

Some Recent Developments in Nonparametric Finance

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Preliminary version: comments welcome

October 26, 2008

Abstract

This paper gives a selective review on the recent developments of nonparametric methods in both continuous and discrete time finance, particularly in the areas of nonparametric estimation and testing of diffusion processes, nonparametric testing of parametric diffusion models, nonparametric pricing of derivatives, nonparametric estimation and hypothesis testing for nonlinear pricing kernel, and nonparametric predictability of asset returns. For each financial context, the paper discusses the suitable statistical concepts, models, and modeling procedures, as well as some of their applications to financial data. Their relative strengths and weakness are discussed. Much theoretical and empirical research is needed in this area, and more importantly, the paper points to several aspects that deserve further investigation.

Keywords: Continuous time model; derivative pricing; jump process; kernel smoothing; nonparametric test; nonparametric pricing kernel; non-stationarity; options; predictability; stochastic discount factor; time-dependent model.

The authors thank Federico M. Bandi and Haitao Li for their valuable and helpful comments, suggestions and discussions. Also, the authors thank participants at the seminars at Graduate School of Business, University of Chicago and Department of Economics at Columbia University and NYU as well as Institute of Economics at Academia Sinica, Taiwan, and audiences at the 7th Annual Advances in Econometrics Conference (November 2008 at Louisiana State University) for their helpful comments. Cai's research was supported, in part, by the National Science Foundation grant DMS-0404954 and funds provided by the University of North Carolina at Charlotte.

1 Introduction

Nonparametric methods have become a core area in statistics and econometrics in the last two decades; see the books by Härdle (1990), Fan and Gijbels (1996) and Li and Racine (2007) for general statistical methodology and theory as well as applications. They have been used successfully to various fields such as economics and finance due to their advantage of requiring little prior information on the data generating process; see the books by Pagan and Ullah (1999), Mittelhammer, Judge and Miller (2000), Tsay (2005), Taylor (2005) and Li and Racine (2007) for real examples in economics and finance. Recently, nonparametric techniques have been proved to be the most attractive way to conduct research and gain economic intuition in certain core areas in finance, such as asset and derivative pricing, term structure theory, portfolio choice, risk management, and predictability of asset returns, particularly, in modeling both continuous and discrete financial time series models; see the books by Campbell, Lo and MacKinlay (1997), Gouriéroux and Jasiak (2001), Duffie (2001), Tsay (2005) and Taylor (2005).

Finance is characterized by time and uncertainty. Both continuous and discrete time series modelling has been a basic analytic tool in modern finance since the seminar papers by Sharpe (1964), Fama (1970), Black and Scholes (1973) and Merton (1973). The rationale behind it is that most of time, news arrives at financial markets in both continuous and discrete manner. More importantly, derivative pricing in theoretical finance is generally much more convenient and elegant in a continuous time framework than through binomial or other discrete approximations. However, statistical analysis based on continuous time financial models has just emerged as a field in less than a decade, although it has been used for more than 4 decades for discrete financial time series. This is apparently due to the difficulty of estimating and testing continuous time models using discretely observed data. The purpose of this survey is to review some recent developments of nonparametric methods in both continuous and discrete time finance for recent years, and particularly in the areas of nonparametric estimation and testing of diffusion models, nonparametric derivative pricing and its tests, and predictability of asset returns based on nonparametric approaches. Financial time series data have some distinct important stylized facts, such as persistent volatility clustering, heavy tails, strong serial dependence, and occasionally sudden but large jumps. In addition, financial modelling is often closely embedded in a financial theoretical framework. These features suggest that standard statistical theory may not be readily applicable to both continuous and discrete financial time series. This is a promising and fruitful area for both financial economists and statisticians to interact each other.

Section 2 introduces various continuous-time diffusion processes and nonparametric estimation methods for diffusion processes. Section 3 reviews the estimation and testing of a parametric diffusion model using nonparametric methods. Section 4 discusses nonparametric estimation and hypothesis testing of derivative and asset pricing, particularly the nonparametric estimation of risk neutral density functions and nonlinear pricing kernel. Nonparametric predictability of asset returns is presented in Section 5. Section 6 concludes.

2 Diffusions and Nonparametric Estimation

2.1 Models

Modeling the dynamics of interest rates, stock prices, foreign exchange rates, and macroeconomic factors, *inter alia*, is one of the most important topics in asset pricing studies. The instantaneous risk-free interest rate or the so-called short rate is, for example, the state variable that determines the evolution of the yield curve in an important class of term structure models, such as Vasicek (1977) and Cox, Ingersoll and Ross (1985, CIR). It is of fundamental importance for pricing fixed-income securities. Many theoretical models have been developed in mathematical finance to describe the short rate movement.¹

In the theoretical term structure literature, the short rate or the underlying process of interest, $\{X_t, t \geq 0\}$, is often modelled as a time-homogeneous diffusion process, or stochastic differential equation,

$$dX_t = \mu(X_t) dt + \sigma(X_t) dB_t, \quad (2.1)$$

where $\{B_t, t \geq 0\}$ is a standard Brownian motion. The functions $\mu(\cdot)$ and $\sigma^2(\cdot)$ are respectively the drift (or instantaneous mean) and the diffusion (or instantaneous variance) of the process, which determine the dynamics of the short rate. Indeed, the model (2.1) can be applied to many core areas in finance, such as options, derivative pricing, asset pricing, term structure of interest rates, dynamic consumption and portfolio choice, default risk, stochastic volatility, and exchange rate dynamics, and others.

There are two basic approaches to identifying $\mu(\cdot)$ and $\sigma(\cdot)$. The first is a parametric approach, which assumes some parametric forms $\mu(\cdot, \theta)$ and $\sigma(\cdot, \theta)$, and estimates the unknown model parameters, say θ . Most existing models in the literature assume that the interest rate exhibits mean-reversion and that the drift $\mu(\cdot)$ is a linear or quadratic function of the interest rate level. It is also often assumed that the diffusion $\sigma(\cdot)$ takes the form of $\sigma |X_t|^\gamma$,

¹Other theoretical models are studied by Brennan and Schwartz (1979), Constantinides (1992), Courtadon (1982), Cox, Ingersoll and Ross (1980), Dothan (1978), Duffie and Kan (1996), Longstaff and Schwartz (1992), Marsh and Rosenfield (1983), and Merton (1973). Heath, Jarrow and Morton (1992) consider another important class of term structure models which use the forward rate as the underlying state variable.

where γ measures the sensitivity of interest rate volatility to the interest rate level. This specification, in modelling interest rate dynamics, captures the so-called “level effect”; i.e., the higher the interest rate level, the larger the volatility. With $\gamma = 0$ and 0.5 , the model (2.1) reduces to the well-known Vasicek and CIR models, respectively. The forms of $\mu(\cdot, \theta)$ and $\sigma(\cdot, \theta)$ are typically chosen due to theoretical wisdom or convenience. They may not be consistent with the data generating process and there might be a risk of misspecification.

The second approach is a nonparametric one, which does not assume any restrictive functional form for $\mu(\cdot)$ and $\sigma(\cdot)$ beyond regularity conditions. In the last few years, great progress has been made in estimating and testing continuous-time models for the short term interest rate using nonparametric methods.² Despite many studies, empirical analysis on the functional forms of the drift and diffusion is still not conclusive. For example, recent studies by Ait-Sahalia (1996b) and Stanton (1997) using nonparametric methods, overwhelmingly reject all linear drift models for the short rate. They find that the drift of the short rate is a nonlinear function of the interest rate level. Both studies show that for the lower and middle ranges of the interest rate, the drift is almost zero, i.e., the interest rate behaves like a random walk. But the short rate exhibits strong mean-reversion when the interest rate level is high. These findings lead to the development of nonlinear term structure models such as those of Ahn and Gao (1999).

However, the evidence of nonlinear drift has been challenged by Pritsker (1998) and Chapman and Pearson (2000), who find that the nonparametric methods of Ait-Sahalia (1996b) and Stanton (1997) have severe finite sample problems, especially near the extreme observations. The finite sample problems with nonparametric methods cast doubt on the evidence of nonlinear drift. On the other hand, the findings in Ait-Sahalia (1996b) and Stanton (1997) that the drift is nearly flat for the middle range of the interest rate are not much affected by the small sample bias. The reason is that near the extreme observations, the nonparametric estimation might not accurate due to the sparsity of data in this region. Also, this region is close to the boundary point, so that the Nadaraya-Watson (NW) estimate suffers a boundary effect. Chapman and Pearson (2000) point out that this is a puzzling fact, since “there are strong theoretical reasons to believe that short rate cannot exhibit the asymptotically explosive behavior implied by a random walk model.” They conclude that “time series methods alone are not capable of producing evidence of nonlinearity in the drift.” Recently, to overcome the boundary effect, Fan and Zhang (2003) fit a nonparametric model

²Empirical studies on the short rate include Ait-Sahalia (1996a, b), Andersen and Lund (1997), Ang and Bekaert (2002a, b), Brenner, Harjes and Kroner (1996), Brown and Dybvig (1986), Chan, Karolyi, Longstaff and Sanders (1992), Chapman and Pearson (2000), Chapman, Long and Pearson (1999), Conley, Hansen, Luttmmer and Scheinkman (1997), Gray (1996), and Stanton (1997).

using a local linear technique and apply the generalized likelihood ratio test of Cai, Fan and Yao (2000) and Fan, Zhang and Zhang (2001) to test whether the drift is linear. They support Chapman and Pearson's (2000) conclusion. However, the generalized likelihood ratio test is developed by Cai, Fan and Yao (2000) for discrete time series and Fan, Zhang and Zhang (2001) for iid samples but it is still unknown whether it is valid for continuous time series contexts, which is warranted for a further investigation. Interest rate data are well-known for persistent serial dependence. Pritsker (1998) uses Vasicek's (1977) model of interest rates to investigate the performance of a nonparametric density estimation in finite samples. He finds that asymptotic theory gives poor approximation even for a rather large sample size.

Controversies also exist on the diffusion $\sigma(\cdot)$. The specification of $\sigma(\cdot)$ is important, because it affects derivative pricing. Chan, Karolyi, Longstaff and Sanders (1992) show that in a single factor model of the short rate, γ roughly equals to 1.5 and all the models with $\gamma \leq 1$ are rejected. Ait-Sahalia (1996b) finds that γ is close to 1, Stanton (1997) finds that in his semiparametric model γ is about 1.5, and Conley, Hansen, Luttmer and Scheikman (1997) show that their estimate of γ is between 1.5 and 2. However, Bliss and Smith (1998) argue that the result that γ equals to 1.5 depends on whether the data between October 1979 to September 1982 are included. From the above discussions, it seems that the value of γ might change over the time.

2.2 Nonparametric Estimation

Under some regularity conditions; see Jiang and Knight (1997) and Bandi and Nguyen (2000), the diffusion process in (2.1) is a one dimensional, regular, strong Markov process with continuous sample paths and time invariant stationary transition density. The drift and diffusion are respectively the first two moments of the infinitesimal conditional distribution of X_t :

$$\mu(X_t) = \lim_{\Delta \rightarrow 0} \Delta^{-1} E[Y_t | X_t], \quad \text{and} \quad \sigma^2(X_t) = \lim_{\Delta \rightarrow 0} \Delta^{-1} E[Y_t^2 | X_t], \quad (2.2)$$

where $Y_t = X_{t+\Delta} - X_t$. See, e.g., Øksendal (1985) and Karatzas and Shreve (1988). The drift describes the movement of X_t due to time changes, whereas the diffusion term measures the magnitude of random fluctuations around the drift.

Using the Dynkin (infinitesimal) operator; see, e.g., Øksendal (1985) and Karatzas and Shreve (1988), Stanton (1997) shows that the first order approximation

$$\mu(X_t)^{(1)} = \frac{1}{\Delta} E\{X_{t+\Delta} - X_t | X_t\} + O(\Delta),$$

the second order approximation

$$\mu(X_t)^{(2)} = \frac{1}{2\Delta} [4 E \{Y_t | X_t\} - E \{X_{t+2\Delta} - X_t | X_t\}] + O(\Delta^2),$$

and the third order approximation

$$\mu(X_t)^{(3)} = \frac{1}{6\Delta} [18 E \{Y_t | X_t\} - 9 E \{X_{t+2\Delta} - X_t | X_t\} + 2 E \{X_{t+3\Delta} - X_t | X_t\}] + O(\Delta^3),$$

etc. Fan and Zhang (2003) derive higher-order approximations. Similar formulas hold for the diffusion; see Stanton (1997). Bandi and Nguyen (2000) argue that approximations to the drift and diffusion of any order display the same rate of convergence and limiting variance, so that asymptotic argument in conjunction with computational issues suggest simply using the first order approximations in practice. As indicated by Stanton (1997), the higher the order of the approximations, the faster they will converge to the true drift and diffusion. However, as noted by Bandi and Nguyen (2000) and Fan and Zhang (2003), higher order approximations can be detrimental to the efficiency of the estimation procedure in finite samples. In fact, the variance grows nearly exponentially fast as the order increases and they are much more volatile than their lower order counterparts. For more discussions, see Bandi (2000), Bandi and Nguyen (2000) and Fan and Zhang (2003).

