

How would an historian do time-series econometrics? Analogy weighting using machine learning

Jim Savage

12 October 2015

Abstract

This paper introduces Analogy Weighting, a weighting scheme in which the most relevant historical episodes disproportionately influence parameter estimates in time series models. The algorithm makes use of the proximity matrix of a Random Forest to generate analogy weights. Consequently the weights are both supervised (the weights are measures of historical closeness in terms of the variables that matter to the prediction of the dependent variable), and metric-free. Using analogy weights improves the predictive performance of forecasting models. Also, analogy weights allow the researcher to acknowledge when the present state of the economy is unlike the historical data on which the model is estimated. During these phases, researchers may wish to allow ignorance priors to take over, blowing up prediction intervals. Applications are given to measuring core inflation and portfolio volatility.

Introduction¹

In building reduced-form empirical macroeconomic models for policy simulation and forecasting *today*, how much history should be included? Too much history and the analyst risks estimating parameters that describe history well but have little relevance to today. With too little history, estimates risk missing out on useful information. Implicit in doing time-series work is an assumption that history provides a useful guide for parameterising economic models.

The method presented in this paper takes this one step further: I assert that if history provides a useful guide to reduced-form models, then more analogous histories should provide a better guide. Conversely, we should have little confidence in model projections being made if the current state of the world is different from the world from which the model parameters were estimated.

How should we decide which histories are most relevant to today? The techniques most commonly used to address this problem aim to account for unmodelled structural change in the underlying data generating process. They come in varying degrees of complexity.

A naive or un-principled approach is to assume that co-movements between variables over the recent history are likely to emerge from a structural relationship that is relevant for the near future. Making such an assumption motivates techniques often used in industry and government work (Alloway 2013), such as estimating models over a rolling window, or using exponential smoothing to give less weight to more distant historical observations. These techniques are easy to understand and implement, but it is not clear what an optimal window selection rule should be.

A second class of strategies includes all historical data, but explicitly models the time variation in model parameters, either using regime-switching frameworks or continuous-state approaches. These approaches are more principled, in that the parameters and data come

¹I would like to thank my advisor, Chris Edmond, for his input. Errors and conceptual shortcomings are all my own. Also thanks to Sue Marquez, Siddhartha Chib, Ben Goodrich, Bob Carpenter, Joshua Chan, Jim Minifie, Eugene Dubossarsky, Jack Mullumby, David Stephan, David Miller and Sarah Tan for their input at various stages in the project.

from a generative model (requiring that the researcher be explicit in their beliefs). Yet they are lacking on other dimensions. The models are typically difficult to implement, are richly parameterised (necessitating informative priors), and, often, the generative models themselves give non-zero probability to implausible outcomes, such as explosive states.

How would historians build econometric models? The ‘historianist’ approach proposed in this paper is to analogise today with periods of history that have similar characteristics, and draw inference based on those analogies. There is a large problem in applying this historianist approach formally to time-series models: how can we determine which time-periods are “similar” to today? Estimating similarity typically involves selecting a metric with which we can measure the similarity between periods. But how much weight should be given to the various dimensions? And should these weights change over time? Perhaps one variable is very important in some states of the world and not in others.

This paper describes an algorithm that makes use of the Random Forest (Breiman 2001), a popular method from machine learning, to construct “analogy weights”. These weights can be used to help regression models learn more from similar periods of history. In the applications below, analogy weighted models outperform benchmark models in out-of-sample predictive comparisons. Yet the real advantage of using analogy weights is to inform the user when the world is in an unprecedented state. When used with Bayesian regression models with informative priors on (large) residual variances, this algorithm can help increase uncertainty when the model is being asked to generate projections in a world that is not similar to the world used to estimate the model.

The weights have two important features. First, their construction does not make use of a metric; a variable’s importance to the weight can vary with states of the world. Second, there are automated out-of-sample robustness checks in the construction of the weights, and so weights represent only similarities on dimensions that appear to have a robust relationship to the dependent variable. The algorithm also requires very few tuning parameters, making it a useful turn-key solution to many time-series modelling tasks.

The paper also presents two applications of the algorithm. The first is in estimating “core” inflation—the permanent component of inflation, using a Vector Autoregression (VAR) with

analogy weights. The second is an application to portfolio risk management, in particular estimating a conditional volatility model for a small portfolio of stocks during the Global Financial Crisis. This second application illustrates how researchers can easily incorporate analogy weights into multivariate GARCH models (below, my own augmentation of the CCC-GARCH method of (Bollerslev 1990)). The applications show that using analogy scores generates the greatest benefit in the first few periods following an unprecedented regime change.

1: Historical analogies in the literature

The premise of analogy weighting is that periods of history that are similar to today may provide more information about the structure of contemporaneous economic relationships than less similar periods of history. This is not a new revelation—many theoretical models allow similar states of the world to emerge. Consequently, many empirical models explicitly model regime shifts. It is also not a revelation only relevant to econometrics. In the finance, machine learning and signal processing literature, similarities between time series are often used in predictive tasks.

Why may similar periods be more relevant to today than dissimilar periods? This is a largely theoretical question left for further research, yet the literature contains many appealing models that give some support to the notion that the behaviour of the economy may depend on the values of macroeconomic and market variables, particularly during crises. The most prominent example of this is in Kiyotaki and Moore (1995) in which exogenous technology shocks reduce the economic value of capital and so affect asset prices. Faced with binding collateral constraints, firms cease to be able to finance investment, and so reduce their asset purchases; this reduces asset prices and amplifies the cycle. Bianchi (2010) proposes a similar mechanism, in which borrowers of foreign currency face similar collateral constraints when their home currency depreciates against the creditor's, reducing the availability of foreign credit and thus making exchange rate depreciations contractionary. In Brunnermeier and Sannikov (2012), “expert” firms can absorb small shocks by adjusting payouts and re-building

capital, but under large or serially correlated shocks these firms may be forced to stop investing or liquidate their capital. Consequently, the impact of a shock in their model depends on the sign and magnitude of the shock, and the economy's distance from its steady state.

While the channels by which these models transmit shocks vary, their implications for reduced-form modelling are similar: the behaviour of economic agents probably varies across states of the world, and there are potentially many such states.

Regime-switching models as in Hamilton (1988) and as reviewed in Kuan (2002) and Hamilton and Raj (2013) attempt to model changes in the state of the economy by allowing parameters to vary across a discrete number of regimes. These models have been applied in a variety of settings, such as examining fiscal multipliers in recessions vs expansions (Auerbach and Gorodnichenko 2012) and the permanent and transitory impacts of business cycles across different sub-periods of the business cycle (Kim and Murray 2002). Another interesting application to currency volatility is in Klaassen (2002). In that paper, standard GARCH forecasts of volatility tend to be too high during high-volatility periods; this is because long periods of relative calm result in a GARCH model's persistence parameter being very high. Allowing GARCH parameters to vary across states results in better out-of-sample volatility forecasts.

The basic structure adopted by the regime-switching literature is to allow for some model to depend on parameters that vary across $s \in S$ different states: $y = f(\theta_s, x)$. The states themselves are not known, but need to be inferred based on observations of y . Of central interest is the probability of transitioning from state j to state k $p(s_t = k | s_{t-1} = j, \dots, s_1 = l, y_{t-1}, \dots, y_1)$. If this probability is equivalent to $p(s_t = k | s_{t-1} = j)$, then the state transitions according to a Markov Chain.

