide with periods of major change in domestic market policies. Second, both parametric and nonparametric test indicate that taste changes moved against grains and eggs in favor of fish and milk. Results for fruits and vegetables were mixed, but both approaches failed to demonstrate significant structural change for meats. Third, separability tests were also mixed, with the parametric tests rejecting asymmetric separability for all commodities except fish. The parametric tests also rejected separability for meats, meat and eggs, and fruits and vegetables. The nonparametric tests also found evidence of asymmetric separability for fish; however, asymmetric separability for milk also could not be rejected. Of the four commodity groupings tested with nonparametric methods, only the meat and egg aggregate was supported by the data.

Much of the recent empirical work in Chinese food consumption has employed cross-sectional and panel data because of the richness of these data sets compared to the available aggregate time series data. While this study is one of the first to empirically test for evidence of structural change and separability in Chinese food demand, the relevance of our results for the existing literature would be enhanced if the tests utilized in this paper were applied to survey and panel data. Moreover, the diverse results obtained in this study using different parametric specifications and aggregation assumptions indicate that researchers in the future should seek to uncover separable structures in the data and employ alternative functional forms before settling on a final empirical specification.

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Table 1.Maximum Likelihood Structural Change Points

	Dynamic AIDS Model	Rotterdam Model
Seven GroupCommodities	(1990, 1993)	(1992, 1994)
Five Commodity groups	(1984, 1987)	(1993, 1994)
-Additional Points	[(84, 85, 86), (87, 88,93, 94)]*	(1986, 1987)†
	[(82, 83, 88, 91, 92), (93,94)]*	

*The numbers in the first parenthesis are possible values for $\tau 1$, and the numbers in the second parenthesis are possible values for $\tau 2$. These combinations are structural change points that cannot be rejected.
†indicates the possible structural change points cannot be rejected at 0.025 significance level. For AIDS model, (88, 94) cannot be rejected at 0.025 level; all other cannot be rejected at 0.05 significance level.

Table 2. Likelihood Ratios for Structural Change Tests for AIDS and Rotterdam Models with Seven/Five Commodity groups

Hypothesis	Restrictions	Likelihood Ratio	$\chi^2_{0.05}$
AIDS Model			
No Structural Change in:			
-all parameters			
-Seven group	33	326.93	43.77
-Five group	18	74.85	28.87
-intercept parameters			
-Seven group	6	290.24	12.59
-Five group	4	12.32	9.49
-price parameters			
-Seven group	21	317.36	32.67
-Five group	10	45.24	18.31
-expenditure parameters			
-Seven group	6	244.77	12.59
-Five group	4	27.48	9.49
Rotterdam Model			
No Structural Change in:			
-all parameters			
-Seven group	27	209.68	40.11
-Five group	14	68.39	11.07
-price parameters			
-Seven group	21	193.65	32.67
-Five group	10	10.23*	18.31
-expenditure parameters			
-Seven group	6	154.78	12.59
-Five group	4	53.11	9.49

* indicates cannot be rejected at 0.05 significance level.

Table 3. Average Marshallian Elasticities for AIDS Model with 7 Commodity groups

	Grain	Meat	Egg	Fish	Vegetable	Fruit	Other Foods	Expenditure
				Before Structural Change				
Grain	0.218	-0.212	0.279	-0.232	0.265	0.126	-0.521	0.076
Meat	-0.255	-0.757	0.391	-0.156	0.441	0.165	-0.435	0.607
Egg	0.593	1.293	-0.261	0.492	-1.486	-0.084	-2.357	1.810
Fish	-0.737	-0.521	0.580	-0.408	-0.099	0.203	0.853	0.129
Vegetable	0.517	1.064	-0.565	-0.006	-0.024	0.215	-0.559	-0.642
Fruit	0.135	0.399	-0.014	0.104	0.151	-0.405	-1.318	0.947
Other Foods	-1.120	-1.218	-0.545	-0.138	-0.841	-0.602	-0.780	5.243
				After Structural Change				
Grain	0.499	-0.408	-0.129	-0.155	-0.103	0.049	0.980	-0.733
Meat	-0.522	-0.430	0.048	-0.249	0.451	0.001	-0.632	1.334
Egg	-0.611	0.309	-0.495	0.174	0.624	0.142	-0.978	0.834
Fish	-0.364	-0.394	0.098	-1.724	2.339	-0.270	-0.034	0.350
Vegetable	-0.456	0.638	0.180	1.674	-1.900	-0.074	-2.236	2.174
Fruit	-0.172	-0.025	0.051	-0.381	-0.042	-0.507	-0.419	1.497
Other Foods	-0.877	-1.043	-0.384	-0.329	-0.723	-0.323	-2.419	5.540

*structural change point is (1990, 1993).

