1 Introduction

Consider a high frequency time series with observed values $y_1, \ldots, y_{T+S}$, collectively denoted by $y_{1:T+S}$, with a forecaster interested in the forecast distribution of the future value y_{T+S+h}, where $S \geq 1$. In many applications forecasting uncertainty can be reduced by conditioning on additional data, so knowledge of the distribution of $p(y_{T+S+h}|y_{1:T+S})$ is preferable to that of $p(y_{T+S+h}|y_{1:T})$.

Many time series models of interest contain a set of global parameters $\boldsymbol{\theta}$ and observation specific latent variables $x_{0:T}$, such that x_t is assumed to follow a Markovian structure and y_t is conditionally independent of $y_{1:t-1}$, given $\boldsymbol{\theta}$ and x_t. One of the most common computational techniques for models with latent variables is Markov Chain Monte Carlo (MCMC), which involves iterative sampling from a Markov Chain designed to converge to the true posterior distribution, however these approaches can have slow convergence rates. In this paper it is assumed that an MCMC algorithm will require a time period of length S to converge, so at time $T + S$ the posterior distribution of $\boldsymbol{\theta}$ and $x_{0:T+S}$ can be conditioned only on $y_{1:T}$ and hence only $p(y_{1:T+S+h}|y_{1:T})$ is available.

The additional data $y_{T+1:T+S}$ can be included in the forecast by using filtering techniques such as the Kalman filter or particle filtering, however this will only update knowledge on $x_{0:T+S}$ but not $\boldsymbol{\theta}$. As an alternative to this method, this paper will instead approximate the posterior distribution $p(\boldsymbol{\theta}, x_T|y_{1:T})$ with $q_{\boldsymbol{\lambda}}(\boldsymbol{\theta}, x_T|y_{1:T})$ and the updated posterior distribution $p(\boldsymbol{\theta}, x_{T+S}|y_{1:T+S})$ with $q_{\boldsymbol{\lambda}}(\boldsymbol{\theta}, x_{T+S}|y_{1:T+S})$. The parameters of this approximation, denoted by $\boldsymbol{\lambda}$, are optimised by gradient descent to minimise the Kullback-Leibler divergence from q to p, a technique known as Variational Inference, which has seen wide use in the literature (references).

Variational Bayes involves an implicit trade-off: reducing the forecast horizon of y_{T+S+h} from $S + h$ to h reduces the uncertainty of the forecast, but the approximation will increase the statistical error. This paper additionally explores that trade-off for a Stochastic Volatility Model, a common latent state time series model.

- Some motivation about forecasting and updates, can probably be copied from the MIVB file
- Not focusing on MCMC sample
- VB estimate at time T, VB update to time T+S
- End goal: An approximation $q(\theta, x_T)$ for use in forecasting.

2 MCMC and Particle Filtering

- MCMC background

- Stochastic Volatility Model
- PMCMC
- Data driven particle filters?
- Whatever else

(Arulampalam et al., 2002)

3 Variational Inference

Variational Inference posits a divergence function between the true posterior distribution $p(\boldsymbol{\theta}, x_{0:T}|y_{1:T})$ and some approximating distribution $q_{\boldsymbol{\lambda}}(\boldsymbol{\theta}, x_{0:T}|y_{1:T})$, choosing the parameters $\boldsymbol{\lambda}$ for a given functional form q that minimises the divergence function.
This paper will follow the traditional approach, where the divergence function is the Kullback-Leibler (KL) divergence (Kullback and Leibler, 1951) from $q_{\boldsymbol{\lambda}}(\boldsymbol{\theta}, x_{0:T}|y_{1:T})$ to the true posterior $p(\boldsymbol{\theta}, x_{0:T}|y_{1:T})$ The KL divergence is defined by

$$KL[q_{\boldsymbol{\lambda}}(\boldsymbol{\theta}, x_{0:T}|y_{1:T}) \,||\, p(\boldsymbol{\theta}, x_{0:T}|y_{1:T})] = \int q_{\boldsymbol{\lambda}}(\boldsymbol{\theta}, x_{0:T}|y_{1:T}) \ln\left(\frac{q_{\boldsymbol{\lambda}}(\boldsymbol{\theta}, x_{0:T}|y_{1:T})}{p(\boldsymbol{\theta}, x_{0:T}|y_{1:T})}\right) d\boldsymbol{\theta} dx_{0:T} \tag{3.1}$$

and can alternatively be expressed as

$$KL[q(\boldsymbol{\theta}, x_{0:T}|\boldsymbol{\lambda}) \,||\, p(\boldsymbol{\theta}, x_{0:T}|y_{1:T})] = \ln(p(y_{1:T})) - \mathcal{L}(\boldsymbol{\lambda}) \tag{3.2}$$

where $\mathcal{L}(\boldsymbol{\lambda})$ is referred to as the Evidence Lower Bound (ELBO), as it provides a lower bound on the unknown constant $\ln(p(y_{1:T}))$. $\mathcal{L}(\boldsymbol{\lambda})$ is defined by