Regime switching models are clearly a powerful approach at capturing un-modelled variation across time, especially if the differences between parameters' values in one state and another are large (this scenario may be difficult to model in the continuous state approach below). Yet the main drawback is that the approach relies on all regimes having been observed in the historical series. This limitation reveals itself in overfitting (in which out of sample

performance is significantly worse than in-sample performance). This anecdote, from Fabozzi and Mann (2012) illustrates the point well:

A survey participant at a firm where regime-shifting models have been experimented with commented,

Everyone knows that returns are conditioned by market regimes, but the potential for overfitting when implementing regime-switching models is great. If you could go back with fifty years of data—but we have only some ten years of data and this is not enough to build a decent model.

As the number of potential states of the economy grows, it may make more sense to use continuous-state models, as described in Koop and Korobilis (2010). In this approach, rather than having a set of parameters for each state of the economy, the “state” of the economy *is* the set of parameters; these are allowed to vary over time. A common implementation of these models, as in Cogley and Sargent (2002), Cogley and Sargent (2005), Koop, Leon-Gonzalez, and Strachan (2009) and Koop and Korobilis (2013) models parameters as changing over time in a random walk process, rather than modelling parameter values as responding directly to changes in exogenous variables. These papers implement some version of the following model. Taking some vector of variables y , they estimate the state space vector autoregression

$$y_t = Z_t\beta_t + \epsilon_t$$

Where Z_t contains an intercept and p lags of the dependent variables. The state evolves according to a random walk

$$\beta_t = \beta_{t-1} + \eta_t$$

and the innovations to both the VAR and state processes are multivariate normal. In some implementations (such as Cogley and Sargent’s 2005 model) the covariance of VAR innovations is itself time-varying, in which case these form a part of the evolving state.

$$\epsilon \sim \mathcal{MVN}(0, \Sigma_{(t)})$$

$$\eta \sim \mathcal{MVN}(0, Q)$$

Given priors on β_0 , Σ and Q , the papers above estimate the time-varying β using a Kalman Filter. One modification (as in Koop and Korobilis (2013)) is to include in the filter a “forgetting factor”—a parameter that determines how much past information influences the current estimate of β_t . The implicit assumption made when using this approach is that the distant past is irrelevant to parameter estimates today.

Both methods described above—regime shifting models and continuous-state-space models—offer some real benefits to a researcher modelling unobserved structural changes. Yet the models have drawbacks. Like typical reduced-form models, discrete regime-switching models are unable to generate states that have not been observed. If the true data generating process has not revealed all her states, these methods are prone to inductive errors. In contrast, continuous-state models give positive weight to the possibility of as-yet unobserved states. But they are hard to implement, and tend to be very richly parameterised. A large number of parameters relative to the number of observations requires that informative priors must be used to estimate continuous-state models (see Koop and Korobilis, 2009).

These problems are not unique to econometrics. In the machine learning and signal processing literature, “time-series-clustering” techniques aim to group time-series together into clusters, each of which should contain time-series that behave similarly. When researchers are looking for analogous periods of a given time-series, they use “subsequence clustering”, in which a rolling window generates many time-series which can be clustered.

One approach, summarised in Keogh and Lin (2005), has been to compute the values of a point-wise distance (usually a L1 or L2 norm) between all adjacent points in two time-series, and summarise the “similarity” between the two series with the average value of the distance. Time series are then clustered using K-means or hierarchical clustering. This approach is extremely computationally expensive when the length of time-series is large, and impossible

when there are missing values or time-series are different lengths. In Keogh and Lin (2005), this method is shown in many cases to generate “meaningless” clusters—clusters of time series that perform for prediction no better than random groupings.

Another approach has been to summarise time-series in terms of their “characteristics”, and then cluster on these. This is less computationally draining, and allows the researcher to cluster series that have different lengths, missing data, etc. In Deng, Moore, and Nechyba (1997), each series is characterised by the parameters of an ARMA model fitted to it, and series are clustered based on these parameter estimates. In Wang, Smith, and Hyndman (2006), each time-series is summarised according to characteristics including frequency, serial correlation, trend, skewness, kurtosis, seasonality, etc., and series are clustered based on the values of these characteristics. This approach to time-series clustering has been more successful at generating useful clusters.

There are a few aspects of time-series clustering approaches that make them unappealing for applied time-series economics research. First, clustering methods are used to cluster *univariate* time series with other univariate time series. The economic approaches discussed above aim instead to generate similar *histories*, which may be defined by the values of many time series. Second, while clustering may make sense if univariate time series behave in a deterministic manner, many interesting economic time series (including exchange rates, stock prices and consumption) look like random-walk series in the short-run. As in Keogh and Lim (2003), clustering random walk series predictably results in spurious clusters. Finally, most applied economic work is concerned with causal analysis, not prediction. Even if building univariate clusters was able to improve prediction of economic time series, it is not clear why this should help in building causal models.

How this paper relates to the literature

The algorithm described below aims to overcome the difficulties associated with both regime-switching models and continuous-state models, while retaining some of the advantages of both. Like regime-switching, the model allows large jumps in the values of parameters as the state shifts. Like a continuous state model, parameters are continuous across states of the

world (including as-yet unseen states). The big advantage is the ease of application. While the existing methods require a great deal of expertise, time and highly informative priors, analogy weighting can be as simple as using (analogy) weighted least squares. The cost of the approach is that we move away from a generative model.

The method works by building a weighting vector that relates one period in history (normally today) with all other periods in observed history. This score serves as a “distance between histories”, and is defined on the characteristics of the economy or system at that point in time.

To avoid having to define a metric that measures similarity between periods of history, the algorithm proposed makes use of the Random Forest (due to Breiman (2001)), a tool from machine learning. A byproduct of fitting a Random Forest is a proximity matrix, which for all pairwise observations provides a measure of proximity in the Random Forest space. The main benefit of this tool is that the distance measure is the distance between observations *in terms of the variables that matter to the prediction of the outcome variable, and the importance of each variable can vary over time*. The algorithm is a time-series generalisation of the approach taken by Miller, Savage, and Tan (Forthcoming), in which Random Forest proximities are used to construct a matching estimator that is less sensitive to the J. Smith and Todd (2001) critique than is Propensity Score Matching.

2 The Analogy Weighting algorithm

The algorithm described below uses the proximity matrix of a random forest. This chapter describes how the Random Forest works, starting by explaining the Classification and Regression Tree method, and then describes how the proximity matrix is estimated. Readers with a familiarity with these methods may skip to the end of the chapter.

Intermediate input 1: Regression Trees

Classification and Regression Trees, known as CART and due to Breiman et al. (1984), is a non-linear supervised recursive partitioning method. While this technique is not widely used in economics, it has recently found interest among some prominent researchers, such as Athey and Imbens (2015), Wager and Athey (2015) and Varian (2014). The aim of the technique is to partition a dataset according to the values of its covariates such that in each of the sub-groups the values of the dependent variable are more similar than in the parent group. The routine described below is a simplified version of the CART routine implemented in many software packages, though one that is much closer to the trees that make up a Random Forest.

Formally, let $x_{i,p}$ be the $p \in P$ -th covariate for individual $i \in I$. Although in some implementations of CART x can be a categorical variable, we consider only the case with a numerical x . The collection of all x s is X . Each individual has a dependent variable $y_i \in Y$, which again we take to be numerical.

