# Goodness-of-Fit Test for Price Duration Distributions

Takayasu Matsuoka*
Graduate School of Economics, Osaka University

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**Abstract**

Is the actual price-setting behavior of an individual commodity item consistent with the assumptions of a sticky-price model? Part of the question may formally be addressed by performing a goodness-of-fit test for price duration distributions. For each of the 429 items in the Japanese retail price data for 2000–2005, we fitted the standard parametric models with or without unobserved heterogeneity to the data and tested the goodness of fit. We found that 8.6 percent of the tested items cannot reject the hypothesis that the underlying distribution is exponential, which corresponds to the time-dependent pricing model of Calvo (1983).

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# 1 Introduction

This paper examines the distributional assumption of the duration of price spells. It forms part of an attempt to evaluate the degree of price stickiness and to construct a formal theory dealing with sticky prices. In fact, our analysis has a theoretical motivation, because existing sticky-price models in macroeconomics explicitly formulate the mechanism of a firm's price change by assuming that the length of price spells follows a certain distribution.[1] In this paper, we examine the distributional assumption using a Pearson-type goodness-of-fit test for censored data, which is based on a maximum likelihood estimation of parametric hazard function (Akritas (1988) and Hjort (1990)). Intuitively, the test statistic compares, on a span-by-span basis, the observed number of price changes and the expected number of price changes implied by the hypothesized hazard function. By performing a goodness-of-fit test, we will investigate whether the actual price-setting behavior is consistent with the implications of sticky-price models.

Empirical studies on price stickiness have recently focused on the voluminous price dataset underlying the computations of the Consumer Price Index (CPI) and produced a rapidly growing literature. (See, for example, Bils and Klenow (2004); Nakamura and Steinsson (2007) for the United States; Dhyne et al. (2006) for the European Union; and Higo and Saita (2007) for Japan.[2]) Since then, researchers have discovered stylized facts at the individual item levels as well as at the aggregated level.[3] The interest of our analysis lies in the the process of price changes at the item level. The reason is that the component items of the CPI are highly heterogeneous and, thereby, the estimates of the hazard rate from the pooled sample of several items are influenced by the heterogeneity effect. In this paper, we conduct the goodness-of-fit test

[1]One example is the Calvo (1983) model, which assumes that the probability of a firm's price change is determined exogenously and does not change over time. This assumption implies that price spell durations have an exponential distribution with a constant hazard rate. The other example is the Dotsey, King, and Wolman (1999) model, which assumes a fixed cost of adjusting price. This model predicts a monotonically increasing hazard function when the general level of prices continues upward.

[2]This subject has been studied extensively in many countries other than those listed above. See also Álvarez (2007) for a useful survey.

[3]A remarkable finding at the aggregated level is that the hazard function exhibits a downward slope. Many authors attribute the discrepancy between this empirical finding and the implication of sticky-price models to heterogeneity in price-setting behavior. Using the finite mixture model, Álvarez et al. (2005) and Ikeda and Nishioka (2007) show that the aggregation of several price-setters facing different hazard functions results in a decreasing hazard function. Nakamura and Steinsson (2007) and Matsuoka (2007) reach the same conclusion using a random effect model in which the item-level heterogeneity is incorporated.

for each of the 429 items available in the *Monthly Report on the Retail Price Survey* for 2000–2005.

For the inference of the unknown distribution at the item level, previous studies frequently examine how the nonparametric hazard rate changes with time. For example, by doing this Higo and Saita (2007) classify the item-level hazard functions into three categories: flexible type, Taylor type, and decreasing hazard type. However, it can be problematic to draw conclusions about the distributional assumptions on the basis of the graphical representation. One reason is that in some items, we are quite uncertain about which of these shapes characterizes the process of price change: decreasing, constant, increasing, or a mixture of them. More seriously, the precision of the estimate of the hazard rate gets lower as the elapsed time becomes longer and, therefore, the judgment can change according to how we evaluate the hazard rates at the longer duration, which is less precisely estimated. The other reason is that unobserved heterogeneity in price-setting behavior may exist at the item level. In that case, we cannot judge from the graphical representation whether the decreasing hazard function exhibits true duration dependence or merely the heterogeneity effect.

In contrast to the classification of Higo and Saita (2007), we draw a clear distinction as to whether an item rejects the hazard function with a particular shape. For the inference of the unknown distribution, we hypothesize two parametric models: the exponential model and the Weibull model. Consequently, our analysis formally addresses such questions as sample of how many items are from exponential distribution with a constant hazard rate or question whether an item that retains the hypothesis of Weibull distribution with monotonically increasing hazard rates exists or not. Furthermore, in order to account for the unobserved heterogeneity, we model the gamma-distributed heterogeneity that appears multiplicatively in the hazard function. Fitting the exponential-gamma mixture model and the Weibull-gamma mixture model to the data, we investigate whether the unobserved heterogeneity, to which we presumably attribute the city characteristics or the pricing strategy of an outlet, can account for the process of price changes at the item level.

Our test results can be summarized in the following four findings. First, 8.6 percent of the tested items (in weighted share) cannot reject the hypothesis that the underlying distribution is exponential, which corresponds to the time-dependent model of Calvo (1983). Second, those items that retain the hypothesis of exponentiality have in common a low frequency of price changes. The frequency of price changes for these items is limited to 11 percent per month. Some items in the subgroups such as Eating out in the Food sector, Repairs and maintenance in the Housing sector, and

Recreational goods and services in the Reading and recreation sector lead to the nonrejection of the hypothesis of exponentiality. The results are also verified by the Shapiro and Wilk (1972) test for exponentiality. Third, the well-fitting Weibull hazard models for our data have a duration-dependent parameter that is nearly equal to or less than one, which provides little support for monotonically increasing hazard function at the item level. Fourth, the performance of the model with unobserved heterogeneity is worse than that of the model without unobserved heterogeneity. In view of the results, it is not easy to attribute the reason for the decreasing hazard estimates simply to the unobservable characteristics of the survey cities.

The remainder of the paper is organized as follows. Section 2 formulates the process of price change using the notion of counting processes, on which our test statistic is based. In Section 3, we show the principle of our goodness-of-fit test and the related graphical method using nonparametric estimates of the cumulative hazard rate. In Section 4, we summarize the properties of our four hypothesized models. In Section 5, we show the nonparametric test for an exponential hypothesis by Shapiro and Wilk (1972). In Section 6, after we describe our price data, we discuss the results of the goodness-of-fit test. Section 7 concludes our analysis.

# 2 Counting processes and Martingales

Our test statistic is based on the theory of counting process.[4] A counting process is a stochastic process whose value counts the number of events. It is an increasing (right continuous) process with jumps of size 1. In order to construct the goodness-of-fit test statistic, a good starting point is to describe two counting processes: the failure process and the at-risk process.

In our analysis, "failure" means an observed price change of an item. The price change of the $i$th item ($i = 1, \ldots, I$) in the $j$th city ($j = 1, \ldots, J(i)$) is defined as $P^{ij}_{t^*} \neq P^{ij}_{t^*-1}$, where $t^*$ is a calendar time. In the following analysis, we do not focus on the calendar time when price changes occur, but on the elapsed time over which price remain unchanged, that is, the time between two consecutive price changes. Formally, the elapsed time of the $k$th price spell of the $i$th item in the $j$th city, $t_{ijk}$ ($k = 1, \ldots, K(i, j)$), is defined as follows: let $t^*_{ijk}$ be the calendar time when the $k$th event of price change occurs in the $j$th city. The event indicates the onset of the risk of a price change for the $k$th price spell. Suppose our observation of the $k$th spell is completed, i.e., we observe the price spell until the $k + 1$th (next)

[4] See Kalbfleisch and Prentice (2002) and Andersen et al. (1993) for details.

price change occurs. Then we have the elapsed time of the $k$th price spell, $t_{ijk} \equiv t^*_{ijk+1} - t^*_{ijk}$.

Next suppose that we can no longer observe the $k$th price spell by reason other than its price change and this incomplete state of the observation arises at (calendar) time $c^*_{ijk}$. In this case, we say that the $k$th price spell is right censored at time $c^*_{ijk}$. Right censoring can occur, for example, when a price spell ends with an item substitution, or when the price change has not occurred by the end of the observation period. The censoring time $c_{ijk}$ is also defined as the elapsed time since the last price change and therefore $c_{ijk} \equiv c^*_{ijk} - t^*_{ijk}$.

Now we are in a position to understand the failure and the at-risk processes. Suppose that the $k$th price spell of the $i$th item in the $j$th city remains unchanged for $t_{ijk}$. The failure process, in other words, the price change process, is defined as

$$N_{ijk}(t) \equiv 1(t_{ijk} \leq t). \tag{1}$$

The process of price change is a sequence of zero and one. For example, a price spell with a five-month duration can be represented by the sequence of $\{0\ 0\ 0\ 0\ 1\}$. The at-risk process of price changes can be written as

$$Y_{ijk}(t) \equiv 1(t_{ijk} \geq t, c_{ijk} \geq t), \tag{2}$$

where $c_{ijk}$ denotes the censoring time. The at-risk process takes one as long as we observe that the price does not change just before time $t$. Hence, if $Y_{ijk}(t) = 1$, the price spell is exposed to the risk of price change. These two processes are the basic building blocks on which our test statistic is constructed.

Superposing these individual processes, we obtain the number of observed price spells changed in $[0, t]$, $N_i(t) = \sum_j \sum_k N_{ijk}(t)$ and the number of price spells at risk just before time $t$, $Y_i(t) = \sum_j \sum_k Y_{ijk}(t)$. Denoting a time just before $t$ by $t^-$, we define $dN_i(t) \equiv N_i(t^- + dt) - N_i(t^-)$ as the increment of $N_i(t)$ over the small interval $[t, t+dt)$. Now we are able to construct a model that describes the transition of price spells. The essential part of the model is the hazard function $\lambda_i(t)$ which, multiplied by $Y_i(t)$, implies the expected number of price changes in the $i$th item at time $t$. The model can be written as follows:

$$E[dN_i(t)|\mathcal{F}_{t^-}] = Y_i(t)\lambda_i(t)dt, \tag{3}$$

where $\mathcal{F}_{t^-}$ is an information set during the period $[0, t)$, which is referred to as the filtration of counting process.

For the model defined in Equation (3), consider the following process

$$M_i(t) = N_i(t) - \int_0^t Y_i(s)\lambda_i(s)ds, \qquad \text{for all } t \geq 0. \tag{4}$$

The process $M_i(t)$ is called a counting process martingale[5] and satisfies the condition $E[dM_i(t)|\mathcal{F}_{t-}] = 0$ for all $t$. The second term on the right-hand side is called the compensator of the counting process, which equals the cumulative sum of the expected number of price changes in $[0, t]$. Denoting the compensator by $E_i(t)$, Equation (4) can be written compactly as $M_i(t) = N_i(t) - E_i(t)$. The property of a martingale process plays a crucial role in our test statistic.

# 3 Tests for distributional assumptions

In this section, we will show the two kinds of model checking. One is a graphical check that compares the theoretical cumulative value of the hazard function with the nonparametric estimates of cumulative hazard rate, which is called the Nelson–Aalen estimator.[6] The other is the goodness of fit test statistic on which our analysis is based. They are conceptually equivalent but the latter is more rigorous. In the following description, we shall omit the subscript $i$ to simplify the notation while all quantities are calculated on an item basis.

[5]The first and second term on the right-hand side of Equation (4) are both the sum of individual processes, which is called a superposed process. The important property of the superposed process is that the sum of individual martingale process is also a martingale: Suppose $E[dM_{ijk}(t)|\mathcal{F}_{t-}] = 0$ for all $t$, $j = 1, \ldots, J(i)$, $k = 1, \ldots, K(i, j)$, then $M_i(t) = \sum_j \sum_k M_{ijk}(t)$ is a martingale with respect to $\mathcal{F}_t$ (See Kalbfleisch and Prentice (2002)).