Now suppose we observe X_t at $t = \tau\Delta$, $\tau = 1, \dots, n$, in a fixed time interval $[0, T]$ with T . Denote the random sample as $\{X_{\tau\Delta}\}_{\tau=1}^n$. Then, it follows from (2.2) that the first order approximations to $\mu(\cdot)$ and $\sigma(\cdot)$ lead to

$$\mu(X_{\tau\Delta}) \approx \frac{1}{\Delta} E[Y_\tau | X_{\tau\Delta}] \quad \text{and} \quad \sigma^2(X_{\tau\Delta}) \approx \frac{1}{\Delta} E[Y_\tau^2 | X_{\tau\Delta}]$$

for all $1 \leq \tau \leq n - 1$, where $Y_\tau = X_{(\tau+1)\Delta} - X_{\tau\Delta}$. Both $\mu(X_{\tau\Delta})$ and $\sigma^2(X_{\tau\Delta})$ become classical nonparametric regressions and a nonparametric kernel smoothing approach can be applied to estimate them.

There are many nonparametric approaches to estimating conditional expectations. Most existing nonparametric methods in finance dwell mainly on the Nadaraya-Watson kernel estimator due to its simplicity. According to Ait-Sahalia (1996a, b), Stanton (1997), Jiang and Knight (1997) and Chapman and Pearson (2000), the NW estimators of $\mu(x)$ and $\sigma^2(x)$ are given, respectively, by

$$\hat{\mu}(x) = \frac{1}{\Delta} \frac{\sum_{\tau=1}^{n-1} Y_\tau K_h(x - X_{\tau\Delta})}{\sum_{\tau=1}^{n-1} K_h(x - X_{\tau\Delta})}, \quad \text{and} \quad \hat{\sigma}^2(x) = \frac{1}{\Delta} \frac{\sum_{\tau=1}^{n-1} Y_\tau^2 K_h(x - X_{\tau\Delta})}{\sum_{\tau=1}^{n-1} K_h(x - X_{\tau\Delta})}, \quad (2.3)$$

where $K_h(u) = K(u/h)/h$, $h = h_n > 0$ is the bandwidth with $h \rightarrow 0$ and $nh \rightarrow \infty$ as $n \rightarrow \infty$, and $K(\cdot) : \mathbb{R} \rightarrow \mathbb{R}$ is a standard kernel. Jiang and Knight (1997) suggest first using

(2.3) to estimate $\sigma^2(\cdot)$. Observe that the drift

$$\mu(X_t) = \frac{1}{2\pi(X_t)} \frac{\partial[\sigma^2(X_t)\pi(X_t)]}{\partial X_t},$$

where $\pi(\cdot)$ is the stationary density of $\{X_t\}$; see, e.g., Ait-Sahalia (1996a), Jiang and Knight (1997), Stanton (1997) and Bandi and Nguyen (2000). Therefore, Jiang and Knight (1997) suggest estimating $\mu(\cdot)$ by

$$\hat{\mu}(x) = \frac{1}{2\hat{\pi}(x)} \frac{\partial\{\hat{\sigma}^2(x)\hat{\pi}(x)\}}{\partial x},$$

where $\hat{\pi}(\cdot)$ is a consistent estimator of $\pi(\cdot)$, say, the classical kernel density estimator. The reason of doing so is based on the fact that in (2.1) the drift is of order dt and the diffusion is of order \sqrt{dt} , as $(dB_t)^2 = dt + O((dt)^2)$. That is, the diffusion has lower order than the drift for infinitesimal changes in time, and the local-time dynamics of the sampling path reflects more of the diffusion than those of the drift term. Therefore, when Δ is very small, identification becomes much easier for the diffusion term than the drift term.

It is well known that the NW estimator suffers from some disadvantages such as larger bias, boundary effects, and inferior minimax efficiency; see, e.g., Fan and Gijbels (1996). To overcome these drawbacks, Fan and Zhang (2003) suggest using the local linear technique, for $k = 1$ and 2,

$$\sum_{\tau=1}^{n-1} \{\Delta^{-1}Y_{\tau}^k - \beta_0 - \beta_1(x - X_{\tau\Delta})\}^2 K_h(x - X_{\tau\Delta}), \quad (2.4)$$

which gives the local linear estimate of $\mu(\cdot)$ for $k = 1$ and $\sigma^2(\cdot)$ for $k = 2$. However, the local linear estimator of the diffusion $\sigma(\cdot)$ cannot be always nonnegative in finite samples. To attenuate this disadvantage of local polynomial method, as suggested by Cai and Hong (2003), a weighted NW method due to Cai (2001) can be used to estimate $\sigma(\cdot)$ but they did not investigate further. Recently, Xu and Phillips (2007) study this approach and its properties.

The asymptotic theory can be found in Jiang and Knight (1997) and Bandi and Nguyen (2000) for the NW estimator and in Fan and Zhang (2003) for the local linear estimator as well as Xu and Phillips (2007) for the weighted NW estimator. To implement kernel estimates, the bandwidth(s) must be chosen. In the iid setting, there are theoretically optimal bandwidth selections. There are no such results for diffusion processes available although there are many theoretic and empirical studies in the literature. As a rule of

thumb, an easy way to choose a data-driven fashion bandwidth is to use the nonparametric version of the Akaike information criterion; see Cai and Tiwari (2000).

One crucial assumption in the above development is the stationarity of $\{X_t\}$. However, it might not hold for real financial time series data. If $\{X_t\}$ is not stationary, Bandi and Phillips (2003) propose using the following estimators to estimate $\mu(x)$ and $\sigma^2(x)$, respectively,

$$\hat{\mu}(x) \frac{\sum_{\tau=1}^n K_h(x - X_{\tau\Delta}) \tilde{\mu}(X_{\tau\Delta})}{\sum_{\tau=1}^n K_h(x - X_{\tau\Delta})}, \quad \text{and} \quad \hat{\sigma}^2(x) \frac{\sum_{\tau=1}^n K_h(x - X_{\tau\Delta}) \tilde{\sigma}^2(X_{\tau\Delta})}{\sum_{\tau=1}^n K_h(x - X_{\tau\Delta})},$$

where

$$\tilde{\mu}(x) = \frac{1}{\Delta} \frac{\sum_{\tau=1}^{n-1} I(|X_{\tau\Delta} - x| \leq b) Y_{\tau}}{\sum_{\tau=1}^n I(|X_{\tau\Delta} - x| \leq b)}, \quad \text{and} \quad \tilde{\sigma}^2(x) = \frac{1}{\Delta} \frac{\sum_{\tau=1}^{n-1} I(|X_{\tau\Delta} - x| \leq b) Y_{\tau}^2}{\sum_{\tau=1}^n I(|X_{\tau\Delta} - x| \leq b)}.$$

See, also Bandi and Nguyen (2000). Here, $b = b_n > 0$ is a bandwidth-like smoothing parameter that depends on the time span and on the sample size, which is called the spatial bandwidth in Bandi and Phillips (2003). This modeling approach is termed as the *chronological local time* estimation. Bandi and Philips's approach can deal well with the situation that the series is not stationary. The reader is referred to the papers by Bandi and Phillips (2003) and Bandi and Nguyen (2000) for more discussions and asymptotic theory.

Bandi and Philips's (2003) estimator can be viewed as a double kernel smoothing method: The first step defines straight sample analogs to the values that drift and diffusion take at the sampled points. Indeed, this step uses the smoothing technique (a linear estimator with same weights) to obtain the raw estimates of the two functions $\tilde{\mu}(x)$ and $\tilde{\sigma}^2(x)$, respectively. This approach is different from classical two-step method in the literature; see Cai (2002a, 2002b). The key is to figure out how important the first is to the second step. To implement this estimator, an empirical and theoretical study on the selection of two bandwidths b and h is needed.

2.3 Time-Dependent Diffusion Models

The time-homogeneous diffusion models in (2.1) have certain limitations. For example, they cannot capture the time effect, as addressed at the end of Section 2.1. A variety of time-dependent diffusion models have been proposed in the literature. A time-dependent diffusion process is formulated as

$$dX_t = \mu(X_t, t) dt + \sigma(X_t, t) dB_t. \quad (2.5)$$

Examples of (2.5) include Ho and Lee (HL) (1986), Hull and White (HW) (1990), Black, Derman and Toy (BDT) (1990), and Black and Karasinski (BK) (1991), among others. They

consider respectively the following models:

$$\begin{aligned}
\text{HL:} \quad dX_t &= \mu(t) dt + \sigma(t) dB_t, \\
\text{HW:} \quad dX_t &= [\alpha_0 + \alpha_1(t) X_t] dt + \sigma(t) X_t^\gamma dB_t, \quad \gamma = 0 \text{ or } 0.5 \\
\text{BDT:} \quad dX_t &= [\alpha_1(t) X_t + \alpha_2(t) X_t \log(X_t)] dt + \sigma(t) X_t dB_t, \\
\text{BK:} \quad dX_t &= [\alpha_1(t) X_t + \alpha_2(t) X_t \log(X_t)] dt + \sigma(t) X_t dB_t,
\end{aligned}$$

where $\alpha_2(t) = \sigma'(t)/\sigma(t)$. Similar to (2.2), one has

$$\mu(X_t, t) = \lim_{\Delta \rightarrow 0} \Delta^{-1} E\{Y_t | X_t\}, \quad \text{and} \quad \sigma^2(X_t, t) = \lim_{\Delta \rightarrow 0} \Delta^{-1} E\{Y_t^2 | X_t\},$$

where $Y_t = X_{t+\Delta} - X_t$, which provide a regression form for estimating $\mu(\cdot, t)$ and $\sigma^2(\cdot, t)$.

By assuming that the drift and diffusion functions are linear in X_t with time varying coefficients, Fan, Jiang, Zhang and Zhou (2003) consider the following time-varying coefficients single factor model

$$dX_t = [\alpha_0(t) + \alpha_1(t) X_t] dt + \beta_0(t) X_t^{\beta_1(t)} dB_t, \quad (2.6)$$

and use the local linear technique in (2.4) to estimate the coefficient functions $\{\alpha_j(\cdot)\}$ and $\{\beta_j(\cdot)\}$. Since the coefficients depend on time, $\{X_t\}$ might not be stationary. The asymptotic properties of the resulting estimators are still unknown. Indeed, the aforementioned models are a special case of the following more general time-varying coefficient multi-factor diffusion models

$$dX_t = \mu(X_t, t) dt + \sigma(X_t, t) dB_t, \quad (2.7)$$

where

$$\mu(X_t, t) = \alpha_0(t) + \alpha_1(t) g(X_t), \quad \text{and} \quad (\sigma(X_t, t)\sigma(X_t, t)^\top)_{ij} = \beta_{0,ij}(t) + \beta_{1,ij}(t)^\top h_{ij}(X_t),$$

and $g(\cdot)$ and $\{h_{ij}(\cdot)\}$ are known functions. This is the time-dependent version of the multi-factor affine models studied in Duffie, Pan and Singleton (2000). It allows time-varying coefficients in multi-factor affine models. A further theoretical and empirical study of the time-varying coefficient multi-factor diffusion model in (2.7) is warranted. It is interesting to point out that the estimation approaches described above are still applicable to the model (2.7) but the asymptotic theory is very challenging due to the nonstationarity of unknown structure of the underlying process $\{X_t\}$.

2.4 Jump Diffusion Models

There has been a vast literature on the study of diffusion models with jumps.³ The main purpose of adding jumps into diffusion models or stochastic volatility diffusion models is to accommodate impact of sudden and large shocks to financial markets, such as macroeconomic announcements, the Asian and Russian finance crisis, the US finance turmoil, an unusually large unemployment announcement, and a dramatic interest rate cut by the Federal Reserve. For more discussions on why it is necessary to add jumps into diffusion models, see, for example, Lobo (1999), Bollerslev and Zhou (2002), Liu, Longstaff and Pan (2002), and Johannes (2004), among others. Also, Jumps can capture the heavy tail behavior of the distribution of the underlying process; see later.