Each node of the tree—a juncture between branches—is associated with a “split-point” \bar{x}_p defined as the point that splits the parent group into two child groups so that the sum of sum of squares for the child groups is reduced by as much as possible:

$$\bar{x}_p = \operatorname{argmax}_{x_p} \left\{ \sum_{i \in I} (y_i - \bar{y})^2 - \left(\sum_{i \in I_L} (y_i - \bar{y}_L)^2 + \sum_{i \in I_R} (y_i - \bar{y}_R)^2 \right) \right\}$$

Where I_L is the “left” part of the data; the individuals i for whom $x_{i,p} < \bar{x}_p$. The mean of those individuals’ dependent variable is \bar{y}_L . I_R is the remainder of the observations. These split points are simply values of the covariates that appear to be useful in terms of splitting the group into more similar sub-groups.

CART is a recursive algorithm, in that once a split has been estimated, the two sub-groups are then split, recursively, until further splits result in child groups groups that are no more similar than their parents, or until a minimum node size has been reached. In “pruned” tree, there is some cost attached to making further splits; if the gain in terms of decreased sum of squares does not meet some cost, then the split is not made. This is an efficient way of

decreasing over-fitting.

Once a tree is estimated, it can produce predicted values for a new observation *even if that observation is missing elements of X* . A new observation is simply “dropped” down the tree until it either reaches a terminal node (one with no child nodes) or until it reaches a node whose split is defined on a variable that is missing for the individual. Once it reaches this node, its predicted value for y is simply the average y for observations in that node in the training set.

If two observations reach the same terminal node, they are said to be *proximate*. Two proximate observations are similar in terms of all variables on which splitting decisions are made.

Intermediate input 2: Random Forests

While CART models are extremely easy to understand and visualise, they do tend to result in overfitting. They also do not capture smooth relationships well (there will only be as many distinct predicted values as there are terminal nodes).

Random Forests Breiman (2001) were designed to address these problems. They are essentially a collection of (normally thousands of) CART models, with two small twists. First, in a Random Forest, each tree is estimated on a random subset of the observations I . Second, at each node, only a proportion of all candidate x_p are eligible to be split on. All trees are grown to their full extent—that is, not pruned. This results in the trees that make up a Random Forest being quite distinct, there being no individual observation or x variable that has an enormous influence over all trees as a group.

To generate predicted values from a Random Forest, a new observation x_i is passed down all trees, and the fitted value is the average of the fitted values in all trees.

Intermediate input 3: Random Forest proximity matrices

This paper is not concerned with the (impressive, see Caruana and Niculescu-Mizil (2006)) predictive abilities of the Random Forest. Rather, the Random Forest generates a “proximity matrix” that gives a distance between all observations in the dataset.

When two observations fall into the same terminal node as one another, they are said to be *proximate*. The (i,j) th entry in a Random Forest’s proximity matrix is the proportion of terminal nodes that observations i and j share across all trees. For additional robustness, it is desirable to use only “out-of-bag” observations in the calculation of proximities. These are the observations excluded when each tree is grown. In this formulation of the proximity matrix, the proximity score is the proportion of terminal nodes shared by two observations *for the trees that the observations were out-of-bag*.

The proximity matrix is an attractive candidate for a distance measure between historical observations. It does not require defining a distance metric or estimating importance weights for the covariates. Because the Random Forest is supervised (it has a dependent variable), the proximities can be thought of as the distance between observations *in terms of the variables that matter for the prediction of y* . And because of the robust manner in which a Random Forest is estimated, (stationary) independent variables whose relationship with y is coincidental are less likely to influence the proximity score. Consequently, the analyst can afford to include many independent variables; those without a robust relationship will tend to be excluded.

Estimating analogy scores—the algorithm

This section describes the primary contribution of this paper: how to construct analogy scores. The aim of the following algorithm is to generate a vector corresponding to periods t_1, \dots, T that contains weights denoting the similarity between all historical periods and period T . Period T will typically be chosen to be the most recent period, as the model builder will generally want to build a useful model of the current state of the economy.

The first step is to construct a $T \times P$ matrix X that contains P variables that could be

though of as describing the “state” of the economy/market conditions. These variables should be stationary.

An aggregated proximity matrix is then constructed like so:

```

initialise pmat as a T by T matrix of 0s

for p in 1 to P:
    build Random Forest model of the p-th variable in
    X using all other variables (and possibly lags of
    all variables), call this RF
    pmat = pmat + proximity matrix of RF
end
pmat = 1/P * pmat

```

The T -th row of this matrix contains the proximities between period T and all periods beforehand. This vector can be then used for weighting observations in a regression model.

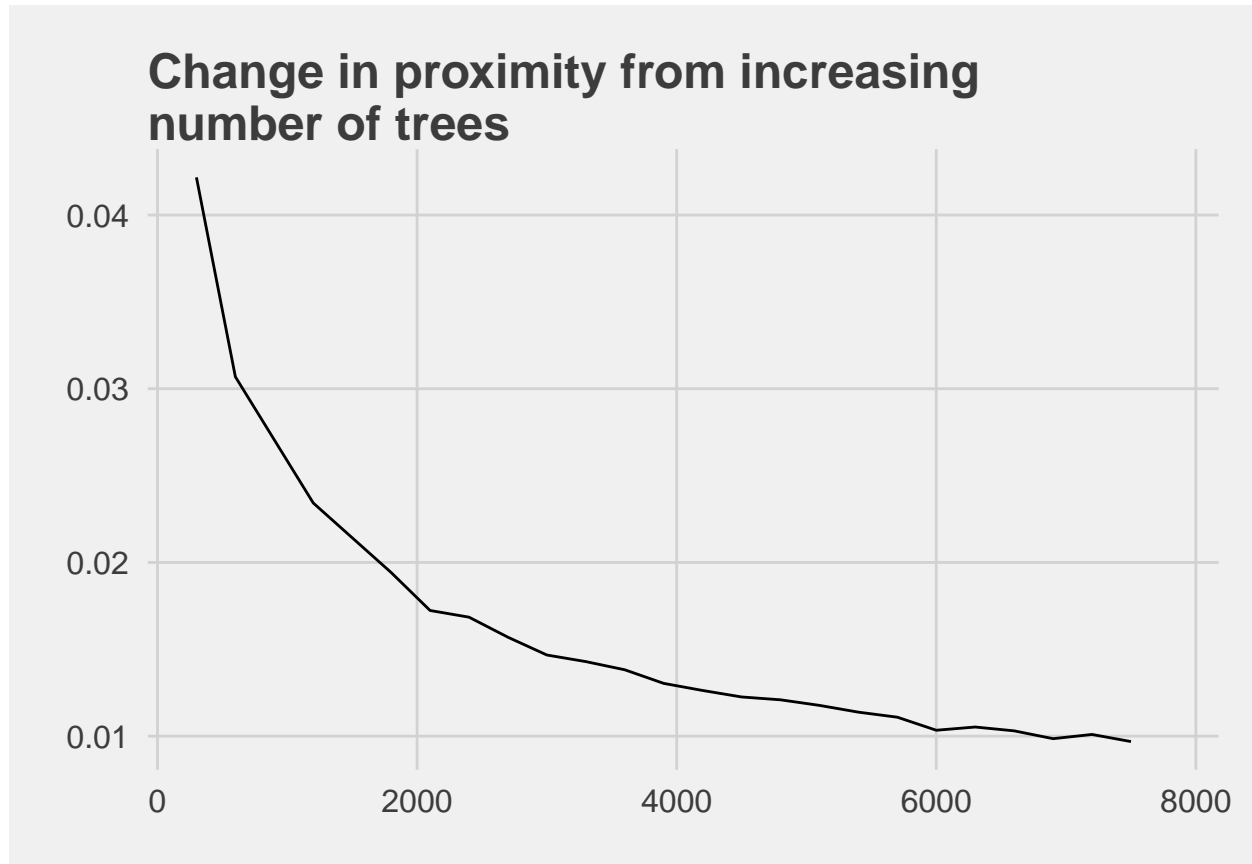
Tuning parameters

Given the introduction of randomness in estimating the random forests that we use to estimate analogy weights, how sure should we be that a given proximity is meaningful? That is, is there a sense of “significance” of the proximity estimate?