[6]Andersen et al. (1993) illustrate the details of the graphical check.

### 3.1 Graphical check using the Nelson-Aalen estimator

The correct specification of a certain parametric hazard function can be checked by graphically comparing the Nelson–Aalen estimator[7]

$$\hat{A}(t) = \int_0^t \{J(s)/Y(s)\}dN(s) \tag{5}$$

with the parametric cumulative hazard function

$$A(t;\theta) = \int_0^t J(s)\lambda(s;\hat{\theta})ds, \tag{6}$$

where $J(s) \equiv 1(Y(s) > 0)$ so that we can define the processes only in the regions where a price spell is observed. In Equation (6), $\lambda(s;\hat{\theta})$ is a parametric hazard function with the maximum likelihood estimator $\hat{\theta}$. We will specify the hypothesized hazard functions in Section 4. If the difference $\hat{A}(t) - A(t;\hat{\theta})$ is considerably large over $t \in [0, T]$, where $T$ denotes the largest of the observation time, the parametric model is judged as a poor one. The goodness-of-fit test described below gives a criterion for this kind of graphical check.

### 3.2 Goodness-of-fit test

Now we describe the goodness of fit test for the censored data. Our null hypothesis is the following composite hypothesis

$$H_0 : \lambda(t) \in \mathcal{L} = \{\lambda(t;\theta);\ \theta \in \Theta \subset \mathcal{R}^p\}, \tag{7}$$

that is, the true hazard function $\lambda(t)$ belongs to the parametric family of hypothesized hazard functions $\mathcal{L}$. When the null hypothesis holds, $\lambda(t)$ can be specified by $\lambda(t;\theta_0)$, where $\theta_0$ is the $p$-dimensional true parameter vector. Since $\theta_0$ is unknown, we employ the maximum likelihood estimator $\hat{\theta}$ corresponding to the hypothesized hazard function. This test statistic, which is based on the maximum likelihood estimator, is originally derived by Akritas

[7]Equation (5) is expressed by using the notion of the Lebesque–Stieltjes integral. When the integrator $N(s)$ is a right continuous step function, it will have many jumps at each of the points $a_1, a_2, \ldots$, where $\Delta N(a_n) = N(a_n) - N(a_n^-) > 0$. In this case, the Stieltjes integral becomes

$$\int_0^t \{J(s)/Y(s)\}dN(s) = \sum_{n:0<a_n\leq t} J(s)\frac{\Delta N(a_n)}{Y(a_n)}.$$

See Fleming and Harrington (1991) for detail.

(1988) but its general framework is proposed by Hjort (1990). We follow Hjort (1990) and present the principle of the test.

Hjort (1990) derives the limit distribution of the following process

$$H_n(t) = \sqrt{n} \int_0^t K_n(s)J(s)\{(1/Y(s))dN(s) - \lambda(s;\hat{\theta})ds\}, \tag{8}$$

where $K_n(t)$ is an almost surely bounded weighting process.[8] When $K_n(t) = 1$, the statistic reduces to $H_n(t) = \sqrt{n}\{\hat{A}(t) - A(t;\hat{\theta})\}$, which compares the Nelson–Aalen estimates and the parametric cumulative hazard function. Setting $K_n(t) = Y(t)/n$, the statistic now becomes $H_n(t) = \{N(t) - E(t;\hat{\theta})\}/\sqrt{n}$ and resolves into the comparison between the observed and expected number of price changes. Our analysis employs the latter case.

As we mentioned above, the test statistic compares these quantities on a span-by-span basis. Let $0 = a_0 < \cdots < a_m = T$ be a division of analysis time into $m$ cells $I_l = (a_{l-1}, a_l]$, $l = 1, \ldots, m$. We then define

$$Q_{n,l} = \frac{1}{\sqrt{n}}\left[N(a_{l-1}, a_l] - \int_{I_l} Y(s)\lambda(s;\hat{\theta})ds\right] = \frac{1}{\sqrt{n}}(N_l - E_l), \tag{9}$$

and $Q_n = (Q_{n,1}, \ldots, Q_{n,m})'$. The numerator of $Q_{n,l}$ is the difference between the observed and expected number of events in the $l$th interval.[9] Hjort (1990) shows that

$$Q_n \xrightarrow{\mathcal{D}} N(0, R), \tag{10}$$

as $n \to \infty$. On this basis we obtain the following test statistic

$$X_n^2 = Q_n' \hat{R}^- Q_n, \tag{11}$$

where $\hat{R}^-$ is the generalized inverse of any consistent estimator of the covariance matrix $R$. The $m \times m$ covariance matrix $R$ can be written in the following form

$$R = D - S'\Sigma^{-1}S. \tag{12}$$

Note that Equation (12) implies the generalized inverse of $R$

$$R^- = D^{-1} + D^{-1}S'G^- SD^{-1}, \tag{13}$$

[8] The derivation of the limit distribution requires a fundamental convergence theorem, the so-called martingale central limit theorem. For the details of the theorem, see Kalbfleisch and Prentice (2002). The formal proof of the weak convergence of the process in Equation (8) is given by Hjort (1990) and is also available in Andersen et al. (1993).

[9] Hjort (1990) proposes that we should choose the $m$ cells so that each cell contains at least five observations.

where $G^{-}$ is the $p \times p$ generalized inverse of $G = \Sigma - SD^{-1}S'$. The component parts of $R$ are as follows: $D$ is a diagonal matrix with elements $d_l = \int_{I_l} y(s)\lambda(s;\theta_0)ds$, where $y(s)$ is a probability limit of $Y(t)/n$; $S$ is the $p \times m$ matrix $(b_1, \ldots, b_m)$, in which $b_l = \int_{I_l} y(s)\psi(s;\theta_0)\lambda(s;\theta_0)ds$, where $\psi(s;\theta_0)$ is the $p$-dimensional score vector of $\log\lambda(s;\theta_0)$; and $p \times p$ covariance matrix $\Sigma = \int_0^T y(s)\psi(s;\theta_0)\psi(s;\theta_0)'\lambda(s;\theta_0)ds$. These quantities are naturally estimated by

$$\hat{d}_l = E_l/n = n^{-1} \int_{I_l} Y(s)\lambda(s;\hat{\theta})ds, \tag{14}$$

$$\hat{b}_l = n^{-1} \int_{I_l} Y(s)\psi(s;\hat{\theta})\lambda(s;\hat{\theta})ds, \tag{15}$$

$$\hat{\Sigma} = n^{-1} \int_0^T Y(s)\psi(s;\hat{\theta})\psi(s;\hat{\theta})'\lambda(s;\hat{\theta})ds. \tag{16}$$

Using Equations (9) and (13)-(16), we can simplify the test statistic

$$\begin{aligned} X_n^2 &= Q_n'\hat{D}^{-1}Q_n + Q_n'\hat{D}^{-1}\hat{S}'\hat{G}^{-}\hat{S}\hat{D}^{-1}Q_n \\ &= \sum_{l=1}^{m} \frac{(N_l - E_l)^2}{E_l} + V_n'\hat{G}^{-}V_n, \end{aligned} \tag{17}$$

where $V_n = \sqrt{n}\sum_{l=1}^{m}\{(N_l - E_l)/E_l\}\hat{b}_l$ and $\hat{G}^{-} = (\hat{\Sigma} - \sum_{l=1}^{m} \hat{b}_l\hat{b}_l'/\hat{d}_l)^{-}$.

Under the null hypothesis (7), the test statistic converges in distribution to chi-squared distribution with degrees of freedom $df = \text{Rank}(R)$, which is equal to the number of cells, $m$. For a large value of the statistic, say $X_n^2 \geq \chi^2_{\alpha,m}$, where $\chi^2_{\alpha,m}$ is the upper $\alpha$ critical point of a chi-squared distribution with $m$ degrees of freedom, we reject the hypothesis.[10]

**[Table 1 about here.]**

Table 1 illustrates the procedure of the goodness-of-fit test at the item level. Suppose we wish to test the hypothesis that the price duration of Coffee (eating out) is a sample from an exponential distribution. The transition of price spells is summarized in the statistics shown in the first four columns. From this data, we obtain the maximum likelihood estimator that corresponds to the exponential model $\hat{\theta}$ (in this case, $\hat{\theta} = 0.054$, thus the mean duration is 18.5 months). We then divide the observation period into

[10] When the dimension of the parameter vector $p \geq 2$, the computation of the generalized inverse of $\hat{R}$ is quite intractable. Hjort (1990) points out a slightly conservative test procedure that rejects $H_0$ if $X_{0,n}^2 = \sum_{l=1}^{m}(N_l - E_l)^2/E_l \equiv \sum_{l=1}^{m} Z_l \geq \chi^2_{\alpha,m}$. Except for a one-dimensional case, we follow his remark and conduct the conservative version of the test.

m cells. Following the remark shown in Footnote 9, when we observe less than five price changes at a point in the interval, we add the number to the number of price changes at the next point until the total number of price changes in the cell reaches at least five. The boxes in the table indicate that process. For example, we observe 7(=2+5) complete price spells that end with a price change in the first box, which form the observed price changes in the 10th cell, $N_{10}$. Consequently, the number of cells $m$ amounts to 21.

From the calculations shown in the last column in the table, we obtain $X^2_{0,n} = \sum_{l=1}^{m} Z_l = 22.18$. The second term on the right-hand side of Equation (17) is calculated to be 1.28. Therefore, we have the test statistic for exponentiality

$$X^2_n = 22.18 + 1.28 = 23.46, \tag{18}$$

which has approximate chi-squared distribution with $m = 21$ degrees of freedom. The associated p–value is 0.320, which leads us to retain the null hypothesis of exponentiality.

# 4 Parametric models

In order to make the inferences about unknown price duration distributions, we test whether the underlying hazard function belongs to a certain parametric family as shown in the hypothesis (7). For the hypothesized hazard function, we consider the following four parametric models: exponential; Weibull; exponential-gamma mixture; and Weibull-gamma mixture. In this section, we will discuss the features of these models.[11]

## 4.1 Models without unobserved heterogeneity

**Model 1. Exponential model**

$$\lambda(t;\theta) = \mu. \tag{19}$$

The simplest parametric model is the exponential model with a constant hazard rate. In this case, the parameter is one-dimension, i.e., $\theta = \mu$. This model has strong economic implication because it corresponds to the price-setting behavior of Calvo's (1983) model, where price changes occur according to the Poisson process with an incidence rate of $\mu$, which is referred to as the frequency of price changes. It is not difficult to confirm that the maximum likelihood estimator of $\mu$ is identical to the frequency of price changes.[12]

[11] For the details of parametric hazard functions, see Cameron and Trivedi (2005) and Klein and Moeschberger (2003).

[12] See, for example, Fleming and Harrington (1991).