For the expositional purpose, we only consider a single factor diffusion model with jump:

$$dX_t = \mu(X_t) dt + \sigma(X_t) dB_t + dJ_t, \quad (2.8)$$

where J_t is a compensated jump process (zero conditional mean) with arrival rate $\lambda_t = \lambda(X_t) \geq 0$, which is an instantaneous intensity function. There are several studies on specification of J_t . For example, $J_t = \xi P_t$, where P_t is a Poisson process with an intensity $\lambda(X_t)$ or a binomial distribution with probability $\lambda(X_t)$, and the jump size, ξ , has a time-invariant distribution $\Pi(\cdot)$ with mean zero. $\Pi(\cdot)$ can be either normal or uniform. If $\lambda_t(\cdot) = 0$ or $E(\xi^2) = 0$, the jump-diffusion model in (2.8) becomes the diffusion model in (2.1). More generally, Chernov, Gallant, Ghysels and Tauchen (2003) consider a Lévy process for $\{J_t\}$. A simple jump diffusion model proposed by Kou (2002) is discussed in Tsay (2005) by assuming that $J_t = \sum_{i=1}^{n_t} (L_i - 1)$, where n_t is a Poisson process with rate λ , and $\{L_i\}$ is a sequence of independent and identically distributed nonnegative random variables such that $\ln(L_i)$ has a double exponential distribution with probability density function. This simple model enjoys several nice properties. The returns implied by the model are leptokurtic and asymmetric with respect to zero. In addition, the model can reproduce volatility smile and provide analytical formulas for the prices of many options.

In practice, $\lambda(\cdot)$ might be assumed to have a particular form. For example, Chernov, Gallant, Ghysels and Tauchen (2003) consider three different types of special forms, each having the appealing feature of yielding analytic option pricing formula for European type contracts written on the stock price index. There are some open issues for the jump-diffusion model: (i) jumps are not observed and it is not possible to say surely if they exist; (ii) if

³See, to name just a few, Pan (1997), Duffie and Pan (2001), Bollerslev and Zhou (2002), Eraker, Johannes and Polson (2003), Bates (2000), Duffie, Pan and Singleton (2000), Johannes (2004), Liu, Longstaff and Pan (2002), Zhou (2001), Singleton (2001), Perron (2001), Chernov, Gallant, Ghysels and Tauchen (2003).

they exist, a natural question arises how to estimate a jump time τ , which is defined to be the discontinuous time at which $X_{\tau+} \neq X_{\tau-}$, and the jump size ξ , which is $\xi = X_{\tau+} - X_{\tau-}$. Wavelet methods may be potentially useful here.

Similar to (2.2), the first two conditional moments are given by

$$\mu_1(X_t) = \lim_{\Delta \downarrow 0} \Delta^{-1} E[Y_t | X_t] = \mu(X_t) + \lambda(X_t) E(\xi),$$

and

$$\mu_2(X_t) = \lim_{\Delta \downarrow 0} \Delta^{-1} E[Y_t^2 | X_t] = \sigma^2(X_t) + \lambda(X_t) E(\xi^2).$$

Clearly, $\mu_2(X_t)$ is much bigger than $\sigma^2(X_t)$ if there is a jump. This means that adding a jump into the model can capture the heavy tails. Also, it is easy to see that the first two moments are the same as those for a diffusion model by using a new drift coefficient $\tilde{\mu}(X_t) = \mu(X_t) + \lambda(X_t) E(\xi)$ and a new diffusion coefficient $\tilde{\sigma}^2(x) = \sigma^2(x) + \lambda(x) E(\xi^2)$. However, the fundamental difference between a diffusion model and a diffusion model with jumps relies on higher order moments. Using the infinitesimal generator (Øksendal, 1985; Karatzas and Shreve, 1988) of X_t , we can compute, $j > 2$,

$$\mu_j(X_t) = \lim_{\Delta \rightarrow 0} \Delta^{-1} E[Y_t^j | X_t] = \lambda(X_t) E(\xi^j).$$

See Duffie, Pan and Singleton (2000) and Johannes (2004) for details. Obviously, jumps provide a simple and intuitive mechanism for capturing the heavy tail behavior of interest rates. In particular, the conditional skewness and kurtosis are respectively given by

$$s(X_t) \equiv \frac{\lambda(X_t) E(\xi^3)}{[\sigma^2(X_t) + \lambda(X_t) E(\xi^2)]^{3/2}}, \quad \text{and} \quad k(X_t) \equiv \frac{\lambda(X_t) E(\xi^4)}{[\sigma^2(X_t) + \lambda(X_t) E(\xi^2)]^2}.$$

Note that $s(X_t) = 0$ if ξ is symmetric. By assuming $\xi \sim N(0, \sigma_\xi^2)$, Johannes (2004) uses the conditional kurtosis to measure the departures for the treasury bill data from normality and concludes that interest rates exchanges are extremely non-normal.

The NW estimation of $\mu_j(\cdot)$ is considered by Johannes (2004) and Bandi and Nguyen (2003). Moreover, Bandi and Nguyen (2003) provide a general asymptotic theory for the resulting estimators. Further, by specifying a particular form of $\Pi(\lambda) = \Pi_0(\lambda, \theta)$, say, $\xi \sim N(0, \sigma_\xi^2)$, Bandi and Nguyen (2003) propose consistent estimators of $\lambda(\cdot)$, σ_ξ^2 , and $\sigma^2(\cdot)$ and derive their asymptotic properties.

A natural question arises how to measure the departures from a pure diffusion model statistically. That is to test the model (2.8) against the model (2.1). It is equivalent to

checking whether $\lambda(\cdot) \equiv 0$ or $\xi = 0$. Instead of using the conditional skewness or kurtosis, a test statistic can be constructed based on the higher order conditional moments. For example, one can construct the following nonparametric test statistics

$$T_1 = \int \hat{\mu}_4(x) w(x) dx, \quad \text{or} \quad T_2 = \int \hat{\mu}_3^2(x) w(x) dx, \quad (2.9)$$

where $w(\cdot)$ is a weighting function. It needs a further investigation theoretically and empirically.

It is well known that prices fully reflect the available information in the efficient market. Thus, we consider the market information consist of two components. First, the anticipated information that drive market prices' daily normal fluctuation. Second, the unanticipated information that drive prices to exceptional fluctuation, which can be characterized by a jump process. Therefore, Cai and Zhang (2008b) investigate the market information via a jump-diffusion process. The jump term in the dynamic of stock price or return rate reflects the sensitivity of unanticipated information for the related firms. This implies the investigation of the jump parameters for firms with different sizes would help us to find the relationship between firm sizes and information sensitivity. With the nonparametric method as described above, Cai and Zhang (2008b) use the kernel estimation method, and reveal how the nonparametric estimation of the jump parameters reflect the so called information effect. Also, they test the model based on the test statistic formulated in (2.9). Due to the lack of the relevant theory of the test statistics in (2.9), Cai and Zhang (2008b) use the Monte Carlo simulation, and find that a jump diffusion process performs better to model all market information, including anticipated and unanticipated information than the pure diffusion model. Empirically, Cai and Zhang (2008b) estimate the jump intensity and jump variance for portfolios with different firm sizes for data from both the US and Chinese markets, and find some evidences that there exists information effect among different firm sizes, from which we could get valuable references for investors' decision making. Finally, using a Monte Carlo simulation method, Cai and Zhang (2008a) exam the test statistics in (2.9) to see how the discontinuity of drift or diffusion function affects the performance of the test statistics. They find that the discontinuity of drift or diffusion function has an impact on the performance of the test statistics.

More generally, for given a discrete sample of a diffusion process, can one tell whether the underlying model that gave rise to the data was a diffusion, or should jumps be allowed into the model? To answer this question, Ait-Sahalia (2002b) proposes an approach to identifying the sufficient and necessary restriction on the transition densities of diffusions, at the sampling interval of the observed data. This restriction characterizes the continuity

of the unobservable continuous sample path of the underlying process and is valid for every sampling interval including long ones. Let $\{X_t, t \geq 0\}$ be a Markovian process taking values in $D \subseteq \mathbb{R}$. Let $p(\Delta, y | x)$ denote the transition density function of the process over interval length Δ , that is, the conditional density of $X_{t+\Delta} = y$ given $X_t = x$, and it is assumed that the transition densities are time homogenous. Ait-Sahalia (2002b) shows that if the transition density $p(\Delta, y | x)$ is strictly positive and twice-continuously differentiable on $D \times D$ and the following condition

$$\frac{\partial^2}{\partial x \partial y} \ln p(\Delta, y | x) > 0 \quad \text{for all } \Delta > 0 \quad \text{and} \quad (x, y) \in D \times D,$$

which is the so called ‘‘diffusion criterion’’ in Ait-Sahalia (2002b), is satisfied, then, the underlying process is a diffusion. From a discretely sampled time series $\{X_{\tau\Delta}\}$, one could test nonparametrically the hypothesis that the data were generated by a continuous-time diffusion $\{X_t\}$. That is to test nonparametrically the null hypothesis

$$\mathbb{H}_0 : \frac{\partial^2}{\partial x \partial y} \ln p(\Delta, y | x) > 0 \quad \text{for all } x, y$$

versus the alternative

$$\mathbb{H}_a : \frac{\partial^2}{\partial x \partial y} \ln p(\Delta, y | x) \leq 0 \quad \text{for some } x, y.$$

One could construct a test statistic based on checking whether the above ‘‘diffusion criterion’’ holds for a nonparametric estimator of $p(\Delta, y | x)$. This topic is still open. If the model has a specific form, say a parametric form, the diffusion criterion becomes a simple form, say, it becomes just a constraint for some parameters. Then, the testing problem becomes to testing a constraint on parameters; see Ait-Sahalia (2002b) for some real applications.

2.5 Time-Dependent Jump Diffusion Models

Duffie, Pan and Singleton (2000) consider the time-varying coefficient intensity

$$\lambda(X_t, t) = \lambda_0(t) + \lambda_1(t) X_t,$$

and Chernov, Gallant, Ghysels and Tauchen (2003) consider a more general stochastic volatility model with the time-varying stochastic intensity,

$$\lambda(\xi_0, X_t, t) = \lambda_0(\xi_0, t) + \lambda_1(\xi_0, t) X_t,$$

where ξ_0 is the size of the previous jump. This specification yields a class of jump Lévy measures which combine the features of jump intensities depending on, say volatility, as well

as the size of the previous jump. Johannes, Kumar and Polson (1999) also propose a class of jump diffusion processes with a jump intensity depending on the past jump time and the absolute return. Moreover, as pointed out by Chernov, Gallant, Ghysels and Tauchen (2003), another potentially very useful specification of the intensity function would include the past duration, i.e., the time since the last jump, say $\tau(t)$, which is the time that has elapsed between the last jump and t where $\tau(t)$ is a continuous function of t , such as

$$\lambda(\xi_0, X_t, \tau, t) = \{\lambda_0(t) + \lambda_1(t) X_t\} \lambda\{\tau(t)\} \exp\{G(\xi_0)\}, \quad (2.10)$$

which can accommodate the increasing, decreasing or hump-shaped hazard functions of the size of the previous jump, and the duration dependence of jump intensities. However, to the best of our knowledge, there have not been any attempt in the literature to discuss the estimation and test of the intensity function $\lambda(\cdot)$ nonparametrically in the above settings.

A natural question arises is how to generalize the model (2.8) economically and statistically to a more general time-dependent jump diffusion model

$$dX_t = \mu(X_t, t) dt + \sigma(X_t, t) dB_t + dJ_t$$

with the time-dependent intensity function $\lambda(\xi_0, X_t, \tau, t)$ without any specified form or with some nonparametric structure, say, like (2.10). Clearly, they include the aforementioned models as a special case, which are studied by Duffie, Pan and Singleton (2000), Johannes, Kumar and Polson (1999), and Chernov, Gallant, Ghysels and Tauchen (2003), among others. This is still an open problem.

3 Nonparametric Estimation and Testing of Parametric Diffusions

3.1 Nonparametric Estimation of Parametric Diffusion Models

As is well-known, derivative pricing in mathematical finance is generally much more tractable in a continuous-time modelling framework than through binomial or other discrete approximations. In the empirical literature, however, it is an usual practice to abandon continuous-time modeling when estimating derivative pricing models. This is mainly due to the difficulty that the transition density for most continuous-time models with discrete observations has no closed form and therefore the maximum likelihood estimation (MLE) is infeasible.

One major focus of the continuous-time literature is on developing econometric methods

to estimate continuous-time models using discretely-sampled data.⁴ This is largely motivated by the fact that using the discrete version of a continuous-time model can result in inconsistent parameter estimates; see Lo (1988). Available estimation procedures include the MLE method of Lo (1988), the simulated methods of moments of Duffie and Singleton (1993) and Gouriéroux, Monfort and Renault (1993), the generalized method of moments (GMM) of Hansen and Scheinkman (1995), the efficient method of moments (EMM) of Gallant and Tauchen (1996), the Markov chain Monte Carlo (MCMC) of Jacquier, Polson and Rossi (1994), Eraker (1998) and Jones (1998), and the methods based on the empirical characteristic function of Jiang and Knight (2002) and Singleton (2001).