Two of the random forest’s tuning parameters have the greatest influence on proximity estimates. The first is `ntree`, the number of trees in the forest. A smaller number of trees relative to the number of observations is more likely to generate spurious matches. As the number of trees grows, the stability of proximity weights improves. This can be seen by taking the absolute value of the derivative of average proximity scores with respect to `ntree`. That is:

$$\text{instability score} = \text{abs} \left(\frac{\partial \frac{1}{n^2} \sum_i \sum_j \text{proximity}_{i,j}}{\partial \text{ntree}} \right)$$

This is illustrated in the figure below. As is clear, there are diminishing returns to the number of trees.



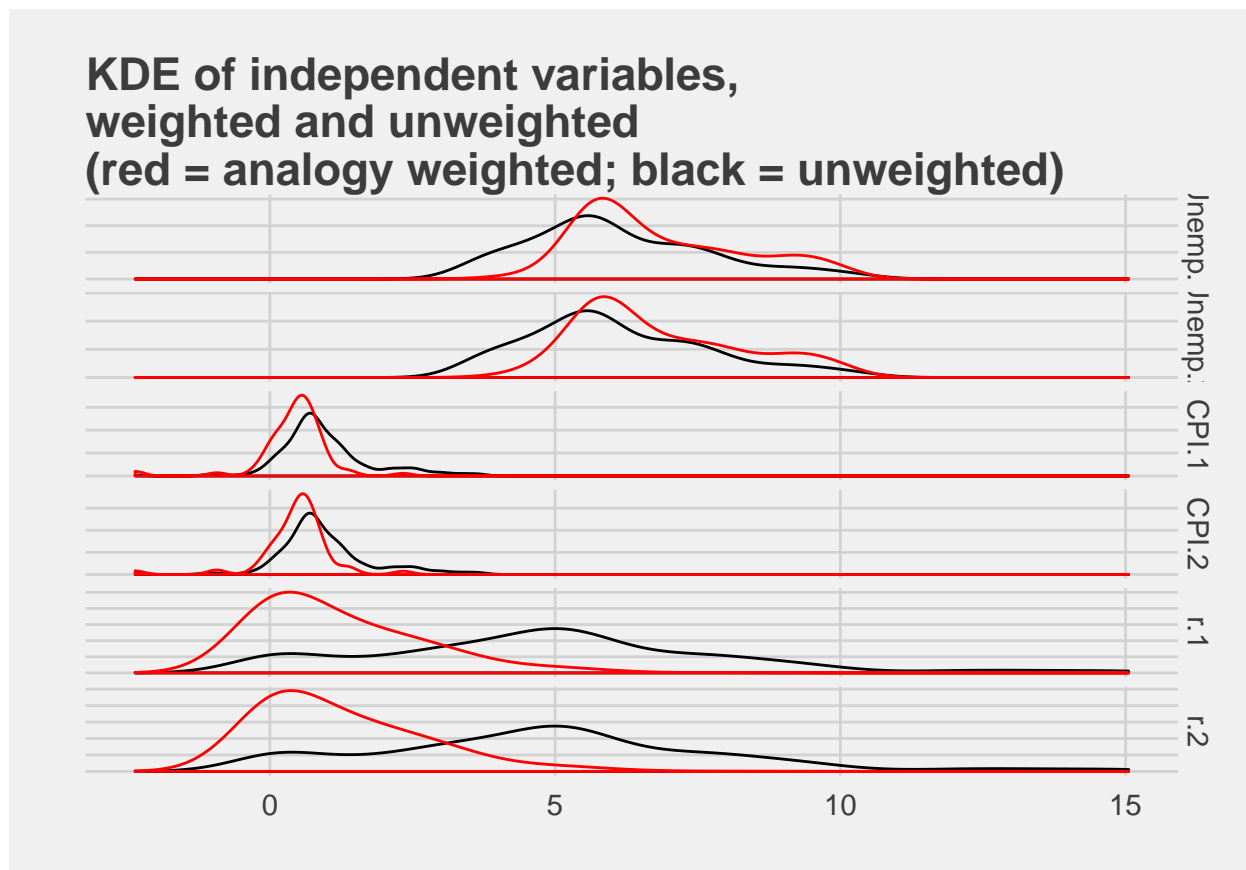
The second main tuning parameter that users should be aware of is the size of terminal nodes in each tree in the random forest, `nodesize`. The default is that trees should grow until a split would not decrease the loss, or until a split would result in a terminal node containing fewer observations than the minimum. The default node size value in R's implementation, `randomForest`, is 5. Increasing this size results in greater average proximity between observations, as the probability that any two observations share a terminal node is greater.

With a single terminal node in each tree, all observations will be trivially proximate, and analogy weights will be 1. This of course would defeat the purpose of using analogy weights! If trees are grown to a full extent (the minimum terminal node size), then a lower bound on the average weight will be the $\frac{5}{\text{number of obs}}$. In practice, the average node size will be considerably larger than the minimum node size. This is because candidate splits must both

decrease terminal-node variance *and* result in two terminal nodes that exceed the minimum terminal node size.

Choosing the optimal terminal node size will involve making a trade-off between including potentially irrelevant (or misleading) histories for larger values of `nodesize`, and having more uncertainty in estimates for smaller values of `nodesize`. The choice should be influenced by the number of parameters in the economic model being estimated, so that on average there are a sufficient number of effective observations (ie. the average weight multiplied by the total number of observations) to provide guidance for parameter values.

This calculation is not straightforward; using analogy weighting can both increase and decrease the precision of parameter estimates. Because analogy scores discount dissimilar observations, they necessarily compress the support on which the parameters are estimated (see the figure below). Taken alone, this should increase uncertainty in the parameter estimates. But the entire notion of proximity weighting is that irrelevant histories are just that—and co-movements during irrelevant periods simply confound our understanding of the true relationship today. If the co-movements between dependent and independent variables during analogous histories are less variable, then the precision of parameter estimates may *increase* under analogy weighting. Out-of-sample forecasting testing is one method of evaluating the relevance of parameter estimates, and is used in the examples below.