**Model 2. Weibull model**

$$\lambda(t;\theta) = \mu\alpha t^{\alpha-1}. \tag{20}$$

A natural generalization of the exponential model is the Weibull model, which allows the hazard function to change monotonically over time. The parameter is $\theta = (\mu, \alpha)'$, which consists of the scale parameter $\mu$ and the shape parameter $\alpha$. The shape parameter $\alpha$ evaluates the duration dependence since the shape of the hazard function is monotonically increasing (decreasing) if $\alpha$ is more (less) than 1. For $\alpha = 1$, the Weibul distribution is an exponential distribution and exhibits no duration dependence. If the price spells of some items are from the Weibull distribution with $\alpha > 1$, we may support the state dependent model of Dotsey, King, and Wolman (1999), in which the hazard function of price changes monotonically increase under the premise that prices continue to rise.[13]

## 4.2 Models with unobserved heterogeneity

The models we mentioned so far presuppose that the samples within an item are homogeneous in terms of the intensity of price changes. However, there is some reason that unobserved heterogeneity exists even at the item level. As Higo and Saita (2007) state, unobserved heterogeneity may arise because of the differences in a city attribute or the pricing strategy of individual outlet. According to this remark, we also fit the model that allows heterogeneity across cities. Generally, the individual hazard function can be written as

$$\lambda_{jk}(t|\upsilon_j) = \upsilon_j \lambda_0(t), \tag{21}$$

where $\upsilon_j$ is the unobserved heterogeneity that is specific to the $j$th city and $\lambda_0(t)$ is the baseline hazard function. As we can see from Equation (21), the heterogeneity effect $\upsilon_j$ appears multiplicatively and thus shifts the individual hazard function proportional to its baseline hazard. We assume that $\upsilon_j$ is independently and identically distributed as Gamma(1,$\delta$). For the baseline hazard $\lambda_0(t)$, we consider exponential and Weibull hazard functions. We will show the aggregated hazard function for each specification below.

**Model 3. Exponential-gamma mixture model**

[13] A full examination of the Dotsey, King, and Wolman (1999) model requires careful consideration of the firm's relative price, which is beyond the scope of our analysis.

$$\lambda(t;\theta) = \mu[1 + \mu\delta t]^{-1}. \tag{22}$$

Here, we specify the baseline hazard function as constant over time, that is, $\lambda_0(t) = \mu$. The parameters $\mu$ and $\delta$ denote the baseline hazard rate and the (normalized) variance of gamma distribution, respectively. The value of $\delta$ indicates the degree of heterogeneity among individuals. In model 3, even if the individual hazard functions are constant overtime, in other words, they exhibit no duration dependence, the aggregated hazard function can be decreasing according to the value of $\delta$. This is an effect of aggregating across heterogeneous individuals.

**Model 4. Weibull-gamma mixture model**

$$\lambda(t;\theta) = \mu\alpha t^{\alpha-1}[1 + (\mu\delta t^{\alpha})]^{-1}. \tag{23}$$

The baseline hazard function of the last model is Weibull, which allows the individual hazard functions to monotonically change over time. The parameters are the scale parameter $\mu$, the duration dependence parameter $\alpha$, and the heterogeneity variance parameter $\delta$. This model can account for both the heterogeneity effect and duration dependence. A well-fitting Weibull-gamma mixture model can explain the interaction of both effects: The aggregated hazard function is decreasing due to the heterogeneity effect while the individual hazard function exhibits positive duration dependence.

# 5 Shapiro and Wilk test for exponentiality

We have seen general framework for testing distributional assumptions for the censored data so far. As for a test of the composite hypothesis of exponential distribution, there is another class of nonparametric tests. Among these tests, Shapiro and Wilk (1972) propose a basic and fairly simple test for exponentiality.[14] The principle of the test, which is similar to their famous test for normality (Shapiro and Wilk (1965)), is to evaluate the adequacy of the linear regression of the ordered observations on the expected values of the order statistics.

Let $x_1 \le x_2 \le \cdots \le x_N$ denote the $N$ order statistics based on the durations of complete price spells that end with a price change. Then the test statistic

$$W = \frac{(\bar{x} - x_1)^2}{(N-1)\hat{\sigma}^2}, \tag{24}$$

[14] An example of the application to economic issues is the investigation of duration dependence in the American business cycle by Diebold and Rudebusch (1990).

where $\bar{x} = \sum_{i=1}^{N} x_i$ and $\hat{\sigma}^2 = \sum_{i=1}^{N}(x_i - \bar{x})^2/N$. The important property of the test statistic is that the distribution is invariant to the parameter values of exponential distribution and depends only on the sample size of complete spells. Shapiro and Wilk (1972) report the simulated percentage points of the test statistic for sample size $n = 3, \ldots, 100$. In order to test items, whose number of complete price spells exceeds 100, we extend the statistical table of the percentage points for $n > 100$. We obtain the null distribution by generating 5000 samples of the test statistic for each sample size of an item.[15] The test statistic is to be used as a two-tailed test and, therefore, we reject the hypothesis at a significance level $\alpha$ if $W \leq h$ or $k \leq W$, where $h$ and $k$ are the $(\alpha/2)100$th and $(1-\alpha/2)100$th percentiles of the null distribution, respectively.

# 6 Empirical Analysis

## 6.1 Data

We employ a sample of the retail prices underlying the computation of the Japanese CPI. These prices are collected on a monthly basis by the Statistics Bureau, Ministry of Internal Affairs and Communications, and appear in the *Monthly Report on the Retail Price Survey.* Prices are reported for each city with a prefectural government and for cities with a population of 150,000 or more. The analysis covers the period from January 2000 to December 2005.

Preferably, we include as many items as possible, but three groups of items are left out of consideration so as not to create a severe bias in our analysis. First, we exclude items that were newly listed as a component item of the CPI and that were not used for calculation of the CPI during the observation period.[16] Second, we exclude seasonal items such as fruits, vegetables, or clothing in which the maximum length of price spells is fairly short due to the right-censoring. Third, we remove items that the statistical agency does not survey every month. These items include, for example, PTA membership fees that are surveyed in April and September every year.

In addition, we exclude two kinds of price data from the original data set. First, we exclude left-censored price spells, whose duration is unknown to us due to the censoring of the starting time. Second, we exclude the price

[15] Shapiro and Wilk (1972) generate 5000 samples of the test statistic for $n \leq 50$ and $[250,000/n]$ samples for $50 < n \leq 100$.

[16] If a new product replaces an old one, we do not exclude the price data of both products, instead, we treat the price spell of the old product as a right-censored spell, censored at the time when the item substitution occurs.

spells whose number of collected prices in the survey exceeds four. Reported prices in the *Monthly Report on the Retail Price Survey* are an arithmetic mean of prices collected at individual outlets. The number of prices collected in each city ranges from one to forty-two, depending on the characteristics of the items and the size of the city. This exclusion limits the maximum number of summand prices to four and, thus, certainly makes our price data represent the process of price changes at the individual outlets within a survey district.[17]

As a consequence of these exclusions, we consider 429 items, covering 60.1 percent of the Japanese CPI in 2000. Table 2 shows the summary statistics.

**[Table 2 about here.]**

## 6.2 Results and Discussions

The overall results of the goodness-of-fit test are summarized in Table 3.[18] We report the results for each group of expenses in the Japanese CPI to find out how the goodness-of-fit of hypothesized models varies over groups.

The second column in the table shows the frequency of price changes by group. The frequency of price changes in a group is calculated as follows: Let $\mathcal{G}_q$ be the $q$th item group, the frequency in the group is $\sum_{i \in \mathcal{G}_q} w_i \hat{\mu}_i$, where $\hat{\mu}_i$ is the parameter estimates of the exponential model, which is identical to the frequency of price changes as we have seen in Section 4, and $w_i$ is the $i$th item's CPI weight in 2000, which satisfies $\sum_{i=1}^{429} w_i = 1$. The figures clearly show that the degree of price stickiness differs across item groups, supporting the finding of many authors including Bils and Klenow (2004).

**[Table 3 about here.]**

In the exponential model (Model 1), 352 items reject the hypothesis of exponentiality at the one percent significance level. In the course of calculation, we cannot calculate the test statistic $X_n^2$ for 19 items because the number of complete spells is too small and thereby we cannot form any single cell. For the rest of the 25 items, which amount to 8.6 percent in weighted

[17]On these exclusions of items and price spells, we follow Higo and Saita (2007) in most respects. A slight difference lies in the last exclusion, in which they use the price data of 55 middle-sized cities. In fact, the number of prices collected in those cities is also less than or equal to four. We include, however, in addition to price data of those middle-sized cities, price data that are surveyed in large-sized cities, but whose number of collected prices is less than or equal to four.

[18]Table 4 in the Appendix shows the test results by item.

share of nonrejected items, we can conclude that exponentiality holds.[19] Table 3 shows that most items in the subgroups of Eating out in the Food sector and Repairs and maintenance in the Housing sector and some items in the Reading and recreation sector lead to the nonrejection of an exponential hypothesis.

It is notable that such items have in common a low frequency of price changes. This fact is clearly illustrated in Figure 1, in which we superpose the distribution of the parameter estimates of the exponential model for the items that lead to nonrejection (denoted by the black bar) on the distribution for all 492 items (gray bar). We can see from Figure 1 that most parameter values of the well-fitting exponential models distribute in the region up to 0.125, whereas in the higher region they do not, except for one item, Camera, with the parameter value $\hat{\mu}$=0.288.

**[Figure 1 about here.]**

The results from the exponentiality test by Shapiro and Wilk (1972) also verify the evidence. Figure 2 depicts the same parameter distribution as the previous figure, but the black bars in Figure 2 denote the number of items that retain the exponentiality in the Shapiro and Wilk (1972) test. The figure illustrates that the items with the higher value of $\hat{\mu}$ uniformly reject the hypothesis at the one percent significance level. The maximum of $\hat{\mu}$ for the items that retain the hypothesis is 0.11, which implies the frequency of price changes for these items is limited to 11 percent per month.[20]

**[Figure 2 about here.]**

As for the Hjort (1990) test, we can verify the results using the conceptually equivalent method that compares the Nelson–Aalen estimate $\hat{A}(t)$ with the parametric cumulative hazard function $A(t;\hat{\theta})$. We choose four items with

[19]Specifically, we cannot reject the null of exponentiality in 58 (=429–(352+19)) items but the figure includes 33 items with high price flexibility. In these flexible items, the test statistic takes a necessarily small value, because the observed and expected number of price changes is nearly equal in each cell over the interval. This can be understood as a trivial case in the sense that any assumed parametric models can lead to the nonrejection of these samples. Thus, we subtract 33 items from 58 items and report 25 items in Table 3.

[20]In Table 3, there is a difference in the number of retained items between the Shapiro and Wilk (1972) test and the Hjort (1990) test for exponentiality. One reasonable explanation for the difference is that the power of the Shapiro and Wilk (1972) test against local alternatives is weaker than that of the Hjort (1990) test. This is because the Shapiro and Wilk (1972) test tends to accept the null of exponentiality too often in the items, in which the Hjort (1990) test confirms that the goodness-of-fit of the Weibull model is plausible, as we can see from Table 4 in the Appendix.

a different value of parameter estimate $\hat{\mu}$ in Figure 3: (a) Coffee (eating out) with $\hat{\mu} = 0.054$, (b) Babies' clothes with $\hat{\mu} = 0.202$, (c) Butter with $\hat{\mu} = 0.399$, and (d) Detergent, laundry with $\hat{\mu} = 0.604$. As we have seen in the last part of Section 3.2, Coffee (eating out) retains the null of exponentiality. Panel (a) shows a typical appearance of both plots when the hypothesis of exponentiality holds. Specifically, the Nelson–Aalen estimate in Panel (a) fluctuates around the predicted value of the cumulative hazard function. We illustrate cases of the rejection in Panels (b)-(d). The predicted values in these panels fail to capture the shapes of the Nelson-Aalen estimate especially in the longer duration.

[**Figure 3 about here.**]

The Weibull model (Model 2) fits well for 54 items, the weighted share of which amounts to 14.6 percent. Now the interest lies in the duration dependence parameter $\alpha$ of the model. If the parameter $\alpha$ exceeds one and the goodness of fit is plausible, we then conclude that the hazard rate of that item tends to increase over time. Figure 4 shows, however, there exist no such items in our dataset. The items that retain the hypothesis that the underlying distribution is Weibull exhibit a decreasing or constant pattern of hazard rates. We can find three items suitable for the Weibull fit (denoted in black) in the region where $\alpha$ exceeds one, but the value of the parameter is nearly equal to one. Thus, the hazard function of these items does not sharply increase over time. In fact, the three items above also retain the hypothesis of exponentiality. When $\alpha$ is close to one, an item that cannot reject a Weibull hypothesis tends to retain an exponential hypothesis as well, which accounts for the results in Table 3 that both hypotheses hold in some items in the group such as Eating out or Repairs and maintenance. The result is highly predictable because Weibull models encompass exponential models.