Below we focus on some nonparametric estimation methods of a parametric continuous-time model

$$dX_t = \mu(X_t, \theta) dt + \sigma(X_t, \theta) dB_t, \quad (3.1)$$

where $\mu(\cdot, \cdot)$ and $\sigma(\cdot, \cdot)$ are known functions, and θ is unknown parameter vector in an open bounded parameter space Θ . Ait-Sahalia (1996b) proposes a minimum distance estimator

$$\hat{\theta} = \arg \min_{\theta \in \Theta} n^{-1} \sum_{\tau=1}^n [\hat{\pi}_0(X_{\tau\Delta}) - \pi(X_{\tau\Delta}, \theta)]^2, \quad (3.2)$$

where

$$\hat{\pi}_0(x) = n^{-1} \sum_{\tau=1}^n K_h(x - X_{\tau\Delta})$$

is a kernel estimator for the stationary density of X_t , and

$$\pi(x, \theta) = \frac{c(\theta)}{\sigma^2(x, \theta)} \exp \left\{ \int_{x_0^*}^x \frac{2\mu(u, \theta)}{\sigma^2(u, \theta)} du \right\}, \quad (3.3)$$

is the marginal density estimator implied by the diffusion model, where the standardization factor $c(\theta)$ ensures that $\pi(\cdot, \theta)$ integrates to 1 for every $\theta \in \Theta$, and x_0^* is the lower bound of the support of X_t . Because the marginal density cannot capture the full dynamics of the diffusion process, one can expect that $\hat{\theta}$ will not be asymptotically most efficient, although it is root- n consistent for θ_0 if the parametric model is correctly specified.

Next, we introduce the approximate likelihood approach, due to Ait-Sahalia (2002a). Let $p_x(\Delta, x | x_0, \theta)$ be the conditional density function of $X_{\tau\Delta} = x$ given $X_{(\tau-1)\Delta} = x_0$ induced

⁴Sundaresan (2001) states that “perhaps the most significant development in the continuous-time field during the last decade has been the innovations in econometric theory and in the estimation techniques for models in continuous time.” For other reviews of the recent literature, see, e.g., Melino (1994), Tauchen (1997, 2001), Campbell, Lo and MacKinlay (1997), Cai and Hong (2003) and Fan (2005).

by the model (3.1). The log-likelihood function of the model for the sample is

$$l_n(\theta) = \sum_{\tau=1}^n \ln p_x(\Delta, X_{\tau\Delta} | X_{(\tau-1)\Delta}, \theta).$$

The MLE estimator that maximizes $l_n(\theta)$ would be asymptotically most efficient if the conditional density $p_x(\Delta, x|x_0, \theta)$ has a closed form. Unfortunately, except for some simple models, $p_x(\Delta, x|x_0, \theta)$ usually does not have a closed form.

Using the Hermite polynomial series, Ait-Sahalia (2002a) proposes a closed form sequence $\{p_x^{(J)}(\Delta, x|x_0, \theta)\}$ to approximate $p_x(\Delta, x|x_0, \theta)$, and then obtains an estimator $\hat{\theta}_n^{(J)}$ that maximizes the approximated model likelihood. The estimator $\hat{\theta}_n^{(J)}$ enjoys the same asymptotic efficiency as the (infeasible) MLE as $J = J_n \rightarrow \infty$. More specifically, Ait-Sahalia (2002a) first considers a transformed process

$$Y_t \equiv \gamma(X_t, \theta) = \int_{-\infty}^{X_t} \frac{1}{\sigma(u, \theta)} du.$$

This transformed process obeys the following diffusion

$$dY_t = \mu_y(Y_t, \theta)dt + dB_t,$$

where

$$\mu_y(y, \theta) = \frac{\mu[\gamma^{-1}(y, \theta), \theta]}{\sigma[\gamma^{-1}(y, \theta), \theta]} - \frac{1}{2} \frac{\partial \sigma[\gamma^{-1}(y, \theta), \theta]}{\partial x}.$$

The transform $X \rightarrow Y$ ensures that the tail of the transition density $p_y(\Delta, y|y_0, \theta)$ of Y_t will generally vanish exponentially fast so that Hermite series approximations will converge. However, $p_y(\Delta, y|y_0, \theta)$ may get peaked at y_0 when the sample frequency Δ gets smaller. To avoid this, Ait-Sahalia (2002a) considers a further transform

$$Z_t = \Delta^{-1/2}(Y_t - y_0)$$

and then approximates the transition density of Z_t by the Hermite polynomials:

$$p_z^{(J)}(z | z_0, \theta) = \phi(z) \sum_{j=0}^J \eta_z^{(j)}(z_0, \theta) H_j(z),$$

where $\phi(\cdot)$ is the $N(0, 1)$ density, and $\{H_j(z)\}$ is the Hermite polynomial series. The coefficients $\{\eta_z^{(j)}(z_0, \theta)\}$ are specific conditional moments of process Z_t , and can be explicitly computed using the Monte Carlo method or using a higher Taylor series expansion in Δ .

The approximated transition density of X_t is then given as follows:

$$\begin{aligned} p_x(x | x_0, \theta) &= \sigma(x, \theta)^{-1} p_y(\gamma(x, \theta) | \gamma(x_0, \theta), \theta) \\ &= \Delta^{-1/2} p_z(\Delta^{-1/2}(\gamma(x, \theta) - \gamma(x_0, \theta)) | \gamma(x_0, \theta), \theta). \end{aligned}$$

Under suitable regularity conditions, particularly when $J = J_n \rightarrow \infty$ as $n \rightarrow \infty$, the estimator

$$\hat{\theta}_n^{(J)} = \arg \min_{\theta \in \Theta} \sum_{\tau=1}^n \ln p_x^{(J)}(X_{\tau\Delta} | X_{(\tau-1)\Delta}, \theta)$$

will be asymptotically equivalent to the infeasible MLE. Ait-Sahalia (1999) applies this method to estimate a variety of diffusion models for spot interest rates, and finds that $J = 2$ or 3 already gives accurate approximation for most financial diffusion models. Egorov, Li and Xu (2003) extend this approach to stationary time-inhomogeneous diffusion models. Ait-Sahalia (2008) extends this method to general multivariate diffusion models and Ait-Sahalia and Kimmel (2007) to affine multi-factor term structure models.

Finally, in a rather general continuous-time setup which allows for stationary multi-factor diffusion models with partially observable state variables, Gallant and Tauchen (1996) propose an efficient method of moment estimator that also enjoys the asymptotic efficiency as the MLE. The basic idea of EMM is to first use a Hermite-polynomial based semi-nonparametric (SNP) density estimator to approximate the transition density of the observed state variables. This is called the auxiliary model and its score is called the score generator, which has expectation zero under the model-implied distribution when the parametric model is correctly specified. Then, given a parameter setting for the multi-factor model, one may use simulation to evaluate the expectation of the score under the stationary density of the model and compute a chi-square criterion function. A nonlinear optimizer is used to find the parameter values that minimize the proposed criterion.

Specifically, suppose $\{X_t\}$ is a stationary possibly vector valued process such that the true conditional density function $p_0(\Delta, X_{\tau\Delta} | X_{s\Delta}, s \leq \tau - 1) = p_0(\Delta, X_{\tau\Delta} | Y_{\tau\Delta})$, where $Y_{\tau\Delta} \equiv (X_{(\tau-1)\Delta}, \dots, X_{(\tau-d)\Delta})'$ for some fixed integer $d \geq 0$. This is a Markovian process of order d . To check the adequacy of a parametric model in (3.1), Gallant and Tauchen (1996) propose to check whether the following moment condition holds:

$$M(\beta_n, \theta) \equiv \int \frac{\partial \log f(\Delta, x, y; \beta_n)}{\partial \beta_n} p(\Delta, x, y; \theta) dx dy = 0, \quad \text{if } \theta = \theta_0 \in \Theta, \quad (3.4)$$

where $p(\Delta, x, y; \theta)$ is the model-implied joint density for $(X_{\tau\Delta}, Y_{\tau\Delta})'$, θ_0 is the unknown true parameter value, and $f(\Delta, x, y; \beta_n)$ is an auxiliary model for the conditional density of

$(X_{\tau\Delta}, Y'_{\tau\Delta})'$. Note that β_n is the parameter vector in the SNP density model $f(\Delta, x, y; \beta_n)$ and generally does not nest the parametric parameter θ . By allowing the dimension of β_n to grow with the sample size n , the SNP density $f(\Delta, x, y; \beta_n)$ will eventually span the true density $p_0(\Delta, x, y)$ of $(X_{\tau\Delta}, Y'_{\tau\Delta})'$, and thus is free of model misspecification asymptotically. Gallant and Tauchen (1996) use a Hermite polynomial approximation for $f(\Delta, x, y; \beta_n)$, with the dimension of β_n determined by such model selection criteria as BIC. The integration in (3.4) can be computed by simulating a large number of realizations under the distribution of the parametric model $p(\Delta, x, y; \theta)$.

The efficient method of moment estimator is defined as follows:

$$\hat{\theta} = \arg \min_{\theta \in \Theta} M(\hat{\beta}_n, \theta)' \hat{I}^{-1}(\theta) M(\hat{\beta}_n, \theta),$$

where $\hat{\beta}$ is the quasi-MLE for β_n , the coefficients in the Hermite polynomial expansion of the SNP density model $f(x, y, \beta_n)$ and the matrix $\hat{I}(\theta)$ is an estimate of the asymptotic variance of $\sqrt{n}\partial M_n(\hat{\beta}_n, \theta)/\partial\theta$ (Gallant and Tauchen, 2001). This estimator $\hat{\theta}$ is asymptotically as efficient as the (infeasible) MLE.

The EMM has been applied widely in financial applications. See, for example, Andersen and Lund (1997), Dai and Singleton (2000), Ahn, Dittmar and Gallant (2002) for interest rate applications, Liu (2000), Andersen, Berzoni and Lund (2002), Chernov, Gallant, Ghysels and Tauchen (2003) for estimating stochastic volatility models for stock prices with such complications as long memory and jumps, Chung and Tauchen (2001) for estimating and testing target zero models of exchange rates, Jiang and van der Sluis (2000) for price option pricing, and Valderamma (2001) for a macroeconomic application. It would be interesting to compare the EMM method and Ait-Sahalia's (2002a) approximate MLE in finite sample performance and this topic is still open.

3.2 Nonparametric Testing of Diffusion Models

In financial applications, most continuous-time models are parametric. It is important to test whether a parametric diffusion model adequately captures the dynamics of the underlying process. Model misspecification generally renders inconsistent estimators of model parameters and their variance-covariance matrix, leading to misleading conclusions in inference and hypothesis testing. More importantly, a misspecified model can yield large errors in hedging, pricing and risk management.

Unlike the vast literature of estimation of parametric diffusion models, there are relatively few test procedures for parametric diffusion models using discrete observations. Suppose

$\{X_t\}$ follows a continuous-time diffusion process in (2.5). Often it is assumed that the drift and diffusion $\mu(\cdot, t)$ and $\sigma(\cdot, t)$ have some parametric forms $\mu(\cdot, t, \theta)$ and $\sigma(\cdot, t, \theta)$, where $\theta \in \Theta$. We say that models $\mu(\cdot, t, \theta)$ and $\sigma(\cdot, t, \theta)$ are correctly specified for the drift and diffusion $\mu(\cdot, t)$ and $\sigma(\cdot, t)$ respectively if

$$\mathbb{H}_0 : P [\mu(X_t, t, \theta_0) = \mu(X_t, t), \sigma(X_t, t, \theta_0) = \sigma(X_t, t)] = 1 \text{ for some } \theta_0 \in \Theta. \quad (3.5)$$

As noted earlier, various methods have been developed to estimate θ_0 , taking (3.5) as given. However, these methods generally cannot deliver consistent parameter estimates if $\mu(\cdot, t, \theta)$ or $\sigma(\cdot, t, \theta)$ is misspecified in the sense that

$$\mathbb{H}_a : P [\mu(X_t, t, \theta) = \mu(X_t, t), \sigma(X_t, t, \theta) = \sigma(X_t, t)] < 1 \text{ for all } \theta \in \Theta. \quad (3.6)$$

Under \mathbb{H}_a of (3.6), there exists no parameter value $\theta \in \Theta$ such that the drift model $\mu(\cdot, t, \theta)$ and the diffusion model $\sigma(\cdot, t, \theta)$ coincide with the true drift $\mu(\cdot, t)$ and the true diffusion $\sigma(\cdot, t)$ respectively.