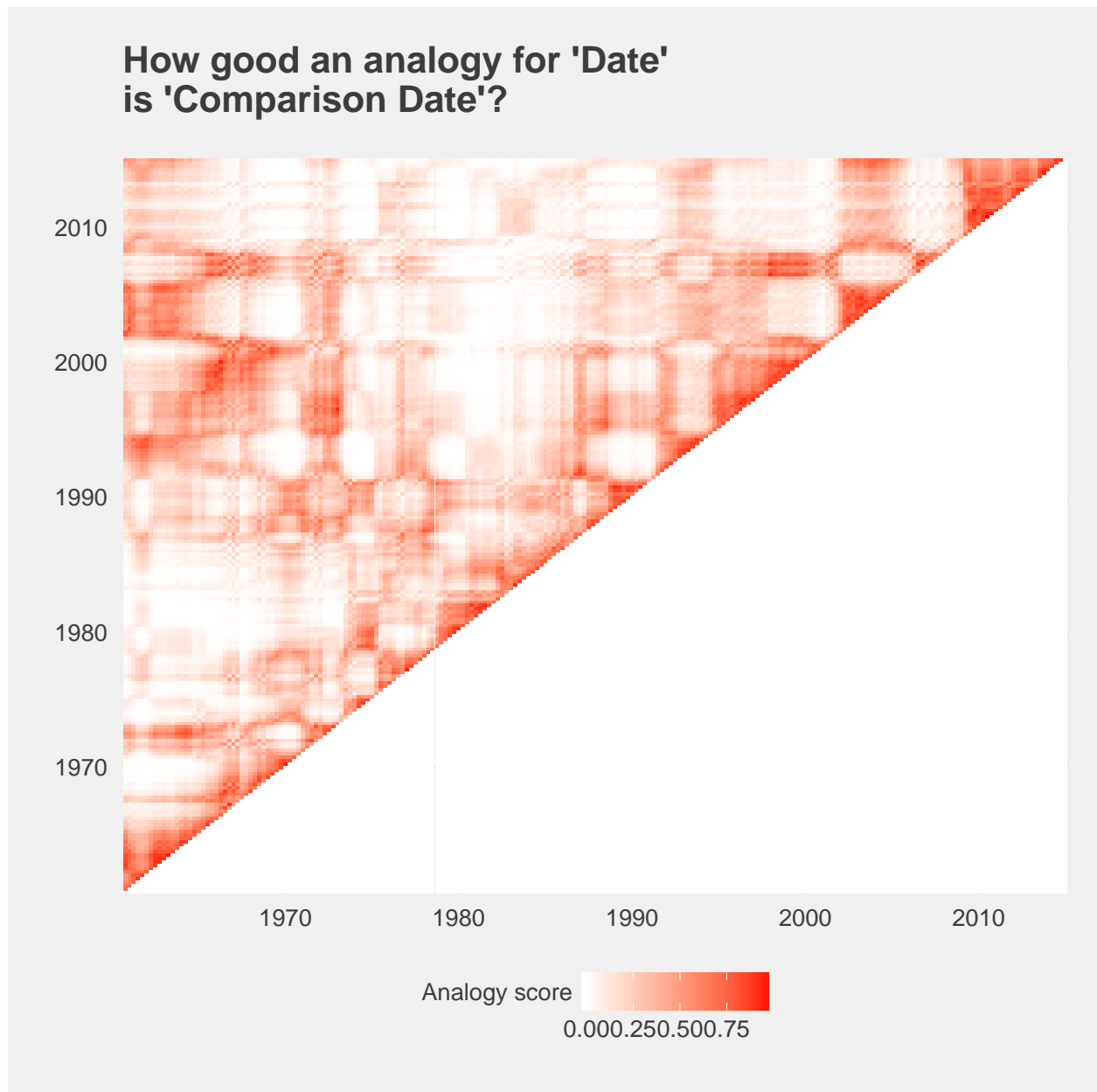
An example of estimating proximity weights

The application below illustrates how one could use analogy weights in estimating the parameters of a simple economic model, and discusses the consequences of modelling choices.

The procedure of building analogy weights discussed above was carried out on quarterly US data, specifically, the unemployment rate, the rate on short-term US Treasuries, and the quarter-on-quarter percentage change of the major cities consumer price index, taking T to be the first quarter of 2015. Random forests were trained on each of these variables, taking the first and second lags of itself and the other variables to be predictors, and the out-of-bag proximity matrices were saved. The terminal node size used was 40, about a fifth of the sample; the average proximity between points should be roughly a fifth.

The plot below shows the resulting proximity matrix. Along the vertical axis is the date for which we would like historical analogies. Along the horizontal axis is the period being

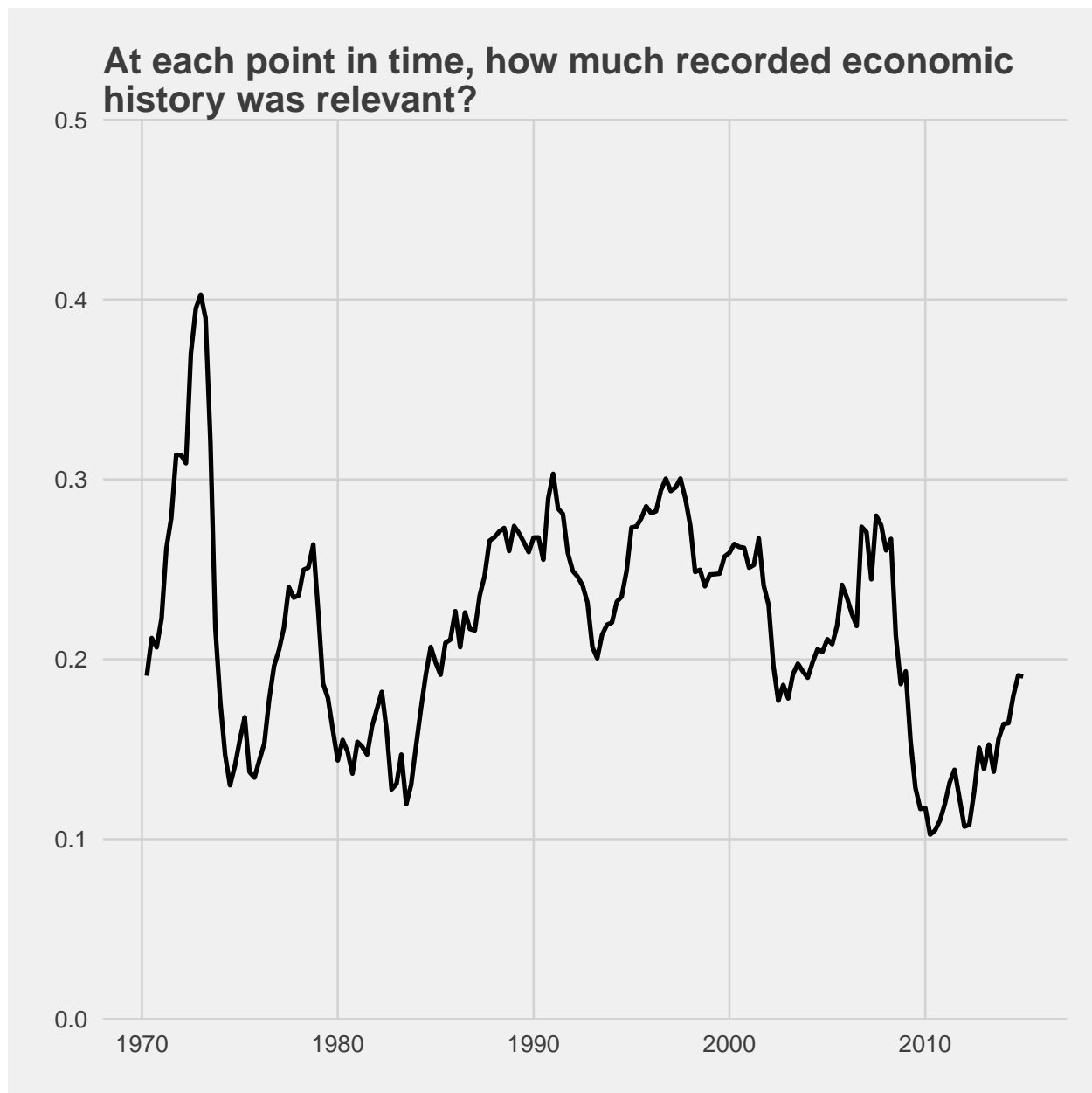
evaluated for similarity. The shading indicates the degree of similarity.