Table 4. Average Marshallian Elasticities for AIDS Model with 5 Commodity groups

	Grain	Meat/Egg	Fish	Vegetable/Fruit	Other Foods	Expenditure
			Before Structural Change			
Grain	-0.275	-0.171	-0.055	-0.181	0.638	0.043
Meat/Egg	-0.192	-0.003	-0.150	0.336	-0.603	0.612
Fish	-0.123	-0.544	0.142	0.142	0.556	-0.173
Veg/Fruit	-0.259	0.442	-0.006	-0.372	-0.516	0.710
Other Foods	-0.518	-1.615	-0.190	-1.096	-1.635	5.055
			After Structural Change			
Grain	0.385	-0.346	0.076	-0.005	0.257	-0.367
Meat/Egg	-0.399	-0.382	0.082	0.023	-0.698	1.374
Fish	-0.058	0.258	-1.040	-0.292	-0.043	1.174
Veg/Fruit	-0.141	0.128	-0.110	-0.750	0.014	0.860
Other Foods	-0.646	-1.136	-0.327	-0.755	-2.760	5.625

*structural change point is (1991,1993).

Table 5. Bias of Structural Change from AIDS Model

Commodity		Bias	Standard Error	Mean Share Before Structural Change
Seven Groups				
	Grain	-0.047	0.020	0.169
	Meat	0.020	0.017	0.205
	Egg	-0.010	0.005	0.052
	Fish	0.044	0.007	0.052
	Vegetable	-0.044	0.018	0.111
	Fruit	-0.015	0.013	0.074
	Other Foods*	0.005		0.336
Five Groups				
	Grain	-0.025	0.008	0.164
	Meat/Egg	0.009	0.014	0.259
	Fish	0.012	0.004	0.053
	Veg/Fruit	0.003	0.009	0.187
	Other Foods	0.001		0.338

*bias of other foods is recovered from the other groups and we did not calculate the standard errors. For seven groups, the structural change point is (1990, 1993); and for five groups the structural change point is (1991,1993).

Table 6. Size-Corrected Likelihood Ratio Tests of Separability with Seven and Five Commodity groups

Commody separable from others	Size-corrected LR	
	AIDS model	Rotterdam model
Seven Commodity Groups		
Grain	134.96	113.76
Meat	139.34	54.86
Egg	136.38	95.16
Fish	135.57	112.45
Vegetables	180.45	131.9
Fruit	134.71	117.05
Other foods	127.9	124.14
Meat/egg	122.65	144.34
Vegetables/fruit	143.13	96.36
Number of restrictions	5	
$\chi^2_{0.05}$	11.07	
Five Commodity Groups		
Grain	145.27	30.6
Meat	237.53	43.22
Fish	170.83	9.24*
Vegetables	226.77	44.37
Other foods	345.83	13.45
Number of restrictions	3	
$\chi^2_{0.05}$	7.81	
$\chi^2_{0.025}$	9.35	

* indicates cannot be rejected at 0.025 level.