There is a growing interest in testing whether a continuous-time model is correctly specified using a discrete sample $\{X_{\tau\Delta}\}_{\tau=1}^n$. Next we will present some test procedures for testing the continuous-time models. Ait-Sahalia (1996b) observes that for a stationary time-homogeneous diffusion process in (3.1), a pair of drift and diffusion models $\mu(\cdot, \theta)$ and $\sigma(\cdot, \theta)$ uniquely determines the stationary density $\pi(\cdot, \theta)$ in (3.3). Ait-Sahalia (1996b) compares a parametric marginal density estimator $\pi(\cdot, \hat{\theta})$ with a nonparametric density estimator $\hat{\pi}_0(\cdot)$ via the quadratic form

$$M \equiv \int_{x_0^*}^{x_1^*} [\hat{\pi}_0(x) - \pi(x, \hat{\theta})]^2 \hat{\pi}_0(x) dx, \quad (3.7)$$

where x_1^* is the upper bound for X_t , $\hat{\theta}$ is the minimum distance estimator given by (3.2). The M statistic, after demeaning and scaling, is asymptotically normal under \mathbb{H}_0 .

The M test makes no restrictive assumptions on the data generating process and can detect a wide range of alternatives. This appealing power property is not shared by parametric approaches such as generalized method of moment tests (e.g., Conley *et al.* 1997). The latter has optimal power against certain alternatives (depending on the choice of moment functions) but may be completely silent against other alternatives. In an application to Euro-dollar interest rates, Ait-Sahalia (1996b) rejects all existing one-factor linear drift models using asymptotic theory and finds that “the principal source of rejection of existing models is the strong nonlinearity of the drift,” which is further supported by Stanton (1997).

However, several limitations of this test may hinder its empirical applicability. First, as Ait-Sahalia (1996b) has pointed out, the marginal density cannot capture the full dynamics of $\{X_t\}$. It cannot distinguish two diffusion models that have the same marginal density but different transition densities.⁵ Second, subject to some regularity conditions, the asymptotic distribution of the quadratic form M in (3.7) remains the same whether the sample $\{X_{\tau\Delta}\}_{\tau=1}^n$ is iid or highly persistently dependent (Ait-Sahalia, 1996b). This convenient asymptotic property unfortunately results in a substantial discrepancy between the asymptotic and finite sample distributions, particularly when data display persistent dependence (Pritsker, 1998). This discrepancy and the slow convergence of kernel estimators are the main reasons identified by Pritsker (1998) for the poor finite sample performance of the M test. They cast some doubt on the applicability of first order asymptotic theory of nonparametric methods in finance, since persistent serial dependence is a stylized fact for interest rates and many other high frequency financial data. Third, a kernel density estimator produces biased estimates near the boundaries of the data (e.g., Härdle, 1990 and Fan and Gijbels, 1996). In the present context, the boundary bias can generate spurious nonlinear drifts, giving misleading conclusions on the dynamics of $\{X_t\}$.

Recently, Hong and Li (2005) have developed a nonparametric test for the model in (2.5) using the transition density, which can capture the full dynamics of $\{X_t\}$ in (3.1). Let $p_0(x, t | x_0, s)$ be the true transition density of the diffusion process X_t ; that is, the conditional density of $X_t = x$ given $X_s = x_0$, $s < t$. For a given pair of drift and diffusion models $\mu(\cdot, t, \theta)$ and $\sigma(\cdot, t, \theta)$, a certain family of transition densities $\{p(x, t | x_0, s, \theta)\}$ is characterized. When (and only when) \mathbb{H}_0 in (3.5) holds, there exists some $\theta_0 \in \Theta$ such that $p(x, t | x_0, s, \theta_0) = p_0(x, t | x_0, s)$ almost everywhere for all $t > s$. Hence, the hypotheses of interest \mathbb{H}_0 in (3.5) versus \mathbb{H}_a in (3.6) can be equivalently written as follows:

$$\mathbb{H}_0 : p(x, t | y, s, \theta_0) = p_0(x, t | y, s) \quad \text{almost everywhere for some } \theta_0 \in \Theta \quad (3.8)$$

versus the alternative hypothesis

$$\mathbb{H}_a : p(x, t | y, s, \theta) \neq p_0(x, t | y, s) \quad \text{for some } t > s \text{ and for all } \theta \in \Theta. \quad (3.9)$$

Clearly, to test \mathbb{H}_0 in (3.8) versus \mathbb{H}_a in (3.9) would be to compare a model transition density estimator $p(x, t | x_0, s, \hat{\theta})$ with a nonparametric transition density estimator, say $\hat{p}_0(x, t | x_0, s)$. Instead of comparing $p(x, t | x_0, s, \hat{\theta})$ and $\hat{p}_0(x, t | x_0, s)$ directly, Hong and

⁵A simple example is the Vasicek model, where if we vary the speed of mean reversion and the scale of diffusion in the same proportion, the marginal density will remain unchanged, but the transition density will be different.

Li (2005) first transform $\{X_{\tau\Delta}\}_{\tau=1}^n$ via a probability integral transform. Define a discrete transformed sequence

$$Z_\tau(\theta) \equiv \int_{-\infty}^{X_{\tau\Delta}} p[x, \tau\Delta | X_{(\tau-1)\Delta}, (\tau-1)\Delta, \theta] dx, \quad \tau = 1, \dots, n. \quad (3.10)$$

Under (and only under) \mathbb{H}_0 in (3.8) there exists some $\theta_0 \in \Theta$ such that $p[x, \tau\Delta | X_{(\tau-1)\Delta}, (\tau-1)\Delta, \theta_0] = p_0[x, \tau\Delta | X_{(\tau-1)\Delta}, (\tau-1)\Delta]$ almost surely for all $\Delta > 0$. Consequently, the transformed series $\{Z_\tau \equiv Z_\tau(\theta_0)\}_{\tau=1}^n$ is iid $U[0, 1]$ under \mathbb{H}_0 in (3.8). This result is first proven, in a simpler context, by Rosenblatt (1952), and is more recently used to evaluate out-of-sample density forecasts (e.g., Diebold, Gunther and Tay, 1998) in a discrete-time context. Intuitively, we may call $\{Z_\tau(\theta)\}$ “generalized residuals” of the model $p(x, t | y, s, \theta)$.

To test \mathbb{H}_0 in (3.8), Hong and Li (2005) check whether $\{Z_\tau\}_{\tau=1}^n$ is both iid and $U[0, 1]$. They compare a kernel estimator $\hat{g}_j(z_1, z_2)$ defined in (3.11) below for the joint density of $\{Z_\tau, Z_{\tau-j}\}$ with unity, the product of two $U[0, 1]$ densities. This approach has at least three advantages. First, since there is no serial dependence in $\{Z_\tau\}$ under \mathbb{H}_0 in (3.8), nonparametric joint density estimators are expected to perform much better in finite samples. In particular, the finite sample distribution of the resulting tests is expected to be robust to persistent dependence in data. Second, there is no asymptotic bias for nonparametric density estimators under \mathbb{H}_0 in (3.8). Third, no matter whether $\{X_t\}$ is time-inhomogeneous or even nonstationary, $\{Z_\tau\}$ is always iid $U[0, 1]$ under correct model specification.

Hong and Li (2005) employ the kernel joint density estimator,

$$\hat{g}_j(z_1, z_2) \equiv (n-j)^{-1} \sum_{\tau=j+1}^n K_h(z_1, \hat{Z}_\tau) K_h(z_2, \hat{Z}_{\tau-j}), \quad j > 0, \quad (3.11)$$

where $\hat{Z}_\tau = Z_\tau(\hat{\theta})$, $\hat{\theta}$ is any \sqrt{n} -consistent estimator for θ_0 , and for $x \in [0, 1]$,

$$K_h(x, y) \equiv \begin{cases} h^{-1}k\left(\frac{x-y}{h}\right) / \int_{-(x/h)}^1 k(u)du, & \text{if } x \in [0, h), \\ h^{-1}k\left(\frac{x-y}{h}\right), & \text{if } x \in [h, 1-h], \\ h^{-1}k\left(\frac{x-y}{h}\right) / \int_{-1}^{(1-x)/h} k(u)du, & \text{if } x \in (1-h, 1] \end{cases}$$

is the kernel with boundary correction (Rice, 1986) and $k(\cdot)$ is a standard kernel. This avoids the boundary bias problem, and has some advantages over some alternative methods such as trimming and the use of the jackknife kernel.⁶ To avoid the boundary bias problem, one

⁶One could simply ignore the data in the boundary regions and only use the data in the interior region. Such a trimming procedure is simple, but in the present context, it would lead to the loss of significant

might apply other kernel smoothing methods such as local polynomial (Fan and Gijbels, 1996) or weighted NW (Cai, 2001).

Hong and Li's (2005) test statistic is

$$\widehat{Q}(j) \equiv \left[(n-j)h \int_0^1 \int_0^1 [\widehat{g}_j(z_1, z_2) - 1]^2 dz_1 dz_2 - A_h^0 \right] / V_0^{1/2},$$

where A_h^0 and V_0 are non-stochastic centering and scale factors that depends on h and $k(\cdot)$.

In a simulation experiment mimicking the dynamics of U.S. interest rates via the Vasicek model, Hong and Li (2005) find that $\widehat{Q}(j)$ has rather reasonable sizes for $n = 500$ (i.e., about two years of daily data). This is a rather substantial improvement over Ait-Sahalia's (1996b) test, in lights of Pritsker's (1998) simulation evidence. Moreover, $\widehat{Q}(j)$ has better power than the marginal density test. Hong and Li (2005) find extremely strong evidence against a variety of existing one-factor diffusion models for the spot interest rate and affine models for interest rate term structures. Egorov, Hong and Li (2006) have recently extended Hong and Li (2005) to evaluate out-of-sample of density forecasts of a multivariate diffusion model possibly with jumps and partially unobservable state variables.

Because the transition density of a continuous-time model generally has no closed form, the probability integral transform $\{Z_\tau(\theta)\}$ in (3.10) is difficult to compute. However, one can approximate the model transition density using the simulation methods developed by (e.g.) Pedersen (1995), Brandt and Santa-Clara (2002), and Elerian, Chib and Shephard (2001). Alternatively, we can use Ait-Sahalia's (2002a) Hermite expansion method to construct a closed-form approximation of the model transition density.

When a misspecified model is rejected, one may like to explore what are possible sources for the rejection. For example, is the rejection due to misspecification in the drift, such as the ignorance of mean shifts or jumps? Is it due to the ignorance of GARCH effects or stochastic volatility? Or is it due to the ignorance of asymmetric behaviors (e.g., leverage effects)? Hong and Li (2005) consider to examine the autocorrelations in the various powers of $\{Z_\tau\}$, which are very informative about how well a model fits various dynamic aspects of

amount of information. If $h = sn^{-\frac{1}{5}}$ where $s^2 = \text{var}(X_t)$, for example, then about 23, 20 and 10 of a uniformly distributed sample will fall into the boundary regions when $n = 100, 500$ and $5,000$ respectively. For financial time series, one may be particularly interested in the tail distribution of the underlying process, which is exactly contained in (and only in) the boundary regions!

Another solution is to use a kernel that adapts to the boundary regions and can effectively eliminate the boundary bias. One example is the so-called jackknife kernel, as used in Chapman and Pearson (2000). In the present context, the jackknife kernel, however, has some undesired features in finite samples. For example, it may generate negative density estimates in the boundary regions because the jackknife kernel can be negative in these regions. It also induces a relatively large variance for the kernel estimates in the boundary regions, adversely affecting the power of the test in finite samples.

the underlying process (e.g., conditional mean, variance, skewness, kurtosis, ARCH-in-mean effect, and leverage effect).

Gallant and Tauchen (1996) also propose an EMM-based minimum chi-square specification test for stationary continuous-time models. They examine the simulation-based expectation of an auxiliary SNP score function under the model distribution, which is zero under correct model specification. The greatest appeal of the EMM approach is that it applies to a wide range of stationary continuous-time processes, including both one-factor and multi-factor diffusion processes with partially observable state variables (e.g., stochastic volatility models). In addition to the minimum chi-square test for generic model mis-specifications, the EMM approach also provides a class of individual t -statistics that are informative in revealing possible sources of model misspecification. This is perhaps the most appealing strength of the EMM approach.

Another feature of the EMM tests is that all EMM test statistics avoid estimating long-run variance-covariances, thus resulting in reasonable finite sample size performance (cf. Anderson, Chung and Sorensen, 1999). In practice, however, it may not be easy to find an adequate SNP density model for financial time series, as is shown in Hong and Lee (2003b). For example, Andersen and Lund (1997) find that an AR(1)-EGARCH model with a number of Hermite polynomials adequately captures the full dynamics of daily S&P 500 return series, using a BIC criterion. However, Hong and Lee (2003a) find that there still exists strong evidence on serial dependence in the standardized residuals of the model, indicating that the auxiliary SNP model is inadequate. This affects the validity of the EMM tests, because their asymptotic variance estimators have exploit the correct specification of the SNP density model.⁷

There has been also an interest in separately testing the drift model and the diffusion model in (3.1). For example, it has been controversial whether the drift of interest rates is linear. To test the linearity of the drift term, one can write it as a functional coefficient form (Cai, Fan and Yao, 2000) $\mu(X_t) = \alpha_0(X_t) + \alpha_1(X_t) X_t$. Then, the null hypothesis is $\mathbb{H}_0 : \alpha_0(\cdot) \equiv \alpha_0$ and $\alpha_1(\cdot) \equiv \alpha_1$. Fan and Zhang (2003) apply the generalized likelihood ratio test developed by Cai, Fan and Yao (2000) and Fan, Zhang and Zhang (2001). They find that \mathbb{H}_0 is not rejected for the short-term interest rates. It is noted that the asymptotic theory for the generalized likelihood ratio test is developed for the iid samples but it is still unknown whether it is valid for a time series context. One might follow the idea from

⁷Chen, Gao and Tang (2008) consider kernel-based simultaneous specification testing for both mean and variance models in a discrete-time setup with dependent observations. The empirical likelihood principle is used to construct the test statistic. They apply the test to check adequacy of a discrete version of a continuous-time diffusion model.