[**Figure 4 about here.**]

Figure 5 illustrates the graphical check for Weibull models. In the same manner as Figure 2, we show the case of nonrejection in Panel (a), Women's haircut charges with $\hat{\mu} = 0.058$ and $\hat{\alpha} = 0.758$, and the cases of rejection in Panels, (b) Gas with $\hat{\mu} = 0.130$ and $\hat{\alpha} = 0.998$, (c) Microwave ovens with $\hat{\mu} = 0.366$ and $\hat{\alpha} = 1.21$, and (d) Gasoline (regular) with $\hat{\mu} = 0.371$ and $\hat{\alpha} = 1.51$. Note that the estimated Weibull hazard function of four items has different implications as to duration dependence: (a) negative duration dependence; (b) no duration dependence; and (c)–(d) positive duration dependence. The Weibull model for Women's haircut charges in Panel (a) has a concave cumulative hazard function, which means the hazard rates decrease over time. In Panel (a) we observe that the step function of the Nelson–Aalen estimate hovers around the cumulative hazard function, while in Panels (b)–(d) we do not observe the same.

**[Figure 5 about here.]**

Finally, we discuss the results of the model with unobserved heterogeneity. As we have shown in Section 4, if samples are from a heterogeneous population, the estimated hazard rates decrease over time due to the heterogeneity effect. As Higo and Saita (2007) state, even if we estimate the hazard function by item, the estimates can be affected by the local characteristics of a city or an outlet. That is the reason why we estimate the parametric hazard model with unobserved heterogeneity (Models 3 and 4). However, the performance of these models is worse than that of the models without heterogeneity: The weighted share of Model 3 and 4 is even smaller than those of Models 1 and 2. Most items that retain the hypothesis of a heterogeneous model also retain the hypothesis of a homogeneous model, and on comparing the numbers of nonrejected items by groups, the introduction of the gamma-distributed heterogeneity does not additionally increase the number of items of nonrejection. In view of these results, it is not easy to attribute the reason for the decreasing hazard estimates simply to the unobservable characteristics of the survey cities.

# 7 Concluding remarks

In this paper, we examine the distributional assumptions of price duration distributions. We use the goodness-of-fit test by Hjort (1990), in which we compare, on a span-by-span basis, the number of observed price changes and expected price changes implied by the hypothesized hazard function. For each of the 429 items available in the *Monthly Report on the Retail Price Survey* for 2000–2005, we fit the standard parametric hazard models with or without unobserved heterogeneity and test the goodness-of-fit.

We establish the following four facts. First, 8.6 percent of the tested items (in weighted share) cannot reject the hypothesis that the underlying distribution is exponential, which corresponds to the time-dependent model of Calvo (1983). Second, most of the items with a high frequency of price changes reject the hypothesis of exponentiality at the one percent significance level. This finding is also confirmed by the Shapiro and Wilk (1972) test for exponentiality. Third, a well-fitting Weibull hazard model has a decreasing or almost constant hazard function. In our dataset, there is no item that retains a Weibull hypothesis and that exhibits a sharply increasing hazard function. Fourth, the introduction of unobserved heterogeneity that is specific to a survey city does not improve the performance of the model without unobserved heterogeneity at the item level.

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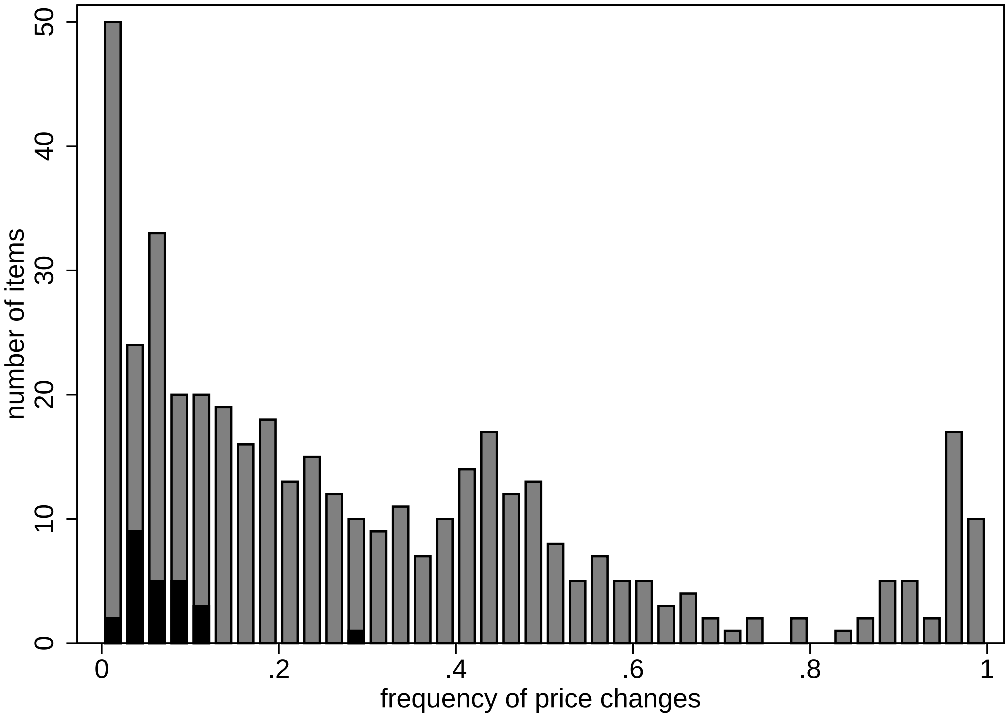


Figure 1: Distribution of $\hat{\mu}$: The black bar denotes the number of items that retain the hypothesis of exponentiality. The test statistic is the Hjort (1990) test of goodness-of-fit. The total number of items is 429. The retail price data are from Jan 2000–Dec 2005.

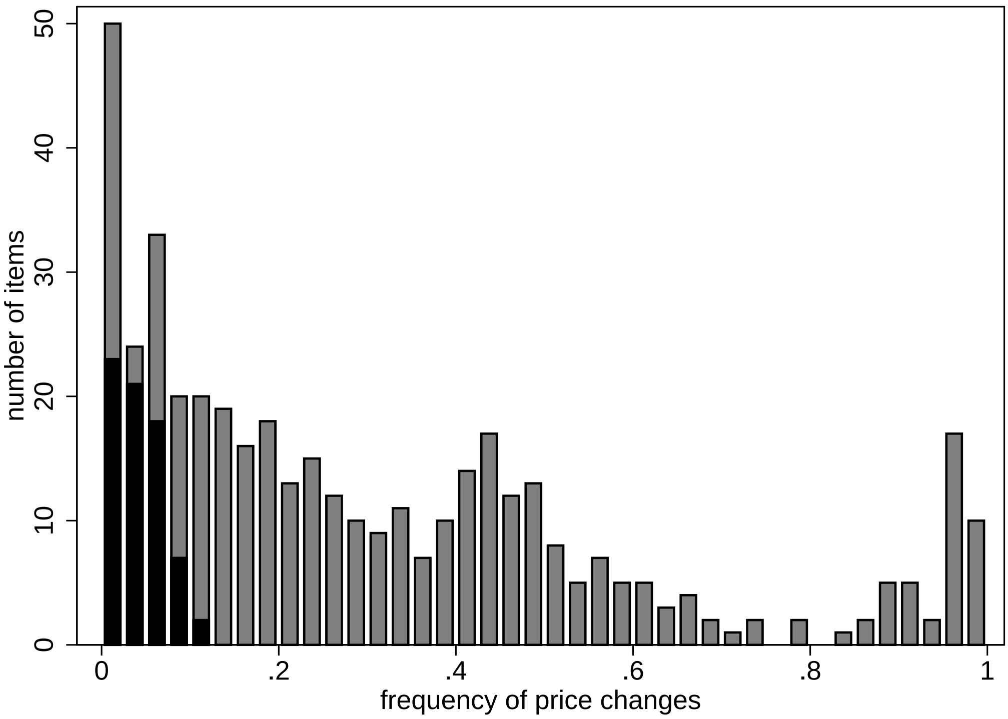


Figure 2: Distribution of $\hat{\mu}$: The black bar denotes the number of items that retain the hypothesis of exponentiality. The test statistic is the Shapiro–Wilk (1972) test for Exponentiality. The total number of items is 429. The retail price data are from Jan 2000–Dec 2005.

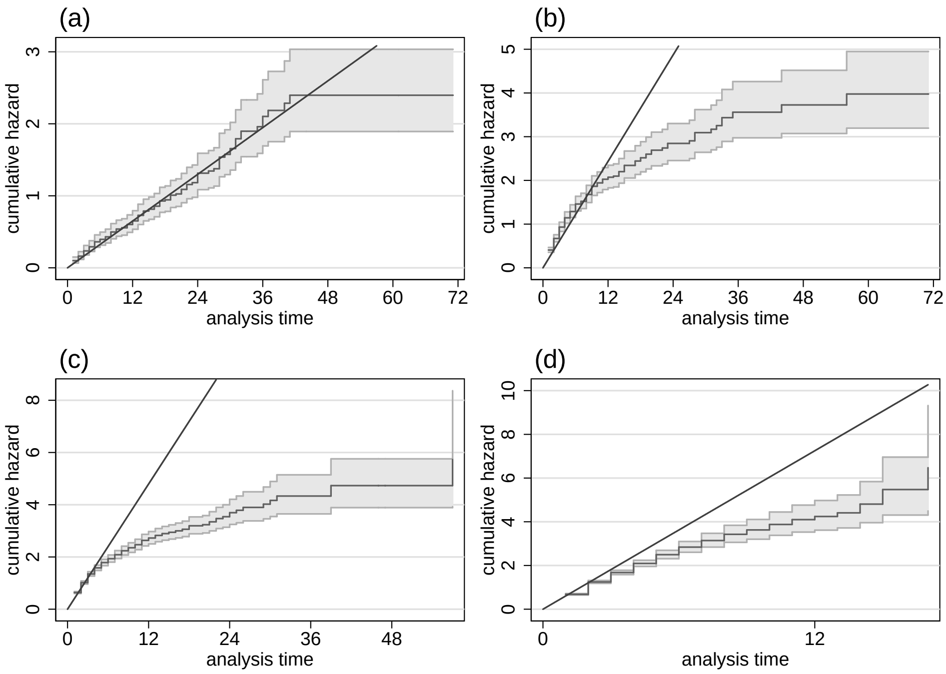


Figure 3: Graphical check for exponentiality: Nelson–Aalen estimate with 95 percent confidence band (step function) and the exponential cumulative hazard function (straight line). (a) Coffee (eating out) with $\hat{\mu} = 0.054$, (b) Babies' clothes with $\hat{\mu} = 0.202$, (c) Butter with $\hat{\mu} = 0.399$, and (d) Detergent, laundry with $\hat{\mu} = 0.604$.