Interestingly, there appear to be three or four distinct states; the first being the post-GFC period in the top right. The second is a period of relative tranquility, roughly coinciding with the period between the early 1990s recession and the GFC, and the period before the stagflation of the 1970s. The third period, first making itself apparent in the late 1960s and continuing through until the early 1990s, is a less defined period. The emergence of fairly discrete states is interesting, given that there is no such restriction being placed on their discovery.

Another important feature of the proximity matrix is its ability to inform the user when the current state of the world is without precedent in terms of the variables that help to generate predictions. When the state of the world differs substantially to the sample, then perhaps the implications of models estimated from (a very different) history should be trusted less.

A metric to evaluate the similarity of the current state to recorded history is to divide the effective sample size (the sum of weights) by the naive sample size (the number of observations) up to a given period. This gives the effective proportion of history up to a period that would be included if you were to use the analogy weights. It is a crude measure for similarity to all of recorded history. It is illustrated in the figure below.



Note the three periods with the largest drop in relevant histories: the oil crisis of the 1970s, the recession of the early 1980s, and the Global Financial Crisis of the 2000s. Using analogy weights during these periods would involve learning less from history, and so being more unsure of model outputs. This feature is elaborated in the portfolio risk modelling example below.

3: Incorporating analogy weights into empirical models

This section gives an overview on how analogy weights may be used to help estimate parametric models. The method by which weights can be incorporated will depend on the sort of model being estimated.

In simple models estimated using maximum likelihood, weights are typically incorporated by scaling the likelihood contribution of each data point by the weight. That is, for some density function that depends on data y and parameters θ , $f(y; \theta)$, the maximum weighted likelihood estimate using a weight vector ω_i will be

$$\text{MLE}(\theta) = \text{argmax}_{\theta} \sum_i \omega_i \log f(y_i; \theta)$$

For historical data points that are given little weight, this has the effect of reducing their importance in the estimation of θ . If a data point is given zero weight, it has no impact on the estimation of the parameter. It is important to adjust the standard errors of parameters estimates when using weighted maximum likelihood. They will be affected by the weights in two ways: first, the variance-covariance matrix of θ will be a weighted variance covariance matrix, and second, there will be fewer effective observations. Analytically, the (weighted) Fisher Information matrix

$$I_{\omega}(\theta) = -\text{E} \left[\frac{\partial^2}{\partial \theta^2} \sum_i \omega_i \log f(y_i; \theta) \right]$$

helps to estimate the weighted covariance matrix

$$\Sigma_{\beta} \approx \frac{1}{\sum_i \omega_i} I_{\omega}(\theta)^{-1}$$

The Hessian matrix inside the square brackets of the first expression tell us about the curvature of the likelihood profile at the estimated point. If the likelihood is quite flat at the estimated θ , then there should be more uncertainty about the parameter estimate. In this case, entries in the Hessian will be small, and its inverse will be large, increasing standard

errors. If there is little weight placed on most historical values, then $\frac{1}{\sum_i \omega_i}$ will be large, increasing standard errors.

For more complex models estimated using Bayesian techniques, incorporating weights is slightly more involved. The below describes Hamiltonian Monte Carlo (the method used to estimate two models below) and how it can be augmented to include weighted observations.

Weighted Hamiltonian Monte Carlo in Stan

In Bayesian analysis, the aim is to conduct inference about a posterior distribution of parameters given data y and controls X $p(\theta|y, X)$. The posterior distribution is proportional to a prior distribution of the parameters multiplied by the likelihood of the data given the parameters $p(\theta|y, X) \propto p(\theta) \times p(y|\theta, X)$. The prior distribution is chosen by the researcher and incorporates information known before observing the data, such as previous study results, hard constraints on the parameter value, or an identifying belief that the parameter value is close to zero. To make inference about a parameter, the posterior needs to be integrated. For example, the mean of some parameter β would be $\hat{\beta} = \int_{\beta} \beta \times p(\beta|y) d\beta$. For all but a few examples, posterior distributions cannot be integrated analytically.

The solution under Markov Chain Monte Carlo is to provide draws from the posterior, which allows numerical integration to be used. That is, if an N -long “chain” of parameter estimates β_i can be drawn from the true posterior, then any given statistic about that parameter can be calculated. For example, an estimate of the parameter mean can be estimated using $\bar{\beta} = \frac{1}{N} \sum_i \beta_i$. A more detailed description including the Markov Chain theory behind the technique can be found in Greenberg (2012).

Hamiltonian Monte Carlo (as described in Betancourt and Girolami (2015) and the Stan development team (2015)) provides draws from the posterior by specifying a joint probability between parameters θ and an auxiliary variable ρ , $p(\rho, \theta|y, X) = p(\rho|\theta, X, y) \times p(\theta|X, y)$. In most applications ρ comes from a multivariate normal distribution with covariance matrix Σ , $\rho \sim \mathcal{MVN}(0, \Sigma)$ and so is independent of θ, x, y . Borrowing from Hamiltonian Mechanics, the joint distribution defines a Hamiltonian:

$$H(\rho, \theta) = -\log p(\rho, \theta) = -\log p(\rho) - \log p(\theta|y, X) = T + V$$

Where T is the “kinetic energy” and V is the “potential energy”. Hamilton’s equations describe how a Hamiltonian evolves over time:

$$\frac{\partial \theta}{\partial t} = +\frac{\partial T}{\partial \rho}$$

$$\frac{\partial \rho}{\partial t} = -\frac{\partial V}{\partial \theta}$$

(the second equation holds as long as ρ is independent of the data and parameters). This system defines a two-state differential equation, which can be integrated using the leapfrog integrator, another tool borrowed from physics. Taking the initial position to be the current draw of θ , a draw of ρ is made from $\mathcal{MVN}(0, \Sigma)$. Initialised with these values, the leapfrog integrator proceeds for L steps, each of length ϵ , according to the following:

$$\rho \leftarrow \rho - \frac{\epsilon}{2} \frac{\partial V}{\partial \theta}$$

$$\theta \leftarrow \theta + \epsilon \Sigma \rho$$

$$\rho \leftarrow \rho - \frac{\epsilon}{2} \frac{\partial V}{\partial \theta}$$

This integration has the effect of “steering” θ towards parts of its distribution with higher density (unlike gradient descent, θ does not converge to the summit of the distribution, but continues over it with a given momentum). Consequently, θ spends a proportionately large amount of time towards the more dense region of its distribution, and less time at the tails. At the end of L steps, there will be some numerical error. This is because while the surface of the density of θ is smooth, movements along it in each step are straight (the error increases with step size). A Metropolis step is therefore used to decide whether to keep the draw, where the probability of keeping a draw that started at (ρ_0, θ_0) and arrived at (ρ_L, θ_L) after L steps is:

$$\min(1, \exp(H(\rho_0, \theta_0) - H(\rho_L, \theta_L)))$$

A vector of weights ω can be included in Hamiltonian Monte Carlo by weighting each observation’s impact on the potential energy term $V_\omega = -\omega \log p(\theta|y, X)$. Because the prior does not rely on the data, this involves weighting each observation’s contribution to the (log) likelihood. In Stan, a software package used to implement HMC, implementing these weights simply involves the following call in the model block of the program:

```
// likelihood (priors are defined above)

for(i in 1:number_of_observations){
  increment_log_prob(weight[i]* a likelihood expression of data and parameters)
}
```