Cai, Fan and Yao (2000) to use the bootstrap or wild bootstrap method instead of the asymptotic theory for time series context. Fan and Zhang (2003) and Fan, Jiang, Zhang and Zhou (2003) conjecture that it would hold based on their simulations. On the other hand, Chen, Härdle and Kleinow (2002) consider an empirical likelihood goodness-of-fit test for time series regression model, and they apply the test to test a discrete drift model of a diffusion process.

There has been also interest in testing the diffusion model $\sigma(\cdot, \theta)$. The motivation comes from the fact that derivative pricing with an underlying equity process only depends on the diffusion $\sigma(\cdot)$, which is one of the most important features of (3.1) for derivative pricing. Kleinow (2002) recently proposes a nonparametric test for a diffusion model $\sigma(\cdot)$. More specifically, Kleinow (2002) compares a nonparametric diffusion estimator $\hat{\sigma}^2(\cdot)$ with a parametric diffusion estimator $\sigma^2(\cdot, \theta)$ via an asymptotically χ^2 test statistic

$$\hat{T}_k = \sum_{l=1}^k \left[\hat{T}(x_l) \right]^2,$$

where

$$\hat{T}(x) = [nh\hat{\pi}(x)]^{1/2} \left[\hat{\sigma}^2(x) / \tilde{\sigma}^2(x, \hat{\theta}) - 1 \right],$$

$\hat{\theta}$ is an \sqrt{n} -consistent estimator for θ_0 and

$$\tilde{\sigma}^2(x, \theta) = \frac{1}{nh\hat{\pi}(x)} \sum_{t=1}^n \sigma^2(x, \hat{\theta}) K_h[(x - X_t)/h]$$

is a smooth version of $\sigma^2(x, \theta)$. The use of $\tilde{\sigma}^2(x, \hat{\theta})$ instead of $\sigma^2(x, \hat{\theta})$ directly reduces the kernel estimation bias in $\hat{T}(x)$, thus allowing the use of the optimal bandwidth h for $\hat{\sigma}^2(x)$. This device is also used in Härdle and Mammen (1993) in testing a parametric regression model. Kleinow (2002) finds that the empirical level of \hat{T}_k is too large relative to the significance level in finite samples and then proposes a modified test statistic using the empirical likelihood approach, which endogenously studentizes conditional heteroscedasticity. As expected, the empirical level of the modified test improves in finite samples, though not necessarily for the power of the test.

Finally, Fan, Jiang, Zhang and Zhou (2003) test whether the coefficients in the time-varying coefficient single factor diffusion model of (2.6) are really time-varying. Specially, they apply the generalized likelihood ratio test to check whether some or all of $\{\alpha_j(\cdot)\}$ and $\{\beta_j(\cdot)\}$ are constant.

4 Derivative Pricing and Risk Neutral Density Estimation

4.1 Risk Neutral Density

In modern finance, the pricing of contingent claims is important given the phenomenal growth in turnover and volume of financial derivatives over the past decades. Derivative pricing formulas are highly nonlinear even when they are available in closed form. Nonparametric techniques are expected to be very useful in this area. In a standard dynamic exchange economy, the equilibrium price of a security at date t with a single liquidating payoff $Y(C_T)$ at date T that is a function of aggregate consumption C_T is given by

$$P_t = E_t [Y(C_T)M_{t,T}], \quad (4.1)$$

where the conditional expectation is taken with respect to the information set available to the representative economic agent at time t , $M_{t,T} = \delta^{T-1}U'(C_T)/U'(C_t)$, the so-called stochastic discount factor (SDF), is the marginal rate of substitution (MRS) between dates t and T , δ is the rate of time preference, and $U(\cdot)$ is the utility function of the economic agent. This is the stochastic Euler equation, or the first order condition of the intertemporal utility maximization of the economic agent with suitable budget constraints (e.g., Cochrane, 1996, 2001). It holds for all securities, including assets and various derivatives. All capital asset pricing models (CAPM) and derivative pricing models can be embedded in this unified framework — each model can be viewed as a specific specification of $M_{t,T}$. See Cochrane (1996, 2001) for an excellent discussion.

There have been some parametric tests for CAMP models (e.g., Hansen and Janaganan, 1997). To our knowledge, there are only a few nonparametric tests available in the literature for testing CAMP models based on the kernel method; see Wang (2002, 2003) and Cai, Kuan and Sun (2008), which will be elaborated in detail in Section 4.3 later. Also, all the tests for CAMP models are formulated in terms of discrete time frameworks. We focus on nonparametric derivative pricing in Section 4.2 and the nonparametric asset pricing will be discussed separately in Section 4.3.

Assuming that the conditional distribution of future consumption C_T has a density representation $f_t(\cdot)$, then the conditional expectation can be expressed as

$$E_t [Y(C_T)M_{t,T}] = \exp(-\tau r_t) \int Y(C_T)f_t^*(C_T)dC_T = \exp(-\tau r_t) E_t^* [Y(C_t)],$$

where r_t is the risk-free interest rate, $\tau = T - t$, and

$$f_t^*(C_T) = \frac{M_{t,T}f_t(C_T)}{\int M_{t,T}f_t(C_T)dC_T}$$

is called the risk neutral density (RND) function. See Taylor (2005, Chapter 16) for details about the definition and estimation methods. This function is also called the risk-neutral pricing probability (Cox and Ross, 1976), or equivalent martingale measure (Harrison and Kreps, 1979), or the state-price density (SPD). It contains rich information on the pricing and hedging of risky assets in an economy, and can be used to price other assets, or to recover the information about the market preferences and asset price dynamics (Bahra 1997, Jackwerth 1999). Obviously, the RND function differs from $f_t(C_T)$, the physical density function of C_T conditional on the information available at time t .

4.2 Nonparametric Derivative Pricing

In order to calculate an option price from (4.1), one has to make some assumption on the data generating process of the underlying asset, $\{P_t\}$. For example, Black and Scholes (1973) assume that the underlying asset follows a geometric Brownian motion:

$$dP_t = \mu P_t dt + \sigma P_t dB_t,$$

where μ and σ are two constants. Applying Ito's Lemma, one can show immediately that P_τ follows a lognormal distribution with parameter $(\mu - \frac{1}{2}\sigma^2)\tau$ and $\sigma\sqrt{\tau}$. Using a no-arbitrage argument, Black and Scholes (1973) show that options can be priced if investors are risk neutral by setting the expected rate of return in the underlying asset, μ , equal to the risk-free interest rate, r . Specifically, the European call option price is

$$\pi(K_t, P_t, r, \tau) = P_t \Phi(d_t) - e^{-r\tau} K_t \Phi(d_t - \sigma\sqrt{\tau}), \quad (4.2)$$

where K_t is the strike price, $\Phi(\cdot)$ is the standard normal cumulative distribution function and $d_t = \{\ln(P_t/K_t) + (r + \frac{1}{2}\sigma^2)\tau\} / (\sigma\sqrt{\tau})$. In (4.2), the only parameter that is not observable at a time t is σ . This parameter, when multiplied with $\sqrt{\tau}$, is the underlying asset return volatility over the remaining life of the option. An knowledge of σ can be inferred from the prices of options traded in the markets: given an observed option price, one can solve an appropriate option pricing model for σ which is essentially a market estimate of the future volatility of the underlying asset returns. This estimate of σ is known as “implied volatility”.

The most important implication of Black-Scholes option pricing is that when the option is correctly priced, the implied volatility σ^2 should be the same across all exercise prices of options on the same underlying asset and with the same maturity date. However, the implied volatility observed in the market is usually a convex function of exercise price, which is often referred to as the “volatility smile”. This indicates that market participants make

more complicated assumptions than the geometric Brownian motion for the dynamics of the underlying asset. In particular, the convexity of “volatility smile” indicates the degree to which the market RND function has a heavier tail than a lognormal density. A great deal of effort has been made to use alternative models for the underlying asset to smooth out the volatility smile and so to achieve higher accuracy in pricing and hedging.

A more general approach to derivative pricing is to estimate the RND function directly from the observed option prices and then use it to price derivatives or to extract market information. To obtain better estimation of the RND function, several econometric techniques have been introduced. These methods are all based on the following fundamental relation between option prices and RNDs: Suppose $G_t = G(K_t, P_t, r_t, \tau)$ is the option pricing formula. Then there is a close relation between the second derivative of G_t with respect to the strike price K_t and the RND function:

$$\frac{\partial^2 G_t}{\partial K_t^2} = \exp(-\tau r_t) f_t^*(P_t). \quad (4.3)$$

This is first shown by Breeden and Litzenberger (1978) in a time-state preference framework.

Most commonly used estimation methods for RNDs are various parametric approaches. One of them is to assume that the underlying asset follows a parametric diffusion process, from which one can obtain the option pricing formula by a no-arbitrage argument, and then obtain the RND function from (4.3) (see, e.g., Bates 1991, 2000, Anagnou, Bedendo, Hodges and Tompkins 2005). Another parametric approach is to directly impose some form for the RND function and then estimate unknown parameters by minimizing the distance between the observed option prices and those generated by the assumed RND function (e.g., Jackwerth and Rubinstein, 1996, Melick and Thomas, 1997, Rubinstein, 1994). A third parametric approach is to assume a parametric form for the call pricing function or the implied volatility smile curve and then apply (4.3) to get the RND function (Bates 1991, Jarrow and Tudd, 1982, Longstaff, 1992, 1995, Shimko, 1993).

The aforementioned parametric approaches all impose certain restrictive assumptions, directly or indirectly, on the data generating process as well as the stochastic discount factor in some cases. The obtained RND function is not robust to the violation of these restrictions. To avoid this drawback, Ait-Sahalia and Lo (1998) use a nonparametric method to extract the RND function from option prices.

Given observed call option prices $\{G_t, K_t, \tau\}$, the price of the underlying asset $\{P_t\}$, and the risk free rate of interest $\{r_t\}$, Ait-Sahalia and Lo (1998) construct a kernel-estimator for

$E(G_t|P_t, K_t, \tau, r_t)$. Under standard regularity conditions, Ait-Sahalia and Lo (1998) show that the RND estimator is consistent and asymptotically normal and they provide explicit expressions for the asymptotic variance of the estimator.

Armed with the RND estimator, Ait-Sahalia and Lo (1998) apply it to the pricing and delta-hedging of S&P 500 call and put options using daily data obtained from the Chicago Board Options Exchange for the sample period from January 4, 1993 to December 31, 1993. The RND estimator exhibits negative skewness and excess kurtosis, a common feature of historical stock returns. Unlike many parametric option pricing models, the RND-generated option pricing formula is capable of capturing persistent “volatility smiles” and other empirical features of market prices. Ait-Sahalia and Lo (2000) use a nonparametrical RND estimator to compute the economic value at risk, that is, the value at risk of the RND function.

The artificial neural network (ANN) has received much attention in economics and finance over the last decade. Hutchinson, Lo and Poggio (1994), Anders, Korn and Schmitt (1998) and Hanke (1999) have successfully applied the ANN models to estimate pricing formulas of financial derivatives. In particular, Hutchinson, Lo and Poggio (1994) use the ANN to address the following question: If option prices are truly determined by the Black-Scholes formula exactly, can ANN “learn” the Black-Scholes formula? In other words, can the Black-Scholes formula be estimated nonparametrically via learning networks with a sufficient degree of accuracy to be of practical use? Hutchinson, Lo and Poggio (1994) perform Monte Carlo simulation experiments in which various ANNs are trained on artificially generated Black-Scholes formula and then compare to the Black-Scholes formula both analytically and in out-of-sample hedging experiments. They begin by simulating a two-year sample of daily stock prices, and creating a across-section of options each day according to the rules used by the Chicago Broad Options Exchange with prices given by the Black-Scholes formula. They find that, even with training sets of only six months of daily data, learning network pricing formulas can approximate the Black-Scholes formula with reasonable accuracy. The nonlinear models obtained from neural networks yield estimates option prices and deltas that are difficult to distinguish visually from the true Black-Scholes values.