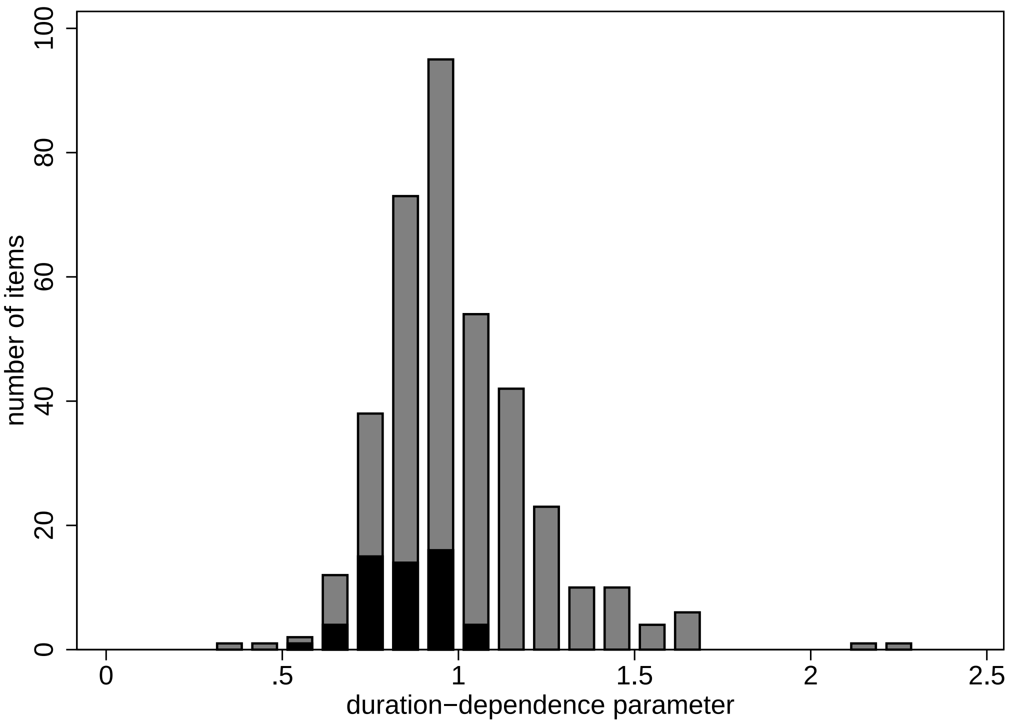


Figure 4: Distribution of $\hat{\alpha}$: The black bar denotes the number of items that retain the hypothesis that the underlying distribution is Weibull distribution. Same data as Figure 1.

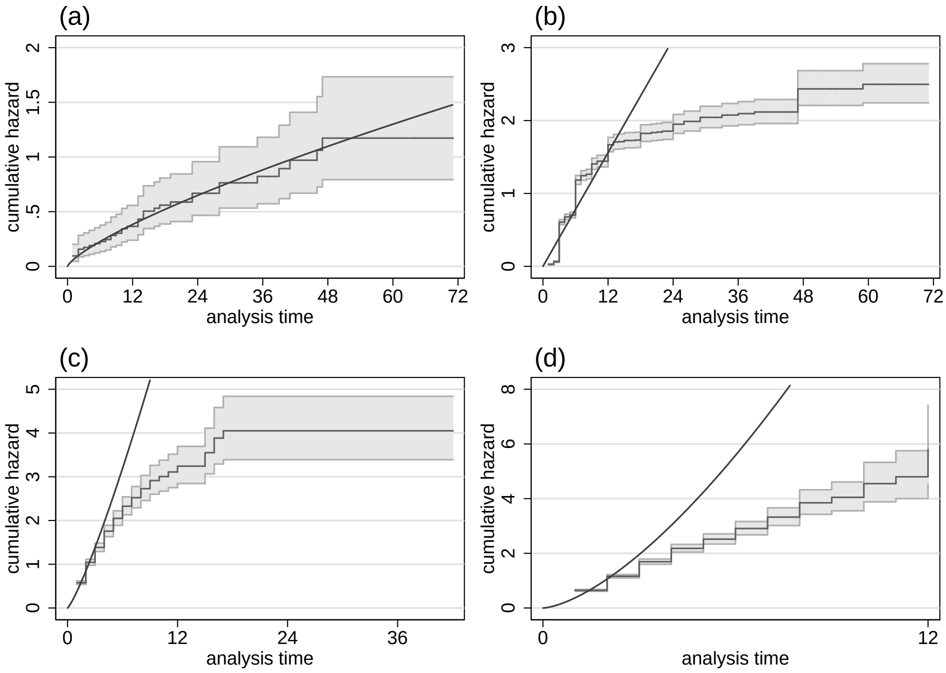


Figure 5: Graphical check for Weibull hazard model: Nelson–Aalen estimate with 95 percent confidence band (step function) and the Weibull cumulative hazard function (smooth line). (a) Women's haircut charges with $\hat{\mu} = 0.058$ and $\hat{\alpha} = 0.758$, (b) Gas with $\hat{\mu} = 0.130$ and $\hat{\alpha} = 0.998$, (c) Microwave ovens with $\hat{\mu} = 0.366$ and $\hat{\alpha} = 1.21$, and (d) Gasoline (regular) with $\hat{\mu} = 0.371$ and $\hat{\alpha} = 1.51$.

| Duration | Number of price spells | | | Cell | Goodness-of-fit test | | |
|---|---|---|---|---|---|---|---|
| $t$ | $Y(t)$ | $N(t)$ | $C(t)$ | $l$ | $N_l$ | $E_l$ | $Z_l$ |
| 1 | 231 | 23 | 0 | 1 | 23 | 12.5 | 8.8 |
| 2 | 208 | 13 | 3 | 2 | 13 | 11.3 | 0.3 |
| 3 | 192 | 14 | 3 | 3 | 14 | 10.4 | 1.3 |
| 4 | 175 | 10 | 6 | 4 | 10 | 9.5 | 0.0 |
| 5 | 159 | 11 | 1 | 5 | 11 | 8.6 | 0.7 |
| 6 | 147 | 5 | 1 | 6 | 5 | 8.0 | 1.1 |
| 7 | 141 | 5 | 2 | 7 | 5 | 7.6 | 0.9 |
| 8 | 134 | 9 | 3 | 8 | 9 | 7.2 | 0.4 |
| 9 | 122 | 5 | 1 | 9 | 5 | 6.6 | 0.4 |
| 10 | 116 | 2 | 5 | 10 | 7 | 12.2 | 2.2 |
| 11 | 109 | 5 | 4 | | | | |
| 12 | 100 | 5 | 4 | 11 | 5 | 5.4 | 0.0 |
| 13 | 91 | 7 | 0 | 12 | 7 | 4.9 | 0.9 |
| 14 | 84 | 5 | 1 | 13 | 5 | 4.5 | 0.0 |
| 15 | 78 | 2 | 4 | 14 | 5 | 8.1 | 1.2 |
| 16 | 72 | 3 | 0 | | | | |
| 17 | 69 | 5 | 1 | 15 | 5 | 3.7 | 0.4 |
| 18 | 63 | 1 | 0 | 16 | 5 | 6.8 | 0.5 |
| 19 | 62 | 4 | 2 | | | | |
| 20 | 56 | 1 | 7 | 17 | 7 | 8.0 | 0.1 |
| 21 | 48 | 3 | 1 | | | | |
| 22 | 44 | 3 | 2 | | | | |
| 23 | 39 | 1 | 0 | 18 | 6 | 4.2 | 0.8 |
| 24 | 38 | 5 | 0 | | | | |
| 26 | 33 | 1 | 0 | 19 | 7 | 5.2 | 0.6 |
| 27 | 32 | 1 | 0 | | | | |
| 28 | 31 | 5 | 0 | | | | |
| 29 | 26 | 1 | 0 | 20 | 6 | 3.9 | 1.1 |
| 30 | 25 | 2 | 1 | | | | |
| 31 | 22 | 3 | 0 | | | | |
| 32 | 19 | 2 | 0 | 21 | 8 | 6.3 | 0.4 |
| 33 | 17 | 0 | 1 | | | | |
| 35 | 16 | 1 | 1 | | | | |
| 36 | 14 | 2 | 0 | | | | |
| 37 | 12 | 1 | 1 | | | | |
| 40 | 10 | 1 | 0 | | | | |
| 41 | 9 | 1 | 0 | | | | |
| 44 | 8 | 0 | 1 | | | | |
| 61 | 7 | 0 | 2 | | | | |
| 71 | 5 | 0 | 5 | | | | |

Table 1: Goodness-of-fit test for exponentiality: Coffee (eating out). $Y(t)$, the number of price spells at risk just before $t$; $N(t)$ the number of price spells end with a price change at $t$; $C(t)$, the number of price spells censored at $t$; $N_l$, the observed number of price changes in the $l$th cell; $E_l$, the expected number of price changes implied by the exponential distribution with hazard rate $\hat{\theta} = 0.054$. $Z_l$ is defined as $(N_l - E_l)^2/E_l$.

| | |
|---|---|
| Number of items | 429 |
| Number of price records | 1,460,823 |
| Number of price spells | 477,210 |
| Number of complete spells | 450,445 |
| Number of right-censored spells | 26,765 |
| Share of right-censored spells | 5.6% |

Table 2: Retail price data used for the calculation of the Japanese CPI from Jan 2000–Dec 2005.

| Groups | Frequency of price changes | Hjort (1990) test of goodness-of-fit | | | | Shapiro-Wilk (1972) test for exponentiality | Number of items tested |
|---|---|---|---|---|---|---|---|
| | | Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | | |
| Food | | | | | | | |
| Cereals | 0.440 | 0 | 0 | 0 | 0 | 0 | 13 |
| Fish & shellfish | 0.741 | 0 | 0 | 0 | 0 | 0 | 29 |
| Meat | 0.477 | 0 | 0 | 0 | 0 | 0 | 10 |
| Dairy products & eggs | 0.477 | 0 | 0 | 0 | 0 | 1 | 8 |
| Vegetables & seaweeds | 0.727 | 0 | 0 | 0 | 0 | 0 | 36 |
| Fruits | 0.946 | 0 | 0 | 0 | 0 | 0 | 5 |
| Oils, fats & seasonings | 0.432 | 0 | 0 | 0 | 0 | 1 | 16 |
| Cakes & candies | 0.325 | 0 | 0 | 0 | 0 | 0 | 17 |
| Cooked food | 0.333 | 0 | 0 | 0 | 0 | 0 | 14 |
| Beverages | 0.338 | 0 | 0 | 0 | 0 | 1 | 14 |
| Alcoholic beverages | 0.250 | 0 | 0 | 0 | 0 | 0 | 13 |
| Eating out | 0.073 | 9 | 12 | 7 | 5 | 12 | 17 |
| Housing | | | | | | | |
| Rent | . | | | | | | |
| Repairs & maintenance | 0.036 | 6 | 9 | 4 | 4 | 9 | 12 |
| Fuel, Light & Water Charges | | | | | | | |
| Electricity | 0.070 | 0 | 0 | 0 | 0 | 0 | 1 |
| Gas | 0.112 | 0 | 0 | 0 | 0 | 0 | 4 |
| Furniture & Household Utensils | | | | | | | |
| Household durables | 0.464 | 0 | 1 | 0 | 0 | 0 | 15 |
| Interior furnishings | 0.227 | 0 | 0 | 0 | 0 | 0 | 5 |
| Bedding | 0.261 | 0 | 0 | 0 | 0 | 0 | 4 |
| Domestic utensils | 0.154 | 1 | 2 | 0 | 0 | 3 | 15 |
| Domestic non-durables | 0.481 | 0 | 0 | 0 | 0 | 0 | 8 |
| Domestic services | 0.006 | 0 | 0 | 0 | 0 | 1 | 3 |
| Cloths & Footwear | | | | | | | |
| Japanese clothing | 0.088 | 0 | 0 | 0 | 0 | 0 | 2 |
| Clothing | 0.256 | 0 | 0 | 0 | 0 | 0 | 4 |
| Shirts & sweaters | 0.215 | 0 | 0 | 0 | 0 | 0 | 1 |
| Underwear | 0.178 | 0 | 0 | 0 | 0 | 0 | 13 |