Table 7. Expenditure Elasticity Bounds by Food Group

Food Group	Lower Bound	Upper Bound	Food Group	Lower Bound	Upper Bound
Pork	0.4	1.2	Grain	0.0	0.4
Beef	0.3	1.3	Vegetables	0.2	1.5
Poultry	0.6	1.3	Fruit	0.7	1.5
Eggs	0.2	1.0	Milk	0.7	2.2
Fish	0.6	1.5	Other	0.2	2.5
Meat	0.3	1.3	Fruit & Veg.	0.2	1.5
Meat & Eggs	0.2	1.3	Milk & Other	0.2	2.5

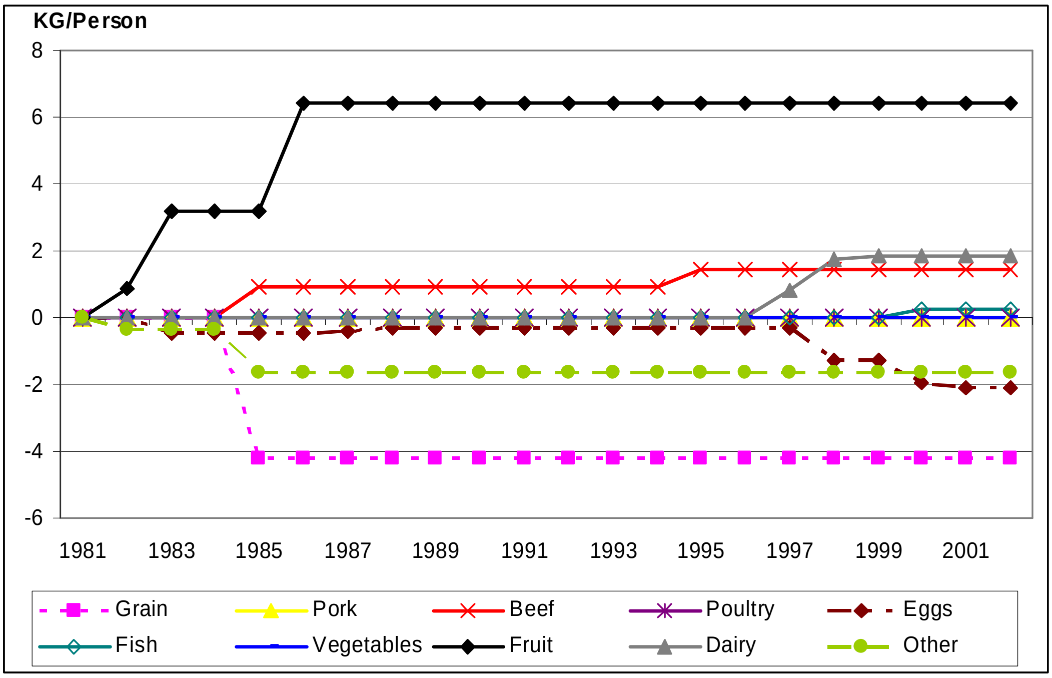

Figure 1. Urban Consumption Cumulative Taste Changes: 10 Food Groups

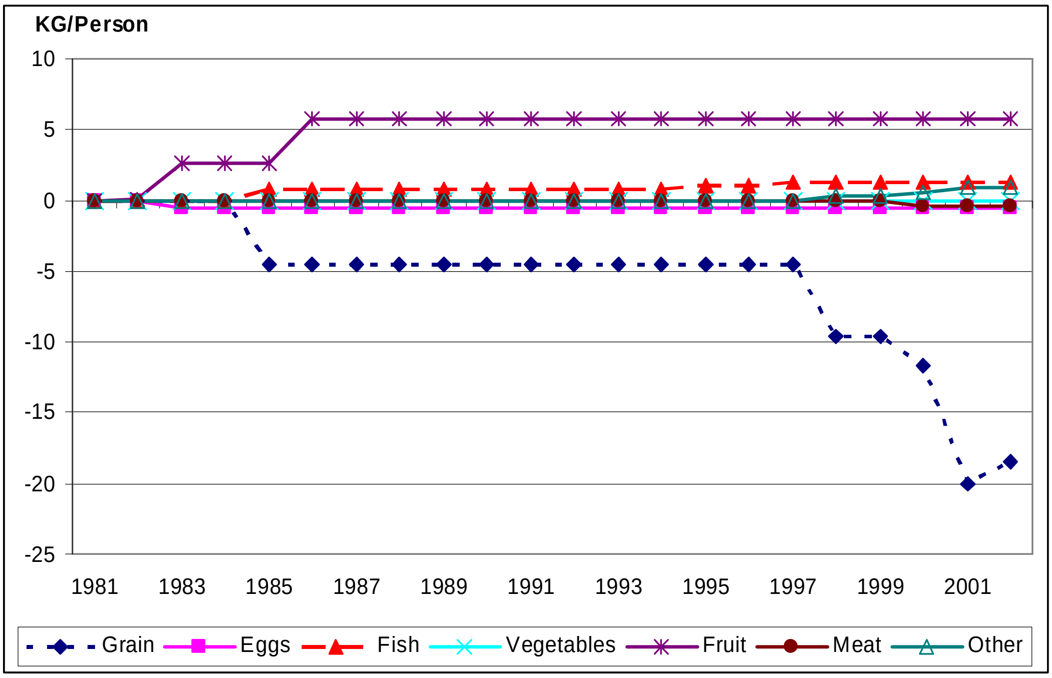

Figure 2. Urban Consumption Cumulative Taste Changes: 7 Food Groups

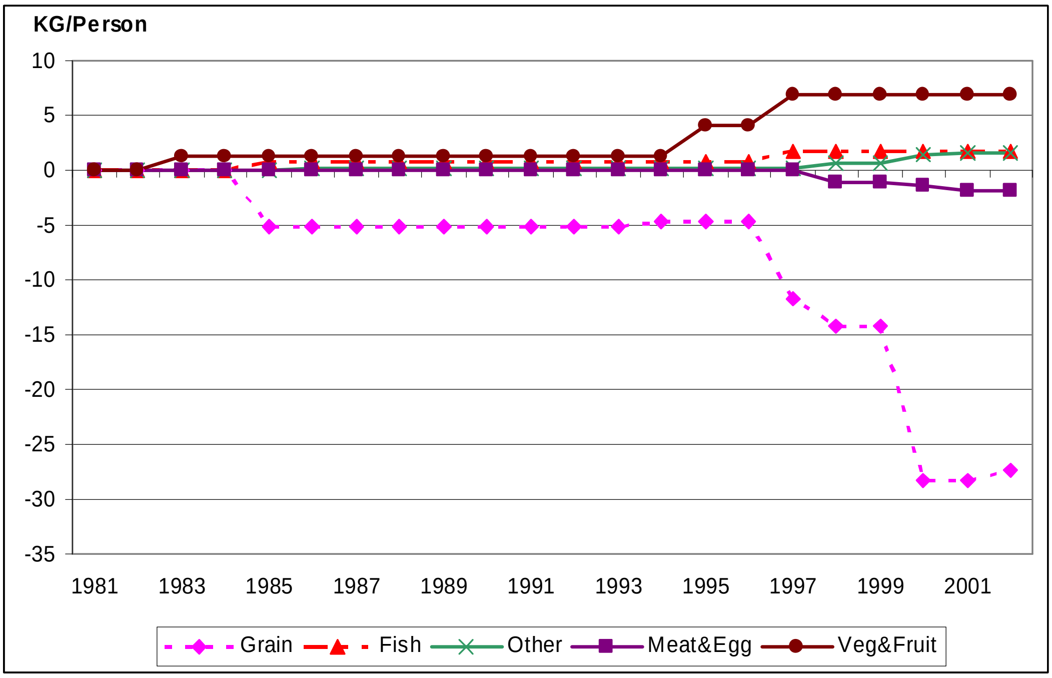

Figure 3. Urban Consumption Cumulative Taste Changes: 5 Food Groups

Table 8. Results from Nonparametric Separability Tests

Separable Commodity/Group	Test of Condition 1	Test of Condition 3
Asymmetric Separability		
Grain	Pass	Fail (5 violations)
Pork	Pass	Fail (12 violations)
Beef	Pass	Fail (1 violation)
Poultry	Fail (1 violation)	Fail (1 violation)
Eggs	Fail (1 violation)	Pass
Fish	Pass	Pass
Vegetables	Pass	Fail (4 violations)
Fruits	Pass	Fail (2 violations)
Milk	Pass	Pass
Other Foods	Fail (1 violation)	Pass
Separable Groups		
Meat	Pass	Fail (1 violation)
Meat & Eggs	Pass	Pass
Fruits & Vegetables	Pass	Fail (1 violation)
Milk & Other	Pass	Fail (11 violations)