Based on the economic theory of option pricing, the price of a call option should be a monotone decreasing convex function of the strike price and the state price density which is proportional to the second derivative of the call function; see (4.3), is a valid density function over future values of the underlying asset price, and hence must be nonnegative and integrate to one. Therefore, Yatchew and Härdle (2006) combine shape restrictions with

nonparametric regression to estimate the call price function and the SPD within a single least squares procedure. Constraints include smoothness of various order derivatives, monotonicity and convexity of the call function and integration to one of the SPD. Confidence intervals and test procedures are implemented using bootstrap methods. In addition, they apply the procedures to option data on the DAX index.

There are several directions of further research on nonparametric estimation and testing of RNDs for derivative pricing. First, how to evaluate the quality of a RND function estimated from option prices? In other words, how to judge how well an estimated RND function reflects the market expected uncertainty of the underlying asset? Because the RND function differs from the physical probability density function of the underlying asset, the valuation of the RND function is rather challenging. The method developed by Hong and Li (2005) cannot be applied directly. One possible way to evaluate the RND function is to assume a certain family of utility functions for the representative investor, as in Rubinstein (1994) and Anagnou, Bedendo, Hodges and Tompkins (2005). Based on this assumption, one can obtain the stochastic discount factor and then the physical probability density function, to which Hong and Li's (2005) test can be applied. However, the utility function of the economic agent is not observable. Thus, when the test delivers a rejection, it may be due to either misspecification of the utility function or misspecification of the data generating process, or both. More fundamentally, it is not clear whether the economy can be a proxy by an representative agent.

A practical issue in recovering the RND function is the limitation of option prices data with certain common characterizations. In other words, the sample size of option price data could be small in many applications. As a result, nonparametric methods should be carefully developed to fit the problems at hand.

Most econometric techniques to estimate the RND function is restricted to European options, while many of the more liquid exchange-traded options are often American. Rather complex extensions of the existing methods, including the nonparametric ones, are required in order to estimate the RND functions from the prices of American options. This is an interesting and practically important direction for further research.

4.3 Nonparametric Asset Pricing

The capital asset pricing model and the arbitrage asset pricing theory (APT) have been the cornerstone in theoretical and empirical finance for decades. The classical CAPM usually assumes a simple and stable linear relationship between an asset's systematic risk and

its expected return; see the books by Campbell, Lo and MacKinlay (1997) and Cochrane (2001) for details. However, this simple relationship assumption has been challenged and rejected by several recent studies based on empirical evidences of time variation in betas and expected returns (as well as return volatilities). As with other models, one considers the conditional CAPMs or nonlinear APT with time-varying betas to characterize the time variation in betas and risk premia. In particular, Fama and French (1992, 1993, 1995) use some instrumental variables such as book-to-market equity ratio and market equity, as proxies for some unidentified risk factors to explain the time variation in returns. Whereas Ferson (1989), Harvey (1989), Ferson and Harvey (1991, 1993, 1998, 1999), Ferson and Korajczyk (1995), and Jagannathan and Wang (1996) conclude that beta and market risk premium vary over time, therefore, static CAPM should incorporate time variation in beta in the model. Although there is a vast amount of empirical evidences on time variation in betas and risk premia, there is no theoretical guidance on how betas and risk premia vary with time or variables that represent conditioning information. Many recent studies focus on modelling the variation in betas using continuous approximation and the theoretical framework of the conditional CAPM; see Cochrane (1996), Jagannathan and Wang (1996, 2002), Wang (2002, 2003) and Ang and Liu (2004) and the references therein. Recently, Ghysels (1998) discusses the problem in detail and stresses the impact of misspecification of beta risk dynamics on inference and estimation. Also, he argues that betas change through time very slowly and linear factor models like the conditional CAPM may have a tendency to overstate the time variation. Further, he shows that among several well-known time-varying beta models, a serious misspecification produces time variation in beta that is highly volatile and leads to large pricing errors. Finally, he concludes that it is better to use the static CAPM in pricing when we do not have a proper model to capture time variation in betas correctly.

It is well documented that large pricing errors could be due to the linear approach used in a nonlinear model and treating a non-linear relationship as a linear could lead to serious prediction problems in estimation. To overcome these problems, some nonlinear models have been considered in the recent literature. Following are some examples. Bansal, Hsieh and Viswanathan (1993) and Bansal and Viswanathan (1993) advocate the idea of a flexible stochastic discount factor model in empirical asset pricing and they focus on nonlinear arbitrage pricing theory models by assuming that the SDF is a nonlinear function of a few state variables. Further, Akdeniz, Altay-Salih and Caner (2003) test for the existence of significant evidence of non-linearity in the time series relationship of industry returns with market returns using the heteroskedasticity consistent Lagrange multiplier test of Hansen (1996) under the framework of the threshold model and they find that there exists statistically significant non-linearity in this relationship with respect to real interest rates. Wang

(2002, 2003) explores a nonparametric form of the SDF model and conducted a test based on the nonparametric model. Parametric models for time-varying betas can be most efficient if the underlying betas are correctly specified. However, a misspecification may cause serious bias and model constraints may distort the betas in local area.

To follow the nations from Bansal, Hsieh and Viswanathan (1993), Bansal and Viswanathan (1993), Ghysels (1998) and Wang (2002, 2003), which are slightly different from those used in (4.1), a very simplified version of the SDF framework for asset pricing admits a basic pricing representation, which is a special case of the model (4.1),

$$E[m_{t+1} r_{i,t+1} | \Omega_t] = 0, \quad (4.4)$$

where Ω_t denotes the information set at time t , m_{t+1} is the SDF or the pricing kernel, and $r_{i,t+1}$ is the excess return on the i -th asset or portfolio. Here $\varepsilon_{t+1} = m_{t+1} r_{i,t+1}$ is called the pricing error. In empirical finance, different models impose different constraints on the SDF. Particularly, the SDF is usually assumed to be a linear function of factors in various applications and then it becomes the well known CAP model; see Jagannathan and Wang (2002) and Wang (2003). Indeed, Jagannathan and Wang (2002) give the detailed comparison of the SDF and CAPM representations. Further, when the SDF is fully parameterized such as linear form, the general method of moments (GMM) of Hansen (1982) can be used to estimate parameters and test the model; see Campbell, Lo and MacKinlay (1997) and Cochrane (2001) for details.

Recently, Bansal, Hsieh and Viswanathan (1993) and Bansal and Viswanathan (1993) assume that m_{t+1} is a nonlinear function of a few state variables. Since the exact form of the nonlinear pricing kernel is unknown, Bansal and Viswanathan (1993) suggest using the polynomial expansion to approximate it and then apply the GMM for estimating and testing. As pointed by Wang (2003), although this approach is intuitive and general, one of the shortcomings is that it is difficult to obtain the distribution theory and the effective assessment of finite sample performance. To overcome this difficulty, instead of considering the nonlinear pricing kernel, Ghysels (1998) focuses on the nonlinear parametric model and uses a set of moment conditions suitable for GMM estimation of parameters involved. Wang (2003) studies the nonparametric conditional CAPM and gives an explicit expression for the pricing kernel m_{t+1} ; that is, $m_{t+1} = 1 - b(Z_t) r_{p,t+1}$, where Z_t is a $k \times 1$ vector of conditioning variables from Ω_t , $b(Z_t) = E(r_{p,t+1}|Z_t)/E(r_{p,t+1}^2|Z_t)$ is an unknown function, and $r_{p,t+1}$ is the return on the market portfolio in excess of the riskless rate. Since the functional form of $b(\cdot)$ is unknown, Wang (2003) suggests estimating $b(\cdot)$ by using the Nadaraya-Watson method to two regression functions $E(r_{p,t+1}|Z_t)$ and $E(r_{p,t+1}^2|Z_t)$. Also, he conducts a simple

nonparametric test about the pricing error. Indeed, his test is the well known F -test by running a multiple regression of the estimated pricing error $\widehat{\varepsilon}_{t+1}$ versus a group of information variables; see (4.9) later for details. Further, Wang (2003) extends this setting to multifactor models by allowing $b(\cdot)$ to change over time; that is, $b(Z_t) = b(t)$. Finally, Bansal, Hsieh and Viswanathan (1993), Bansal and Viswanathan (1993), and Ghysels (1998) do not assume that m_{t+1} is a linear function of $r_{p,t+1}$ and instead they consider a parametric model by using the polynomial expansion.

To combine the models studied by Bansal, Hsieh and Viswanathan (1993), Bansal and Viswanathan (1993), Ghysels (1998) and Wang (2002, 2003), and some other models in the finance literature under a very general framework, Cai, Kuan and Sun (2008) assume that the nonlinear pricing kernel has the form as $m_{t+1} = 1 - m(Z_t)r_{p,t+1}$, where $m(\cdot)$ is unspecified and they focus on the following nonparametric ATP model

$$E[\{1 - m(Z_t)r_{p,t+1}\} r_{i,t+1} | \Omega_t] = 0, \quad (4.5)$$

where $m(\cdot)$ is an unknown function of Z_t which is a $k \times 1$ vector of conditioning variables from Ω_t . Indeed, (4.5) can be regarded as a moment (orthogonal) condition. The main interest of (4.5) is to identify and estimate the function $m(Z_t)$ as well as test whether the model is correctly specified.

Let I_t be a $q \times 1$ ($q \geq k$) vector of conditional variables from Ω_t , including Z_t , satisfying the following orthogonal condition

$$E[\{1 - m(Z_t)r_{p,t+1}\} r_{i,t+1} | I_t] = 0, \quad (4.6)$$

which can be regarded as an approximation of (4.5). It follows from the orthogonality condition in (4.6) that, for any vector function $Q(V_t) \equiv Q_t$ with a dimension d_q specified later,

$$E [Q_t \{1 - m(Z_t)r_{p,t+1}\} r_{i,t+1} | I_t] = 0,$$

and its sample version is

$$\frac{1}{T} \sum_{t=1}^T Q_t \{1 - m(Z_t)r_{p,t+1}\} r_{i,t+1} = 0. \quad (4.7)$$

Therefore, Cai, Kuan and Sun (2008) propose a new nonparametric estimation procedure to combine the orthogonality conditions given in (4.7) with the local linear fitting scheme of Fan and Gijbels (1996) to estimate the unknown function $m(\cdot)$. This nonparametric estimation approach is called by Cai, Kuan and Sun (2008) as the nonparametric generalized method of moment (NPGMM).

For given the grid point z_0 and $\{Z_t\}$ in a neighborhood of z_0 , the orthogonality conditions in (4.7) can be approximated by the following locally weighted orthogonality conditions

$$\sum_{t=1}^T Q_t [1 - (a - b^T (Z_t - z_0))r_{p,t+1}] r_{i,t+1} K_h(Z_t - z_0) = 0, \quad (4.8)$$

where $K_h(\cdot) = h^{-k}K(\cdot/h)$, $K(\cdot)$ is a kernel function in \mathbb{R}^k , and $h = h_n > 0$ is a bandwidth, which controls the amount of smoothing used in the estimation. (4.8) can be viewed as a generalization of the nonparametric estimation equations in Cai (2003) and the locally weighted version of (9.2.29) in Hamilton (1994, p.243). Therefore, solving the above equations leads to the NPGMM estimate of $m(z_0)$, denoted by $\widehat{m}(z_0)$, which is \widehat{a} , where $(\widehat{a}, \widehat{b})$ is the minimizer of (4.8). Cai, Kuan and Sun (2008) discuss how to choose Q_t and derive the asymptotic properties of the proposed nonparametric estimator.

Let $\widehat{e}_{i,t+1}$ be the estimated pricing error; that is, $\widehat{e}_{i,t+1} = \widehat{m}_{t+1} r_{i,t+1}$, where $\widehat{m}_{t+1} = 1 - \widehat{m}(Z_t)r_{p,t+1}$. To test $E(e_{i,t+1} | \Omega_t) = 0$, Wang (2002, 2003) consider a simple test as follows. First, to run a multiple regression

$$\widehat{e}_{i,t+1} = V_t^T \delta_i + v_{i,t+1}, \quad (4.9)$$

where V_t is a $q \times 1$ ($q \geq k$) vector of observed variables from Ω_t ,⁸ and then test if all the regression coefficients are zero; that is, $\mathbb{H}_0 : \delta_1 = \dots = \delta_q = 0$. Also, Wang (2002) discusses two alternative test procedures. Indeed, the above model can be viewed as a liner approximation of $E[e_{i,t+1} | V_t]$. To examine the magnitude of pricing errors, Ghysels (1998) considers the mean square error (MSE) as a criterion to test if the conditional CAPM or APT model is misspecified relative to the unconditional one.