Estimating models by HMC provides several advantages over other Bayesian techniques, such as Gibbs sampling or Metropolis Hastings (MH)—techniques described well in Tanner and Wong (1987), Gelfand and Smith (1990), and Chib and Greenberg (1995). The biggest is that because in a single iteration, θ may plausibly transition from any part of the the posterior to any other, draws are far more likely to be independent (see Betancourt and Girolami (2015)). This is not the case in Gibbs or Metropolis Hastings. Because draws are more likely to be independent, far fewer iterations are needed to draw inference about θ . The other large advantage is that HMC is able to explore enormous parameter spaces—in the order of tens to hundreds of thousands of dimensions—without much trouble. This makes it feasible (if a little slow) to estimate time-varying parameter models by drawing directly from the posterior of the varying state space, rather than using work-arounds like the Kalman Filter. This saves time in model building, and provides more flexibility in specifying richer models (for instance, allowing the state to evolve with non-normal innovations).

Bayesian models can be estimated efficiently using the Stan software package (Stan development team (2015)), which uses an implementation of HMC known as the No U-Turn Sampler (NUTS) of Homan and Gelman (2014). This algorithm automatically chooses the three

control parameters, L , ϵ and Σ , based on a series of adaptive iterations during the warm-up phase. Stan abstracts from the actual implementation of HMC, allowing the user to instead specify the probability model being estimated. Users can call programs written in Stan from within R, Python, Matlab, Julia or Stata.

4 Applications of analogy weighting

4a: An application to estimating core inflation

Most central banks have a mandate to fight inflation, either explicitly, as with a numerical inflation target, or implicitly, as with the Fed’s “just do it” approach (Mishkin 2001). The inflation that is typically targetted is a notion of “core inflation”, the inflation that would remain after self-correcting blips have been ignored. This is a latent quantity, and must be estimated. Various methods of estimating core inflation exist—several are discussed below.

Core inflation

Core inflation is a widely-used concept that lacks a formal definition or common tool for estimation, though two conceptual approaches are common.

The first commonly-used approach is to model core inflation as being some latent factor that drives observed price series. For instance, Kapetanios (2004) defines core inflation as the first factor from a singular value decomposition of a large number of component price series. This is similar to the approach of Stock and Watson (2008), who model core inflation using a dynamic factor model that allows for structural breaks. Simpler variants of this modelling philosophy work by filtering or smoothing price series, or by simply taking averages across price series that have had surprising components removed ad-hoc (see Quah and Vahey (1995)).

A second, more principled approach is to model core inflation as being the permanent component of inflation, and therefore the part that has no effect on medium or long-run economic activity. Quah and Vahey (1995), Ribba (2003) and Cogley and Sargent (2002)

take this approach. Cogley and Sargent’s method is to take core inflation to be the median of posterior predictive distributions for inflation into the distant future, using a Bayesian VAR with time-varying parameters. Quah and Vahey (1995) define core inflation to be the rate of inflation that has no impact on middle to long-run growth (estimated using a restricted classical VAR).

There are problems with both approaches. The dynamic factor model approach does not typically discriminate between periods in which a signal variable contains a lot of information about the unobserved factor and when it contains less. It is also sensitive to the choice of variables used as inputs to the factors. By dropping or censoring signal series, ad-hoc approaches may not have this problem, however are both costly and atheoretic. In the second approach, an economic model must be built that provides the best possible (long-run) forecasts, given information available at the time the forecast is being made. Thus the problem in estimating core inflation using this approach is one of finding a good forecasting model, and parameterising it with the most relevant parameters at the time of making forecasts.

In the demonstration below, I use the second approach to estimating core inflation. The demonstration makes use of three models: the first is a vanilla VAR with static parameters. The second is a time-varying parameter VAR based on the Cogley and Sargent (2002) model. The third model is a vanilla VAR augmented to make use of analogy weights.

Three approaches formally

Vanilla Bayesian VAR

The model describes the co-movement between inflation, unemployment and the yield on short-term Treasuries, using a VAR(2) specification. Taking Y_t to be vector containing all variables, the first model is

$$Y_t = \mu + \beta_1 Y_{t-1} + \beta_2 Y_{t-2} + \epsilon_t$$

where μ and the β s are coefficients and the multivariate residual $\epsilon \sim \mathcal{MVN}(0, \Sigma)$. The parameters are estimated in a Bayesian framework, and have so-called “shrinkage” priors

applied. These are all equal to 0 except the AR(1) terms, which have a prior of 1. The residual covariance matrix Σ is decomposed into a correlation matrix Ω and a scale vector τ such that

$$\Sigma = \text{diag}(\tau)\Omega\text{diag}(\tau)$$

Under this setup, we give independent priors to the correlation and scale: the scale vectors have a prior distribution of $\tau \sim \text{Cauchy}_+(0, 2.5)$, and the correlation comes from the the LKJ distribution $\Omega \sim \text{lkj}(1.5)$. The LKJ distribution of Lewandowski, Kurowicka, and Joe (2009) is a quite new distribution for correlation matrices that addresses some of the shortcomings of the popular Inverse-Wishart prior. The single parameter of the distribution dictates how much cross-correlation the distribution will produce; a value of 1 results in a uniform prior, while as this degrees of freedom parameter goes to infinity, the distribution collapses on the unit-diagonal correlation matrix.

I experimented with two approaches to modelling the covariance between the coefficients μ , β_1 and β_2 . Estimating a full covariance matrix would be too burdensome given the amount of data available (it would involve estimating 210 parameters in the covariance matrix alone). One approach is to use a Littermann prior, which restricts Ω to be diagonal and specifies priors on τ . Another approach is to estimate only the covariances within each “block”; that is, allow μ to have an estimated covariance, and likewise with the β s. This has the advantage that it only requires estimating 78 covariance parameters, and was the approach taken for this model.

The model is estimated using HMC as described above, with no weights.

While being fitted, the model calculates 1-year-ahead posterior predictions; these make use of both the uncertainty in the regression error and uncertainty in parameter estimates. These predictions form the basis for the model comparisons below.

Weighted VAR

The second model fitted takes exactly the same form as the VAR described above, with

one difference: it is fitted using analogy weights. These weights are calculated as described above, using inflation, unemployment and the interest rate as the variables used to estimate analogies.

Time-varying parameter VAR

The third model considered is a Time-varying-parameter VAR, similar to the one described in Cogley and Sargent (2002). The model is

$$Y_t = \mu_t + \beta_{1,t}Y_{t-1} + \beta_{2,t}Y_{t-2} + \epsilon_t$$

with the coefficients themselves evolving over time. Take $\theta = [\mu, \text{vec}\beta_1, \text{vec}\beta_2]'$, where the vec operation stacks the a matrix in column-major order, then:

$$\theta_t = \theta_{t-1} + \eta_t$$

and $\eta \sim \mathcal{MVN}(0, \Sigma_\eta)$. For identifiability, suitable structure must be put on Σ_η ; as with the static VARs, I assume that it is block diagonal. The covariance matrix is modelled as the quadratic form of a scale vector and a correlation matrix, and the correlation matrix is given an LKJ prior with a large degrees of freedom parameter. This specification allows some covariance between parameters within the same block, but shrinks the covariance towards zero.