| Groups (*continued*) | Frequency of price changes | Hjort (1990) test of goodness-of-fit: Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | Shapiro-Wilk (1972) test for exponentiality | Number of items tested |
|---|---|---|---|---|---|---|---|
| Cloth & thread | 0.115 | 0 | 0 | 0 | 0 | 0 | 3 |
| Other clothing | 0.161 | 0 | 1 | 0 | 0 | 0 | 6 |
| Services related to clothing | 0.049 | 0 | 1 | 0 | 0 | 1 | 2 |
| Medical Care | | | | | | | |
| Medicines & health fortification | 0.098 | 0 | 3 | 0 | 0 | 4 | 11 |
| Medical supplies & appliances | 0.252 | 1 | 1 | 2 | 1 | 1 | 9 |
| Medical services | 0.008 | 0 | 1 | 0 | 0 | 1 | 5 |
| Transportation & Communication | | | | | | | |
| Public transportation | 0.001 | 0 | 0 | 0 | 0 | 1 | 9 |
| Private transportation | 0.323 | 2 | 4 | 1 | 2 | 7 | 12 |
| Communication | 0.023 | 0 | 1 | 0 | 1 | 2 | 4 |
| Education | | | | | | | |
| School fees | . | | | | | | |
| Tutorial fees | 0.062 | 0 | 0 | 0 | 0 | 1 | 1 |
| Reading & Recreation | | | | | | | |
| Recreational durables | 0.516 | 1 | 0 | 1 | 0 | 0 | 5 |
| Recreational goods | 0.292 | 3 | 7 | 1 | 1 | 6 | 31 |
| Recreational services | 0.075 | 2 | 5 | 1 | 1 | 13 | 19 |
| Miscellaneous | | | | | | | |
| Personal care services | 0.026 | 0 | 3 | 0 | 1 | 6 | 7 |
| Toilet articles | 0.286 | 0 | 1 | 0 | 0 | 1 | 12 |
| Personal effects | 0.165 | 0 | 2 | 0 | 0 | 1 | 9 |
| Other | 0.005 | 0 | 0 | 0 | 0 | 1 | 5 |
| Total | 0.247 | 25 | 54 | 17 | 16 | 74 | 429 |
| | - | (8.6) | (14.6) | (6.1) | (5.0) | (21.3) | - |

Table 3: The number of items that retain the hypothesis that the underlying distribution belongs to a specified family of distributions. Entries for groups are the subgroups of the ten major groups in the Japanese CPI. The frequency of price changes is calculated as $\sum_{i \in \mathcal{G}_q} w_i \hat{\mu}_i$, where $\mathcal{G}_q$ is the $q$th item group and $w_i$ is $i$th item's CPI weight in 2000 with $\sum_{i=1}^{429} w_i = 1$ ($\hat{\mu}_i$ is defined in the text). The weighted share (in percentage) is in the parentheses in the last row.

# Appendix

| No. | Item | Frequency of price changes | Hjort (1990) test of goodness-of-fit: Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | Shapiro–Wilk (1972) test for exponentiality |
|---|---|---|---|---|---|---|---|
| | | | **[Food]** | | | | |
| 1 | Rice A (domestic), "Koshihikari" | 0.435 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 2 | Rice B (domestic), non-blended rice excluding "Koshihikari" | 0.486 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 3 | Blended rice | 0.382 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 4 | Glutinous rice | 0.340 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 5 | White bread | 0.469 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 6 | Bean-jam buns | 0.276 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 7 | Boiled noodles | 0.373 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 8 | Dried noodles | 0.427 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 9 | Spaghetti | 0.519 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 10 | Instant noodles | 0.523 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 11 | Uncooked Chinese noodles | 0.577 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 12 | Wheat flour | 0.482 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 13 | "Mochi", rice-cakes | 0.411 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 14 | Tuna fish | 0.864 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 15 | Horse mackerel | 0.975 | 0.999† | - | - | - | reject |
| 16 | Sardines | 0.973 | 0.975† | - | - | - | reject |
| 17 | Flounder | 0.963 | 0.145† | - | - | - | reject |
| 18 | Salmon | 0.829 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 19 | Mackerel | 0.955 | 0.931† | - | - | - | reject |
| 20 | Saury | 0.916 | 0.088† | - | - | - | reject |
| 21 | Sea bream | 0.899 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 22 | Yellowtail | 0.916 | 0.000 | - | - | - | reject |
| 23 | Cuttlefish | 0.969 | 0.984† | - | - | - | reject |
| 24 | Octopus | 0.796 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 25 | Prawns | 0.870 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 26 | Short-necked clams | 0.778 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 27 | Scallops | 0.882 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 28 | Salted salmon | 0.700 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 29 | Salted cod roe | 0.539 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 30 | "Shirasu-boshi", dried young sardines | 0.636 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 31 | Dried horse mackerel | 0.742 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 32 | Dried sardines | 0.659 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 33 | "Niboshi", dried small sardines | 0.397 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 34 | Capelin | 0.714 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 35 | "Agekamaboko", fried fish-paste patties | 0.521 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 36 | "Chikuwa", baked fish-paste bars | 0.437 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 37 | "Kamaboko", steamed fish-paste cakes | 0.400 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 38 | Dried bonito fillets | 0.279 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 39 | Pickled fish | 0.663 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 40 | Fish prepared in soy sauce | 0.301 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 41 | Canned fish | 0.248 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 42 | "Shiokara", salted fish guts | 0.261 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 43 | Beef (loin) | 0.516 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 44 | Beef (shoulder) | 0.559 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 45 | Beef (imported) | 0.729 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 46 | Pork (loin) | 0.425 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 47 | Pork (shoulder) | 0.467 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 48 | Chicken | 0.437 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 49 | Liver | 0.268 | 0.000 | 0.000 | 0.000 | 0.000 | reject |

| No. | Item | Frequency of price changes | Hjort (1990) test of goodness-of-fit | | | | Shapiro–Wilk (1972) test for exponentiality |
|---|---|---|---|---|---|---|---|
| | | | Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | |
| 50 | Ham | 0.368 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 51 | Sausages | 0.390 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 52 | Bacon | 0.316 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 53 | Fresh milk (delivered) | 0.051 | 0.000 | 0.000 | 0.000 | 0.000 | * |
| 54 | Fresh milk (sold in stores) | 0.364 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 55 | Powdered milk | 0.213 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 56 | Yogurt | 0.656 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 57 | Butter | 0.399 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 58 | Cheese | 0.428 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 59 | Cheese (imported) | 0.130 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 60 | Hen eggs | 0.880 | 0.179† | - | - | - | reject |
| 61 | Cabbage | 0.989 | 0.997† | - | - | - | reject |
| 62 | Spinach | 0.998 | 1.000† | - | - | - | reject |
| 63 | Chinese cabbage | 0.986 | 0.984† | - | - | - | reject |
| 64 | Welsh onions | 0.992 | 0.990† | - | - | - | reject |
| 65 | Lettuce | 0.995 | 0.977† | - | - | - | reject |
| 66 | Broccoli | 0.997 | 1.000† | - | - | - | reject |
| 67 | Bean sprouts | 0.338 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 68 | Asparagus | 0.976 | 0.971† | - | - | - | reject |
| 69 | Sweet potatoes | 0.969 | 0.481† | - | - | - | reject |
| 70 | White potatoes | 0.957 | 0.451† | - | - | - | reject |
| 71 | Taros | 0.968 | 0.656† | - | - | - | reject |
| 72 | Radishes | 0.985 | 0.989† | - | - | - | reject |
| 73 | Carrots | 0.974 | 0.316† | - | - | - | reject |
| 74 | Burdocks | 0.964 | 0.011† | - | - | - | reject |
| 75 | Onions | 0.958 | 0.560† | - | - | - | reject |
| 76 | Lotus roots | 0.972 | 0.498† | - | - | - | reject |
| 77 | “Naga-imo” yams | 0.921 | 0.009 | - | - | - | reject |
| 78 | Tomatoes | 0.992 | 0.991† | - | - | - | reject |
| 79 | Green peppers | 0.961 | 0.961† | - | - | - | reject |
| 80 | “Shiitake”, Japanese mushrooms, fresh | 0.961 | 0.332† | - | - | - | reject |
| 81 | “Enokidake”, mushrooms | 0.916 | 0.341† | - | - | - | reject |
| 82 | “Shimeji”, mushrooms | 0.940 | 0.993† | - | - | - | reject |
| 83 | “Azuki”, red beans | 0.313 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 84 | “Shiitake”, Japanese mushrooms, dried | 0.266 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 85 | Laver | 0.417 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 86 | “Wakame”, seaweed | 0.349 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 87 | Dried tangle | 0.436 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 88 | Bean curd | 0.297 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 89 | Fried bean curd | 0.405 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 90 | “Natto”, fermented soybeans | 0.499 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 91 | “Konnyaku”, devil’s-tongue jelly | 0.374 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 92 | “Umeboshi”, pickled plums | 0.418 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 93 | Pickled radishes | 0.476 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 94 | Pickled Chinese cabbage | 0.603 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 95 | Sliced vegetables pickled in soy sauce | 0.415 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 96 | Tangle prepared in soy sauce | 0.491 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 97 | Grapefruits | 0.953 | 0.592† | - | - | - | reject |
| 98 | Oranges | 0.952 | 0.955† | - | - | - | reject |
| 99 | Lemons | 0.960 | 0.785† | - | - | - | reject |
| 100 | Bananas | 0.938 | 0.107† | - | - | - | reject |
| 101 | Kiwi fruits | 0.951 | 0.431† | - | - | - | reject |
| 102 | Edible oil | 0.486 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 103 | Margarine | 0.424 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 104 | Salt | 0.020 | 0.000 | 0.001 | 0.000 | 0.000 | * |

| No. | Item | Frequency of price changes | Hjort (1990) test of goodness-of-fit | | | | Shapiro–Wilk (1972) test for exponentiality |
|---|---|---|---|---|---|---|---|
| | | | Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | |
| 105 | Soy sauce | 0.381 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 106 | Soybean paste | 0.436 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 107 | Sugar | 0.434 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 108 | Vinegar | 0.240 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 109 | Worcester sauce | 0.359 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 110 | Ketchup | 0.498 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 111 | Mayonnaise | 0.400 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 112 | Jam | 0.497 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 113 | Instant curry mix | 0.537 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 114 | Instant dried soup | 0.429 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 115 | Flavor seasonings | 0.492 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 116 | "Furikake", granular flavor seasonings | 0.493 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 117 | Liquid seasonings | 0.453 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 118 | "Yokan", sweet bean jelly | 0.061 | 0.000 | 0.007 | 0.000 | 0.000 | reject |
| 119 | "Manju", bean-jam cakes | 0.254 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 120 | "Daifukumochi", rice cakes stuffed with sweetened bean jam | 0.214 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 121 | "Kasutera", sponge cakes | 0.118 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 122 | Cakes | 0.208 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 123 | Jelly | 0.434 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 124 | Pudding | 0.477 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 125 | Cream puffs | 0.218 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 126 | "Sembei", Japanese rice crackers | 0.524 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 127 | "Sembei", Japanese wheat crackers | 0.377 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 128 | Biscuits | 0.660 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 129 | Potato chips | 0.596 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 130 | Candies | 0.550 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 131 | Chocolate | 0.349 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 132 | Ice cream | 0.227 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 133 | Peanuts | 0.370 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 134 | Chewing gum | 0.072 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 135 | Box lunch | 0.096 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 136 | Rice balls | 0.167 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 137 | Bread like sandwiches put cooked food between bread | 0.394 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 138 | Frozen pilaf | 0.493 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 139 | "Kabayaki", broiled eels | 0.632 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 140 | Salad | 0.433 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 141 | Croquettes | 0.339 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 142 | Cutlets | 0.228 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 143 | Fried chicken | 0.471 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 144 | Ch(i)aotzu | 0.467 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 145 | Frozen croquettes | 0.439 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 146 | Cooked curry | 0.436 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 147 | "Mazegohan no moto", prepared materials to boiled rice with assorted ingredients | 0.440 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 148 | Boiled beans | 0.458 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 149 | Green tea ("Bancha") | 0.261 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 150 | Green tea ("Sencha") | 0.316 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 151 | Black tea | 0.339 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 152 | Instant coffee | 0.409 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 153 | Coffee beans | 0.465 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 154 | Coffee beverages | 0.317 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 155 | Fruit juice | 0.509 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 156 | Beverages which contains juice | 0.271 | 0.000 | 0.000 | 0.000 | 0.000 | reject |