$$\mathcal{L}(\boldsymbol{\lambda}) = \int q_{\boldsymbol{\lambda}}(\boldsymbol{\theta}, x_{0:T}|y_{1:T}) \ln\left(\frac{p(y_{1:T}, \boldsymbol{\theta}, x_{0:T})}{q_{\boldsymbol{\lambda}}(\boldsymbol{\theta}, x_{0:T}|y_{1:T})}\right) d\boldsymbol{\theta} dx_{1:T}, \tag{3.3}$$

and as $\ln(p(y_{1:T}))$ is constant with respect to $\boldsymbol{\lambda}$, maximising (3.3) with respect to $\boldsymbol{\lambda}$ is equivalent to minimising (3.1). Maximising (3.3) is more convenient that minimising (3.1) as it is a function of the known joint distribution $p(y_{1:T}, \boldsymbol{\theta}, x_{0:T})$ instead of the unknown posterior $p(\boldsymbol{\theta}, x_{0:T}|y_{1:T})$.

3.1 ELBO Optimisation

Equation (3.3) can be maximised using a gradient ascent approach, where the following update step is iteratively applied until (3.3) converges within some pre-specified tolerance:

$$\boldsymbol{\lambda}^{(m+1)} = \boldsymbol{\lambda}^{(m)} + \rho^{(m)} \frac{\delta}{\delta \boldsymbol{\lambda}} \mathcal{L}(\boldsymbol{\lambda}^{(\boldsymbol{m})}), \tag{3.4}$$

where the derivative is evaluated at $\boldsymbol{\lambda}^{(m)}$. This update requires some initial values $\boldsymbol{\lambda}^{(0)}$ and a sequence $\rho^{(m)}, m = 1, 2, \ldots$ known as the learning rate. If $\rho^{(m)}$ is chosen to satisfy the following conditions the algorithm is guaranteed to converge to a local maximum (Robbins and Monro, 1951).

$$\sum_{m=1}^{\infty} \rho^{(m)} = \infty \tag{3.5}$$

$$\sum_{m=1}^{\infty} (\rho^{(m)})^2 < \infty. \tag{3.6}$$

This paper uses Adam (Kingma and Ba, 2015) to generate the sequence $\rho^{(m)}$.
Ranganath et al. (2014) showed that a Monte Carlo estimate of the derivative of the ELBO can be given by

$$\frac{\delta}{\delta \boldsymbol{\lambda}} \mathcal{L}(\boldsymbol{\lambda}^{(\boldsymbol{m})}) \approx \frac{1}{N} \sum_{i=1}^{N} \frac{\delta}{\delta \boldsymbol{\lambda}} [\ln(q_{\boldsymbol{\lambda}^{(m)}}(\boldsymbol{\theta_i}, x_{0:T,i}|y_{1:T})] \ln \left(\frac{p(y_{1:T}, \boldsymbol{\theta}_i, x_{1:T,i})}{q_{\boldsymbol{\lambda}^{(m)}}(\boldsymbol{\theta_i}, x_{0:T,i}|y_{1:T})} \right) \tag{3.7}$$

where $i = 1, \ldots, N$ indicates independent simulations from $q_{\boldsymbol{\lambda}^{(m)}}(\boldsymbol{\theta}, x_{0:T}|y_{1:T})$. The terms in the sum in (3.7) can have large variances, and in practice a large value of N is required to ensure a precise estimate of the gradient of the ELBO is obtained, slowing computation.
The variance can be reduced by the reparameterisation trick of Kingma and Welling (2014), introducing a random vector $\boldsymbol{\epsilon}$ with a distribution $q(\boldsymbol{\epsilon})$ that contains no free parameters, and a differentiable transform f such that

$$f(\boldsymbol{\epsilon}, \boldsymbol{\lambda}) = \{\boldsymbol{\theta}, x_{1:T}\}. \tag{3.8}$$

Kingma and Welling (2014) show that an f exists to transform $\boldsymbol{\epsilon}$ to any continuous random variable, with examples including a location-scale transform of a standard normal ϵ and an inverse-CDF transform of a uniform ϵ. In this case, (3.3) becomes

$$\mathcal{L}(\boldsymbol{\lambda}) = \int q(\boldsymbol{\epsilon}) \ln \left(\frac{p(y_{1:T}, f(\boldsymbol{\epsilon}, \boldsymbol{\lambda}) |\det J(\boldsymbol{\epsilon}, \boldsymbol{\lambda})|}{q(\boldsymbol{\epsilon})} \right) d\boldsymbol{\epsilon}, \tag{3.9}$$

where $J(\boldsymbol{\epsilon}, \boldsymbol{\lambda})$ is the Jacobian matrix of the transformation f. The derivative of (3.9) can be estimated by