To check the misspecification of the model, Cai, Kuan and Sun (2008) construct a consistent nonparametric test based on a U-Statistics technique, described as follows. The interest is to test the hypotheses

$$\mathbb{H}_0 : E(e_{i,t+1} | I_t) = 0 \quad \text{versus} \quad \mathbb{H}_a : E(e_{i,t+1} | I_t) \neq 0, \quad (4.10)$$

where I_t is a $q \times 1$ ($q \geq k$) vector of observed variables from Ω_t . Similar to Wang (2003), I_t is taken to be Z_t . It is clear that $E(e_{i,t+1} | Z_t) = 0$ if and only if $[E(e_{i,t+1} | Z_t)]^2 f(Z_t) = 0$ if and only if $E[e_{i,t+1} E(e_{i,t+1} | Z_t) f(Z_t)] = 0$, where $f(\cdot)$ is the density of Z_t . Interestingly, the testing problem on conditional moment becomes unconditional. Obviously, the test statistic could be postulated as

$$U_T = \frac{1}{T} \sum_{t=1}^T e_{i,t+1} E(e_{i,t+1} | Z_t) f(Z_t), \quad (4.11)$$

⁸Wang (2003) takes V_t to be Z_t in his empirical analysis.

if $e_{i,t+1}E(e_{i,t+1}|Z_t)f(Z_t)$ would be known. Since $E(e_{i,t+1}|Z_t)f(Z_t)$ is unknown, its leave-one-out Nadaraya-Watson estimator can be formulated as

$$\widehat{E}(e_{i,t+1}|Z_t)f(Z_t) = \frac{1}{T(T-1)} \sum_{s \neq t}^T e_{i,t+1} K_h(Z_s - Z_t). \quad (4.12)$$

Plugging (4.12) into (4.11) and replacing $e_{i,t+1}$ by its estimate $\widehat{e}_{i,t+1} = \widehat{e}_t$, one obtain the test statistic, denoted by \widehat{U}_T ,

$$\widehat{U}_T = \frac{1}{T(T-1)} \sum_{s \neq t} K_h(Z_s - Z_t) \widehat{e}_s \widehat{e}_t, \quad (4.13)$$

which is indeed a second order U-statistics. Finally, Cai, Kuan and Sun (2008) show that this nonparametric test statistic is consistent. In addition, they apply the proposed testing procedure to test if either the CAPM or the Fama and French model, in the flexible nonparametric form, can explain the momentum profit which is the value-weighted portfolio of NYSE stocks as the market portfolio, using the dividend price ratio (DPR), the default premium (DEF), the one-month Treasury bill rate (RTB), and the excess return on the NYSE equally weighted portfolio (EWR) as the conditioning variables.

5 Nonparametric Predictive Models for Asset Returns

The predictability of stock returns has been studied for the last two decades as a cornerstone research topic in economics and finance. In many financial applications such as the mutual fund performance, the conditional capital asset pricing, and the optimal asset allocations, people routinely examine the predictability problem. Tremendous empirical studies document the predictability of stock returns using various lagged financial variables, such as the log dividend-price ratio, the log earning-price ratio, the log book-to-market ratio, the dividend yield, the term spread and default premium, and the interest rates. Important questions are often asked about whether the returns are predictable and whether the predictability is stable over time. Since many of the predictive financial variables are highly persistent and even nonstationary, it is really challenging econometrically or statistically to answer these questions.

Predictability issues are generally assessed in the context of parametric predictive regression models in which rates of returns are regressed against the lagged values of stochastic explanatory variables (or state variables). Mankiw and Shapiro (1986) and Stambaugh (1986) were first to discern the econometric and statistical difficulties inherent in the estimation of predictive regressions through the structure predictive linear model as

$$y_t = \alpha_0 + \alpha_1 x_{t-1} + \varepsilon_t, \quad x_t = \rho x_{t-1} + u_t, \quad 1 \leq t \leq n \quad (5.1)$$

where y_t is the predictable variable, say excess stock return at time t , innovations $\{(\varepsilon_t, u_t)\}$ are independently and identically distributed (iid) bivariate normal $N(0, \Sigma)$ with $\Sigma = \begin{pmatrix} \sigma_\varepsilon^2 & \sigma_{\varepsilon u} \\ \sigma_{\varepsilon u} & \sigma_u^2 \end{pmatrix}$, and x_{t-1} is a financial variable such as the log dividend-price ratio at time $t - 1$, which is commonly modelled by an AR(1) model as the second equation in (5.1).

There are several limitations to model (5.1) that should be seriously considered. First, note that the correlation between two innovations ε_t and u_t in (5.1) is $\phi = \sigma_{\varepsilon u} / \sigma_\varepsilon \sigma_u$, which is unfortunately non-zero for many empirical applications; see, for example, see Table 4 in Campbell and Yogo (2006) and Table 1 in Torous, Valkanov, and Yan (2004) for some real applications. This creates the so called “endogeneity” (x_{t-1} and ε_t may be correlated) problem which makes modelling difficult and produces estimation biased. Another difficulty comes from the parameter ρ , which is the unknown degree of persistence of the variable x_t . That is, x_t is stationary ($|\rho| < 1$); see Viceira (1997), Amihud and Hurvich (2004), Paye and Timmermann (2006), and Dangl and Halling (2007), or it is unit root or integrated ($\rho = 1$), denoted by I(1); see Park and Hanh (1999), Chang and Martinez-Chombo (2003), and Cai, Li and Park (2006), or it is local-to-unity or nearly integrated ($\rho = 1 + c/n$, where $c < 0$), denoted by NI(1); see, Elliott and Stock (1994), Cavanagh, Elliott, and Stock (1995), Torous, Valkanov, and Yan (2004), Campbell and Yogo (2006), Polk, Thompson, and Vuolteenho (2006), and Rossi (2007), and among others. This means that predictive variable x_t is highly persistent, and even nonstationary, which causes a lot of troubles for econometric modelling.

The third difficulty is the instability issue of the return predictive model. In fact, in return predictive models based on financial instruments such as the dividend and earnings yield, short interest rates, term spreads, and default premium, and so on, there have been found many evidences on the instability of prediction model, particularly based on the dividend and earnings yield and the sample from the second half of the 1990s. This leads to the conclusion that the coefficients should change over time; see, for example, Viceira (1997), Lettau and Ludvigsson (2001), Goyal and Welch (2003), Paye and Timmermann (2006), Ang and Bekaert (2007), and Dangl and Halling (2007). While the aforementioned studies found evidences of instability in return predictive models, they did not provide any guideline on how the coefficients change over the time and where the return models may have changed. It is well known that if return prediction models are unstable, one can only assess the economic significance of return predictability provided it can be determined how widespread such instability changes over time and the extent to which it affects the predictability of stock returns. Therefore, all of the foregoing difficulties about the classical predictive regression models motivate us to propose a new varying-coefficient predictive regression model. The

proposed model is not only interesting in its applications to finance and economics but also important in enriching the econometric theory.

As shown in Nelson and Kim (1993), because of the endogeneity, the ordinary least squares (OLS) estimate of the slope coefficient α_1 in (5.1) and its standard errors are substantially biased in finite samples if x_t is highly persistent, not really exogenous, and even nonstationary. Conventional tests based on standard t -statistics from OLS estimates tend to over-reject the null of non-predictability in Monte Carlo simulations. Some improvements have been developed recently, such as the first order bias-correction estimator proposed in Stambaugh (1999), the two-stage least squares (linear projection method) estimator in Amihud and Hurvich (2004), and the conservative bias-adjusted estimator in Lewellen (2004), but all of them still have not overcome the instability difficulty mentioned above. To deal with the instability problems, Paye and Timmermann (2006) analyze the data based on a sample of excess returns on international equity indices related to state variables such as the lagged dividend yield, short interest rate, term spread and default premium, to investigate how widespread the evidence of structural breaks is and what extent breaks affect the predictability of stock returns. Finally, Dangl and Halling (2007) consider equity return prediction model with random coefficients generated from a unit root process, related to 16 state variables.

Cai and Wang (2008a) consider a time-varying coefficient predictive regression model to allow the coefficients α_0 and α_1 in (5.1) to change over time (to be function of time), denoted by $\alpha_0(t)$ and $\alpha_1(t)$. They use a nonlinear projection of ε_t onto u_t and apply the local linear method to find the nonparametric estimates for $\alpha_j(t)$. They derive the asymptotic properties for the proposed estimator and it shows that the limiting distribution is a mixed normal with conditional variance being a function of integrations of an Ornstein-Uhlenbeck process (mean-reverting process). It also demonstrates that the convergence rates for the intercept function (the regular rate at $(nh)^{1/2}$) and the slope function (a faster rate at $(n^2h)^{1/2}$) are totally different due to the NI(1) property of the state variable, although the asymptotic bias, coming from the local linear approximation, is the same as the stationary covariate case. Therefore, to estimate the intercept function efficiently, Cai and Wang (2008a) propose a two-stage optimal estimation procedure similar to the profile likelihood method; see, e.g., Speckmann (1988), Cai (2002a, 2002b), and Cai, Li, and Park (2006), and they also show that the two-stage estimator reaches indeed the optimality.

Cai and Wang (2008b) consider some consistent nonparametric tests for testing the null hypothesis of whether a parametric linear regression model is suitable or if there is no a relationship between the dependent variable and predictors. Therefore, these testing problems

can be postulated as the following general testing hypothesis:

$$\mathbb{H}_0 : \alpha_j(t) = \alpha_j(t, \theta_j), \quad (5.2)$$

where $\alpha_j(t, \theta_j)$ is a known function with unknown parameter θ_j . If $\alpha_j(t, \theta_j)$ is constant and (5.2) is to test if the model (5.1) is appropriate for the given problem. If $\alpha_1(t, \theta_1) = 0$, it is to test if there exists the predictability. If $\alpha_j(t, \theta_j)$ is a piecewise constant function, it is to test whether there is any structural change. Cai and Wang (2008b) propose a nonparametric test which is a U-statistic type, similar to (4.13), and they also show that the proposed test statistic has different asymptotic behaviors depending on the stochastic properties of x_t . Specifically, Cai and Wang (2008b) address the following two scenarios: (a) x_t is non-stationary (either I(1) or NI(1)); (b) x_t contains both stationary and non-stationary components. Cai and Wang (2008a, 2008b) apply the estimation and testing procedures described above to consider the instability of predictability of some financial variables. Their test finds evidence for instability of predictability for the dividend-price and earnings-price ratios. They also find evidence for instability of predictability with the short rate and the long-short yield spread, for which the conventional test leads to valid inference.

For the linear projection used by Amihud and Hurvich (2004), it is implicitly assumed that the joint distribution of two innovations ε_t and u_t in the model (5.1) is normal and this assumption might not hold for all applications. To relax this harsh assumption, Cai (2008) considers a nonlinear project of ε_t onto x_{t-1} instead of u_t as $\varepsilon_t = \phi(x_{t-1}) + v_t$, where $E(v_t|x_{t-1}) = 0$. Therefore, the endogeneity is removed. Then, the model (5.1) becomes to the following classical regression model with nonstationary predictors

$$y_t = g(x_{t-1}) + v_t, \quad x_t = \rho x_{t-1} + u_t, \quad 1 \leq t \leq n \quad (5.3)$$

where $g(x_{t-1}) = \alpha_0 + \alpha_1 x_{t-1} + \phi(x_{t-1})$ and $E(v_t|x_{t-1}) = 0$. Now, for the model (5.3), the testing predictability $\mathbb{H}_0 : \beta_1 = 0$ for the model (5.1) as in Campbell and Yogo (2006) becomes the testing hypothesis $\mathbb{H}_0 : g(x) = c$ for the model (5.3), which is indeed more general. To estimate $g(\cdot)$ nonparametrically, Cai (2008) uses a local linear or local constant method and derives the limiting distribution of the nonparametric estimator when x_t is an I(1) process. Further, Cai (2008) proposes two test procedures. The first one is similar to the testing approach proposed in Sun, Cai and Li (2008) when x_t is integrated. And, the second one is to use the generalized likelihood ratio type testing procedure as in Cai, Fan and Yao (2000) and the bootstrap. Finally, Cai (2008) applies the aforementioned estimation and testing procedures to consider the predictability of some financial instruments. The test

finds evidence for the predictability for the dividend-price and earnings-price ratios, and the short rate and the long-short yield spread.

6 Conclusion

Over the last several years, nonparametric methods for both continuous and discrete time have become an integral part of research in financial economics. The literature is already vast and continues to grow swiftly, involving a full spread of participants for both financial economists and statisticians and engaging a wide sweep of academic journals. The field has left indelible mark on almost all core areas in finance such as asset pricing theory, consumption portfolio selection, derivatives and risk analysis. The popularity of this field is also witnessed by the fact that the graduate students at both Master and doctoral levels in economics, finance, mathematics and statistics are expected to take courses in this discipline or alike and review the important research papers in this area to search for their own research interests, particularly dissertation topics for doctoral students. On the other hand, this area also has made an impact in the financial industry as the sophisticated nonparametric techniques can be of practical assistance in the industry. We hope that this selective review has provided the reader a perspective on this important field in finance and statistics and some open research problems.

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