| No. | Item | Frequency of price changes | Hjort (1990) test of goodness-of-fit: Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | Shapiro–Wilk (1972) test for exponentiality |
|---|---|---|---|---|---|---|---|
| 157 | Vegetable juice | 0.347 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 158 | Carbonated beverages | 0.281 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 159 | Fermented lactic drinks, unsterilized (“Calpis”) | 0.430 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 160 | Fermented lactic drinks, unsterilized (“Yakult”) | 0.001 | - | - | - | - | ∗ |
| 161 | Sports soft drinks | 0.439 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 162 | Mineral water | 0.523 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 163 | “Sake A” (finest quality) | 0.118 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 164 | “Sake B” (high quality) | 0.167 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 165 | “Sake C” (medium quality) | 0.287 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 166 | “Shochu”, distilled spirits | 0.116 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 167 | Beer | 0.286 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 168 | Beer (imported) | 0.160 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 169 | Low-malt beer | 0.331 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 170 | Whisky (imported) | 0.300 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 171 | Whisky (43% vol. and over) | 0.168 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 172 | Whisky (40% or more, but less than 43% vo1.) | 0.300 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 173 | Whisky (38% or more, but less than 40% vo1.) | 0.139 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 174 | Wine | 0.246 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 175 | Wine (imported) | 0.177 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 176 | Japanese noodles (eating out) | 0.045 | 0.048∗ | 0.307∗ | 0.007 | 0.011∗ | ∗ |
| 177 | Chinese noodles | 0.047 | 0.000 | 0.064∗ | 0.000 | 0.000 | ∗ |
| 178 | Spaghetti (eating out) | 0.085 | 0.257∗ | 0.320∗ | 0.093∗ | 0.061∗ | ∗ |
| 179 | “Nigiri-zushi”, hand-rolled “Sushi” | 0.072 | 0.000 | 0.008 | 0.000 | 0.000 | reject |
| 180 | “Norimaki”,“Sushi” rolled in laver | 0.075 | 0.004 | 0.009 | 0.000 | 0.000 | ∗ |
| 181 | Chicken & eggs on rice | 0.047 | 0.001 | 0.185∗ | 0.000 | 0.004 | ∗ |
| 182 | “Tendon”, prawns “Tempura” on rice | 0.063 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 183 | Curry & rice | 0.064 | 0.107∗ | 0.172∗ | 0.038∗ | 0.004 | ∗ |
| 184 | Bowl of rice topped with seasoned beef | 0.136 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 185 | Ch(i)aotzu (eating out) | 0.062 | 0.014∗ | 0.059∗ | 0.002 | 0.001 | ∗ |
| 186 | Hamburg steaks | 0.102 | 0.059∗ | 0.135∗ | 0.010∗ | 0.002 | reject |
| 187 | Fried prawns | 0.076 | 0.192∗ | 0.282∗ | 0.052∗ | 0.028∗ | ∗ |
| 188 | Lunch for children | 0.082 | 0.029∗ | 0.030∗ | 0.027∗ | 0.016∗ | ∗ |
| 189 | Hamburgers | 0.142 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 190 | Sandwiches | 0.080 | 0.068∗ | 0.219∗ | 0.033∗ | 0.048∗ | ∗ |
| 191 | Pizza | 0.061 | 0.000 | 0.185∗ | 0.000 | 0.000 | reject |
| 192 | Coffee (eating out) | 0.054 | 0.320∗ | 0.551∗ | 0.025∗ | 0.007 | ∗ |
| | | | **[Housing]** | | | | |
| 193 | Bathtubs | 0.037 | 0.000 | 0.002 | 0.000 | 0.002 | ∗ |
| 194 | Toilet seat with a hot douche | 0.293 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 195 | Hot-water supply equipment | 0.047 | 0.171∗ | 0.582∗ | 0.186∗ | 0.582∗ | ∗ |
| 196 | Board | 0.061 | 0.149∗ | 0.265∗ | 0.133∗ | 0.116∗ | ∗ |
| 197 | Paint | 0.072 | 0.000 | 0.045∗ | 0.000 | 0.001 | reject |
| 198 | “Tatami” reupholstering | 0.042 | 0.000 | 0.062∗ | 0.000 | 0.000 | ∗ |
| 199 | Plastering | 0.029 | 0.021∗ | 0.056∗ | 0.002 | 0.002 | ∗ |
| 200 | Gardening | 0.029 | 0.010∗ | 0.027∗ | 0.012∗ | 0.010∗ | ∗ |
| 201 | Sheet glass replacement | 0.031 | 0.000 | 0.259∗ | 0.000 | 0.004 | ∗ |
| 202 | “Fusuma”, sliding doors reupholstering | 0.017 | 0.856∗ | 0.902∗ | 0.890∗ | 0.875∗ | ∗ |

| No. | Item | Frequency of price changes | Hjort (1990) test of goodness-of-fit: Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | Shapiro–Wilk (1972) test for exponentiality |
|---|---|---|---|---|---|---|---|
| 203 | Carpentering | 0.024 | 0.012∗ | 0.026∗ | 0.002 | 0.001 | ∗ |
| | **[Fuel, Light & Water Charge]** | | | | | | |
| 204 | Fire insurance premium | 0.000 | - | - | - | - | - |
| 205 | Electricity | 0.070 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 206 | Gas | 0.130 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 207 | Kerosene | 0.425 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 208 | Water charges | 0.005 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 209 | Sewerage charges | 0.010 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| | **[Furniture & Household Utensils]** | | | | | | |
| 210 | Microwave ovens | 0.464 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 211 | Electric rice-cookers | 0.553 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 212 | Electric pots | 0.471 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 213 | Gas cooking tables | 0.248 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 214 | Gas water heaters | 0.152 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 215 | Refrigerators | 0.683 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 216 | Vacuum cleaners | 0.571 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 217 | Washing machines | 0.635 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 218 | Sewing machines | 0.115 | 0.001 | 0.014∗ | 0.000 | 0.000 | reject |
| 219 | Electric irons | 0.472 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 220 | Room air conditioners | 0.537 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 221 | Chests of drawers | 0.208 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 222 | Wardrobes | 0.188 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 223 | Sitting tables | 0.193 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 224 | Kitchen cabinets | 0.195 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 225 | Clocks | 0.107 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 226 | Lighting apparatus | 0.233 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 227 | Carpets | 0.253 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 228 | "Goza", rush floor coverings | 0.196 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 229 | Curtains | 0.226 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 230 | Beds | 0.144 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 231 | Quilts | 0.332 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 232 | Sheets | 0.176 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 233 | Quilt covers | 0.204 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 234 | Rice bowls | 0.148 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 235 | Dishes | 0.183 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 236 | Coffee cups & saucers | 0.137 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 237 | Glasses | 0.184 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 238 | Wine glasses | 0.101 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 239 | Sealed kitchenware | 0.232 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 240 | Pans | 0.160 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 241 | Pans (imported) | 0.018 | 0.000 | 0.116∗ | 0.000 | 0.000 | ∗ |
| 242 | Kettles | 0.155 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 243 | Scrubbing brushes | 0.303 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 244 | Shelves for microwave oven | 0.134 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 245 | Fluorescent lamps | 0.098 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 246 | Towels | 0.140 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 247 | Vinyl hose | 0.063 | 0.018∗ | 0.274∗ | 0.000 | 0.000 | ∗ |
| 248 | Clean water equipment | 0.233 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 249 | Rolled toilet paper | 0.459 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 250 | Liquid detergent, kitchen | 0.571 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 251 | Detergent, laundry | 0.604 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 252 | Food wrap | 0.502 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 253 | Insecticide | 0.209 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 254 | Moth repellent for clothes | 0.393 | 0.000 | 0.000 | 0.000 | 0.000 | reject |

| No. | Item | Frequency of price changes | Hjort (1990) test of goodness-of-fit | | | | Shapiro–Wilk (1972) test for exponentiality |
|---|---|---|---|---|---|---|---|
| | | | Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | |
| 255 | Fabric softener | 0.611 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 256 | Fragrance | 0.241 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 257 | Domestic help | 0.010 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 258 | Charges for treatment of human waste | 0.007 | 0.000 | 0.000 | 0.000 | 0.000 | * |
| 259 | Charges for mop-rental | 0.001 | - | - | - | - | - |
| | **[Clothes & Footwear]** | | | | | | |
| 260 | Women's "Kimono" | 0.092 | 0.000 | 0.002 | 0.000 | 0.000 | reject |
| 261 | Women's "Obi" | 0.078 | 0.000 | 0.002 | 0.000 | 0.000 | reject |
| 262 | Men's slacks (jeans) | 0.149 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 263 | Women's slacks (jeans) | 0.162 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 264 | Boys' short pants | 0.484 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 265 | Babies' clothes | 0.203 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 266 | Men's business shirts (long sleeves) | 0.215 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 267 | Men's undershirts (short sleeves) | 0.125 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 268 | Men's briefs | 0.120 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 269 | Brassieres | 0.178 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 270 | Panties | 0.076 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 271 | Slips | 0.175 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 272 | Children's undershirts | 0.217 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 273 | Men's shoes | 0.177 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 274 | Women's shoes | 0.225 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 275 | Children's shoes | 0.186 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 276 | Canvas shoes (for adults) | 0.144 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 277 | Canvas shoes (for children) | 0.201 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 278 | Sandals | 0.220 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 279 | "Zori", Japanese sandals | 0.139 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 280 | Women's dress materials | 0.159 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 281 | Men's suit materials | 0.079 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 282 | Woolen yarn | 0.101 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 283 | Hats & caps | 0.049 | 0.000 | 0.050* | 0.000 | - | reject |
| 284 | Neckties | 0.314 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 285 | Neckties (imported) | 0.250 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 286 | Women's stockings | 0.084 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 287 | Women's socks | 0.227 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 288 | Belts | 0.134 | 0.000 | 0.001 | 0.000 | 0.000 | reject |
| 289 | Tailoring charges | 0.041 | 0.000 | 0.005 | 0.000 | 0.000 | reject |
| 290 | Laundry charges (men's business shirts) | 0.050 | 0.000 | 0.174* | 0.000 | 0.000 | * |
| | **[Medical Care]** | | | | | | |
| 291 | Medicines for cold | 0.126 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 292 | Antipyretic & analgesic medicines | 0.087 | 0.000 | 0.001 | 0.000 | 0.000 | reject |
| 293 | Gastrointestinal medicines | 0.158 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 294 | Vitamins-A | 0.102 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 295 | Vitamins-B | 0.066 | 0.000 | 0.016* | 0.000 | 0.000 | reject |
| 296 | Health drinks | 0.060 | 0.000 | 0.006 | 0.000 | 0.000 | * |
| 297 | Dermal medicines | 0.065 | 0.000 | 0.046* | 0.000 | 0.000 | * |
| 298 | Plasters | 0.070 | 0.005 | 0.003 | 0.000 | 0.000 | * |
| 299 | Eyewashes | 0.184 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 300 | Breath fresheners | 0.051 | 0.000 | 0.183* | 0.000 | 0.000 | * |
| 301 | Chinese medicines | 0.109 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 302 | Disposable diapers | 0.607 | 0.007 | 0.000 | 0.032* | 0.000 | reject |