$$\frac{\delta}{\delta \boldsymbol{\lambda}} \mathcal{L}(\boldsymbol{\lambda}^{(m)}) \approx \frac{1}{M} \sum_{i=1}^{M} \frac{\delta f(\boldsymbol{\epsilon}_i, \boldsymbol{\lambda}^{(m)})}{\delta \boldsymbol{\lambda}} \frac{\delta \ln(p(y_{1:T}, f(\boldsymbol{\epsilon}_i, \boldsymbol{\lambda}^{(m)})))}{\delta f(\boldsymbol{\epsilon}_i, \boldsymbol{\lambda}^{(m)})} + \frac{\delta \ln(|\det J(\boldsymbol{\lambda}^{(m)}, \boldsymbol{\epsilon}_i)|)}{\delta \boldsymbol{\lambda}}, \tag{3.10}$$

where simulations of $\boldsymbol{\theta}$ and $x_{1:T+S}$ are replaced by simulations of $\boldsymbol{\epsilon}$ from $q(\boldsymbol{\epsilon})$. The variance of this estimator is often orders of magnitude smaller than the estimator in (3.7) (citations), so M can be set much lower than N.

3.2 Randomised Quasi Monte Carlo

(Gunawan et al., 2017) (Matousek, 1998) (Bratley and Bennet, 1988) (Sobol, 1967)

3.3 Choice of q distribution

4 Dimensionality Reduction

ADVI simulation results
As the latent variables are Markovian, once the distribution of x_{T+1} is available, the distributions of each previous $x_t, t \leq T$ is irrelevant to forecasts of y_{T+S+h} and it may be more convenient to construct an approximation for $p(\boldsymbol{\theta}, x_{T+1}|y_{1:T})$ than for $p(\boldsymbol{\theta}, x_{0:T+1}|y_{1:T})$, as they contain the same information about y_{T+S+h} but the dimensionality of the posterior is reduced from $T+k+2$ to $k+1$, where k is the number of elements in $\boldsymbol{\theta}$. This in turn allows for more expressive approximating distributions to be used, as the size of $\boldsymbol{\lambda}$ does not grow with T.
In this case, using $f(\boldsymbol{\epsilon}, \boldsymbol{\lambda}) = \{\boldsymbol{\theta}, x_{T+1}\}$,

$$\ln(p(y_{1:T}, f(\boldsymbol{\epsilon}, \boldsymbol{\lambda}))) = \ln\left(\int_{X_T} p(x_{T+1}|\theta, x_T)p(x_T|y_{1:T}, \boldsymbol{\theta})d_{X_T}\right) + \ln(p(y_{1:T}|\boldsymbol{\theta})) + \ln(p(\boldsymbol{\theta})) \tag{4.1}$$

In many cases, the distributions $p(x_T|y_{1:T}, \boldsymbol{\theta})$ and $p(y_{1:T}|\boldsymbol{\theta})$ are intractable, but Tran et al. (2017) note that the particle filter estimators $\hat{p}(x_T|y_{1:T}, \boldsymbol{\theta})$ and $\hat{p}(y_{1:T}|\boldsymbol{\theta})$ are sufficient substitutes in what they refer to as Variational Bayes with Intractable Likelihood (VBIL). The particle filter estimation of the distribution $\hat{p}(x_T|y_{1:T}, \boldsymbol{\theta})$ is expressed as a discrete set of point masses $x_T^{(k)}$ and weights $\pi_T^{(k)}$, where $k = 1, \ldots, N$ represent the particles of the particle filter. In this case, the integral in (4.1) reduces to the sum

$$\int_{x_T} p(x_{T+1}|\theta, x_T)p(x_T|y_{1:T}, \boldsymbol{\theta})d_{x_T} = \sum_{k=1}^{N} \pi_T^{(k)} p(x_{T+1}|\boldsymbol{\theta}, x_T^{(k)}), \tag{4.2}$$

and hence

$$\ln(p(y_{1:T}, f(\boldsymbol{\epsilon}, \boldsymbol{\lambda}))) \approx \ln\left(\sum_{k=1}^{N} \pi_T^{(k)} p(x_{T+1}|\boldsymbol{\theta}, x_T^{(k)})\right) + \ln(\hat{p}(y_{1:T}|\boldsymbol{\theta})) + \ln(p(\boldsymbol{\theta})) \tag{4.3}$$

can be substituted into (3.10). This derivative of (4.3) with respect to $\boldsymbol{\lambda}$ can be obtained with automatic differentiation tools such as those provided by Stan (Carpenter et al., 2015).

4.1 Importance Sampling

(Sakaya and Klami, 2017)

5 Updating

$$p(y_{1:T+S}, \theta, x_T, x_{T+S}) = p(\theta)p(y_{1:T+S}|\theta)p(x_T|\theta, y_{1:T+S})p(x_{T+S}|x_T, \theta, y_{1:T+S})$$

- Proper updates with filter/smooth
- Secondary approximation - no smoothing
- Straight up particle filter on old approximation
- PF on old MCMC
- Simulation Results

6 Empirical

7 Discussion

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