| No. | Item | Frequency of price changes | Hjort (1990) test of goodness-of-fit | | | | Shapiro–Wilk (1972) test for exponentiality |
|---|---|---|---|---|---|---|---|
| | | | Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | |
| 303 | Sanitary napkins | 0.550 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 304 | Bath preparation | 0.414 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 305 | Contact lenses cleaning solution | 0.112 | 0.000 | 0.002 | 0.000 | 0.000 | reject |
| 306 | Spectacles | 0.066 | 0.000 | 0.008 | 0.000 | 0.000 | reject |
| 307 | Contact lenses | 0.076 | 0.278∗ | 0.339∗ | 0.291∗ | 0.123∗ | ∗ |
| 308 | Bathroom scales | 0.262 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 309 | Thermometers | 0.058 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 310 | Sphygmomanometers | 0.293 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 311 | Medical treatment | 0.000 | - | - | - | - | - |
| 312 | Delivery fees in national hospital | 0.070 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 313 | Delivery fees in public hospital | 0.051 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 314 | Charges for massage | 0.016 | 0.000 | 0.039∗ | 0.000 | 0.000 | ∗ |
| 315 | Fees for complete medical checkup | 0.019 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| | | | **[Transportation & Communication]** | | | | |
| 316 | Railway fares (ordinary fares, excluding "Shinkansen") | 0.000 | - | - | - | - | - |
| 317 | Railway fares (special fares, excluding "Shinkansen") | 0.000 | - | - | - | - | - |
| 318 | Railway fares (students' season tickets) | 0.000 | - | - | - | - | - |
| 319 | Railway fares (commuters' season tickets) | 0.000 | - | - | - | - | - |
| 320 | Railway fares (ordinary passengers) | 0.001 | - | - | - | - | reject |
| 321 | Railway fares (students' season tickets) | 0.000 | - | - | - | - | - |
| 322 | Railway fares (commuters' season tickets) | 0.001 | - | - | - | - | - |
| 323 | Bus fares | 0.003 | - | - | - | - | ∗ |
| 324 | Taxi fares | 0.000 | - | - | - | - | - |
| 325 | Bicycles | 0.141 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 326 | Gasoline (regular) | 0.583 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 327 | Gasoline (premium) | 0.579 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 328 | Tires | 0.077 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 329 | Car wax | 0.106 | 0.000 | 0.002 | 0.000 | 0.000 | reject |
| 330 | Regular inspection | 0.021 | 0.000 | 0.009 | 0.001 | 0.006 | ∗ |
| 331 | Muffler replacement | 0.021 | 0.000 | 0.004 | 0.000 | 0.000 | ∗ |
| 332 | Puncture repairs | 0.019 | 0.006 | 0.046∗ | 0.008 | 0.046∗ | ∗ |
| 333 | Motor oil replacement | 0.040 | 0.013∗ | 0.235∗ | 0.000 | 0.000 | ∗ |
| 334 | Charges for garage rental | 0.026 | 0.391∗ | 0.699∗ | 0.068∗ | 0.043∗ | ∗ |
| 335 | Charges for parking | 0.014 | 0.000 | 0.107∗ | 0.001 | 0.005 | ∗ |
| 336 | Charges for driving license | 0.015 | 0.000 | - | - | - | reject |
| 337 | Telephone charges | 0.001 | - | - | - | - | - |
| 338 | Mobile telephone charges | 0.007 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 339 | Forwarding charges | 0.020 | 0.004 | 0.220∗ | 0.006 | 0.220∗ | ∗ |
| 340 | Communication equipments | 0.604 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| | | | **[Education]** | | | | |
| 341 | Tutorial fees | 0.062 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| | | | **[Reading & Recreation]** | | | | |
| 342 | TV sets | 0.568 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 343 | Stereo phonograph sets | 0.548 | 0.000 | 0.000 | 0.000 | 0.000 | reject |

| No. | Item | Frequency of price changes | Hjort (1990) test of goodness-of-fit | | | | Shapiro–Wilk (1972) test for exponentiality |
|---|---|---|---|---|---|---|---|
| | | | Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | |
| 344 | Mobile audio equipment | 0.596 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 345 | Video tape recorders | 0.544 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 346 | Cameras | 0.288 | 0.133∗ | 0.000 | 0.250∗ | 0.000 | reject |
| 347 | Ball-point pens | 0.036 | 0.057∗ | 0.334∗ | 0.052∗ | 0.072∗ | ∗ |
| 348 | Pencils | 0.061 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 349 | Marking pens | 0.023 | 0.000 | 0.083∗ | 0.000 | 0.000 | ∗ |
| 350 | Notebooks | 0.069 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 351 | Albums | 0.047 | 0.042∗ | 0.067∗ | 0.002 | 0.001 | ∗ |
| 352 | Papers for office automation | 0.092 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 353 | Cellophane adhesive tape | 0.034 | 0.000 | 0.050∗ | 0.000 | 0.000 | ∗ |
| 354 | Pencil cases | 0.110 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 355 | Golf clubs | 0.104 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 356 | Soccer balls | 0.136 | 0.000 | 0.003 | 0.000 | 0.000 | reject |
| 357 | Baseball gloves | 0.156 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 358 | Tennis rackets | 0.079 | 0.000 | 0.043∗ | 0.000 | 0.000 | reject |
| 359 | Tennis rackets (imported) | 0.186 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 360 | Fishing rods | 0.137 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 361 | Pants for exercise | 0.197 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 362 | Swimming suits | 0.182 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 363 | Video games, hardware | 0.320 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 364 | Dolls | 0.046 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 365 | Toy cars | 0.038 | 0.000 | 0.046∗ | 0.000 | 0.000 | ∗ |
| 366 | Building blocks | 0.122 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 367 | Cut flowers (Carnations) | 0.885 | 0.009 | 0.000 | 0.009 | 0.000 | reject |
| 368 | Cut flowers (Chrysanthemums) | 0.903 | 0.605† | - | - | - | reject |
| 369 | Cut flowers (Roses) | 0.899 | 0.006 | 0.000 | 0.006 | 0.000 | reject |
| 370 | Films | 0.102 | 0.014∗ | 0.001 | 0.003 | 0.000 | reject |
| 371 | Media for audio recording | 0.162 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 372 | Video tapes | 0.236 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 373 | Pet foods (dog foods) | 0.409 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 374 | Pet foods (cat foods) | 0.273 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 375 | Flowerpots | 0.074 | 0.000 | 0.014∗ | 0.000 | 0.000 | reject |
| 376 | Gardening earth | 0.175 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 377 | Dry batteries | 0.118 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 378 | Lesson fees (English conversation school) | 0.052 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 379 | Lesson fees (calligraphy school) | 0.018 | 0.000 | 0.045∗ | 0.000 | 0.001 | ∗ |
| 380 | Lesson fees (music school) | 0.023 | 0.000 | 0.001 | 0.000 | 0.000 | ∗ |
| 381 | Lesson fees (swimming school) | 0.018 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 382 | Lesson fees (dressmaking school) | 0.021 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 383 | Lesson fees (cooking school) | 0.030 | 0.012∗ | 0.057∗ | 0.001 | 0.001 | ∗ |
| 384 | Lesson fees, driving school | 0.062 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 385 | Admission, movies | 0.018 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 386 | Charges for practicing golf | 0.026 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 387 | Charges for playing golf | 0.255 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 388 | Tennis court charges | 0.081 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 389 | Game charges, bowling | 0.004 | 0.000 | 0.011∗ | 0.000 | 0.000 | ∗ |
| 390 | Swimming pool charges | 0.078 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 391 | Admission fees to the art museum | 0.250 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 392 | Game charges, mahjong | 0.017 | 0.000 | 0.189∗ | 0.000 | 0.000 | ∗ |
| 393 | “karaoke room” charges | 0.106 | 0.316∗ | 0.091∗ | 0.241∗ | 0.038∗ | ∗ |
| 394 | Photo processing charges | 0.051 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 395 | Charges for video rental | 0.099 | 0.000 | 0.001 | 0.000 | 0.000 | reject |
| 396 | Veterinary surgeon fees | 0.015 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| | **[Miscellaneous]** | | | | | | |
| 397 | Bathing charges (adults) | 0.003 | 0.000 | 0.001 | 0.000 | 0.000 | ∗ |

| No. | Item | Frequency of price changes | Hjort (1990) test of goodness-of-fit | | | | Shapiro–Wilk (1972) test for exponentiality |
|---|---|---|---|---|---|---|---|
| | | | Exponential | Weibull | Exponential-Gamma | Weibull-Gamma | |
| 398 | Bathing charges (children, 6–11 years) | 0.001 | - | - | - | - | ∗ |
| 399 | Bathing charges (children, under 6 years) | 0.000 | - | - | - | - | - |
| 400 | Men's haircut charges | 0.020 | 0.001 | 0.072∗ | 0.002 | 0.002 | ∗ |
| 401 | Permanent wave charges | 0.039 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 402 | Women's haircut charges | 0.026 | 0.007 | 0.318∗ | 0.008 | 0.088∗ | ∗ |
| 403 | Hair dyeing charges | 0.030 | 0.001 | 0.106∗ | 0.000 | 0.000 | ∗ |
| 404 | Electric shavers | 0.412 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 405 | Electric shavers (imported) | 0.358 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 406 | Toothbrushes | 0.376 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 407 | Toilet soap | 0.386 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 408 | Shampoo | 0.412 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 409 | Hair rinse | 0.342 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 410 | Toothpaste | 0.453 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 411 | Hair liquid | 0.170 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 412 | Hair tonic | 0.154 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 413 | Face cream-B | 0.070 | 0.000 | 0.016∗ | 0.000 | 0.000 | ∗ |
| 414 | Toilet lotion | 0.197 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 415 | Hair Dyeing | 0.347 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 416 | Handbags | 0.231 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 417 | Handbags (imported) | 0.223 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 418 | Suitcases | 0.183 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 419 | Rings | 0.077 | 0.000 | 0.209∗ | 0.000 | 0.000 | reject |
| 420 | Wrist watches | 0.070 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 421 | Wrist watches (imported) | 0.113 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 422 | Repair charges of wrist watches | 0.013 | 0.000 | 0.040∗ | 0.000 | 0.000 | ∗ |
| 423 | Men's umbrellas | 0.181 | 0.000 | 0.000 | 0.000 | 0.000 | reject |
| 424 | Handkerchiefs | 0.041 | 0.000 | 0.001 | 0.000 | 0.000 | reject |
| 425 | Nursery school fees | 0.006 | 0.000 | 0.000 | 0.000 | 0.000 | ∗ |
| 426 | Charges for certificates of registered stamps | 0.000 | - | - | - | - | - |
| 427 | Charges for certificates of permanent registration | 0.000 | - | - | - | - | - |
| 428 | Charges for acquisition of passport | 0.000 | - | - | - | - | - |
| 429 | Day service fees of nursing care for the aged | 0.019 | 0.000 | 0.000 | 0.000 | 0.000 | reject |

Table 4: Goodness-of-fit test results by item. Entries are the 429 items in the Japanese CPI for 2000–2005. P–values are reported for the Hjort (1990) test of goodness-of-fit. †, the Hjort (1990) test cannot reject the hypothesis due to the high price flexibility (see also Footnote 19 in the text); ∗, the item cannot reject the hypothesis that the underlying distribution belongs to a specified family of distributions. "reject" in the last column means that the Shapiro and Wilk (1972) test rejects the null hypothesis of exponentiality at the one percent significance level.