Data and method

The data series used include quarterly observations of the quarterly CPI growth rate for major cities; the national non-farm unemployment rate; and the annualised yield on 90 day Treasuries. The raw data series are monthly, and we need to convert these into quarterly observations. For CPI growth, the first month's value for each quarter is assigned to the whole quarter. For unemployment we also take the first month of each quarter. For treasury yields,

the average value during each quarter is used. The data come from the FRED database, and are the current releases (not archival data).²

Using observations starting 1960Q1, the two static parameter models are estimated up to a given date. This date starts at 1970Q1 and extends in 1-year increments. At the end of each estimation period, the models generate 1-year ahead posterior predictive forecasts. For the same comparison, the time-varying parameter model is estimated over the whole period, and one-year-ahead predictions are made using the estimated parameter values for a given period. These forecasts are used to compare the predictive accuracy of each model.

Each model then produces an estimate of the core or permanent inflation by taking a limiting posterior forecast of each series. For both the vanilla VAR and weighted VAR, the limiting forecast is simply the unconditional mean of each series. This is achieved by re-writing the VAR in lag notation:

$$(\mathbf{I} - \beta_1 L - \beta_2 L^2)Y_t = \mu + \epsilon_t$$

Taking expectations and inverting gives

$$E[Y] = (\mathbf{I} - \beta_1 L - \beta_2 L^2)^{-1}\mu = (\mathbf{I} - \beta_1 - \beta_2)^{-1}\mu$$

The second equality holds because $L\mu = \mu$. For the time-varying VAR, this inversion is not trivial, as μ_t varies over time. In this case, a limiting forecast is obtained by forward-simulating the model 100 quarters, with explosive states discarded, as in Cogley and Sargent (2002).

²In an experiment, I sourced the archival series from ALFRED, and back-tested the out-of-sample predictive performance of both the vanilla VAR and analogy-weighted VAR. 1-year-ahead forecasts of inflation and the interest rate were significantly lower for the weighted model, though unemployment forecasts were no better. These results are omitted because of the high computational cost in re-fitting the TVP-VAR model to the archival data.

Forecasting performance of the three approaches

The three models are estimated as described above, yielding two sets of forecasts: the first being one-year-ahead forecasts. The accuracy of the 1-year-ahead forecasts can be measured by comparing against history. The second set of forecasts are the “permanent components” of the series—what is left when the model has been simulated into the distant future. It is from these permanent forecasts that we derive an estimate of core inflation. Yet the quality of short-run forecasts provide some guidance about which model’s estimates to believe.

The table on the following page contains the mean-absolute-error of the one-year ahead forecasts. These show that the TVP-VAR and the weighted VAR both tend to perform better than the unweighted VAR. The dominance of the TVP-VAR should not come as a surprise, as the forecasts are not true out-of-sample. The relative performance of the weighted VAR over the unweighted VAR is suggestive of analogy weights adding value.

Mean absolute errors of three models for one-year-ahead forecasts

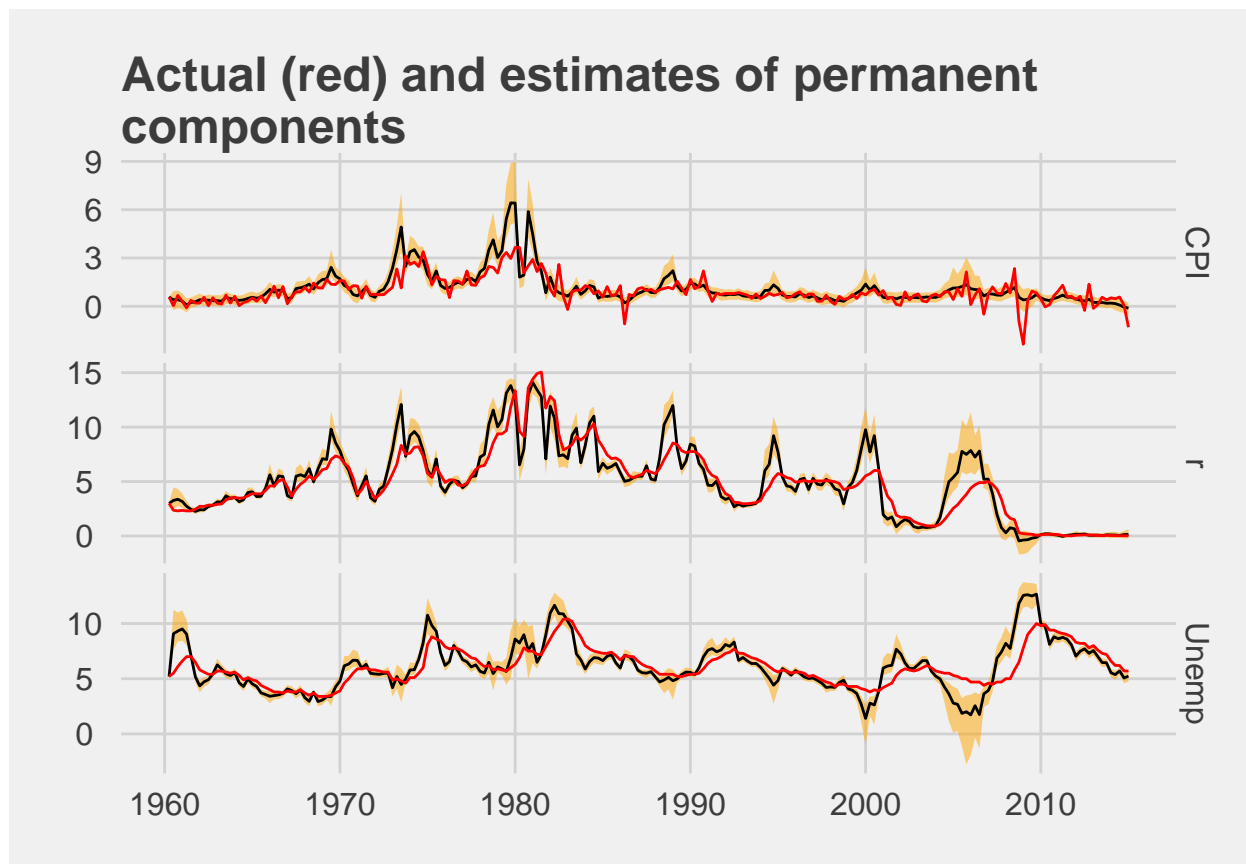
Decade	CPI inflation			Interest rate			Unemployment		
	TVP	VAR	AW-VAR	TVP	VAR	AW-VAR	TVP	VAR	AW-VAR
1970	0.15	0.35	0.35	1.25	1.99	2.07	0.90	1.28	1.44
1980	0.91	0.95	0.96	2.11	1.89	1.79	0.61	0.56	0.59
1990	0.16	0.29	0.28	1.27	0.92	0.86	0.40	0.44	0.42
2000	0.60	0.72	0.62	1.16	1.39	1.26	0.53	0.59	0.48
2010	0.54	0.56	0.51	0.06	1.68	1.09	0.83	1.36	1.35
Whole period	0.51	0.61	0.58	1.3	1.49	1.36	0.59	0.71	0.7

The results from this horse-race suggest that including analogy weights in a time-series model can significantly improve its predictive performance. In this test, the performance of the weighted model approaches (and in some cases exceeds) that of a more complex time-varying-parameter model.

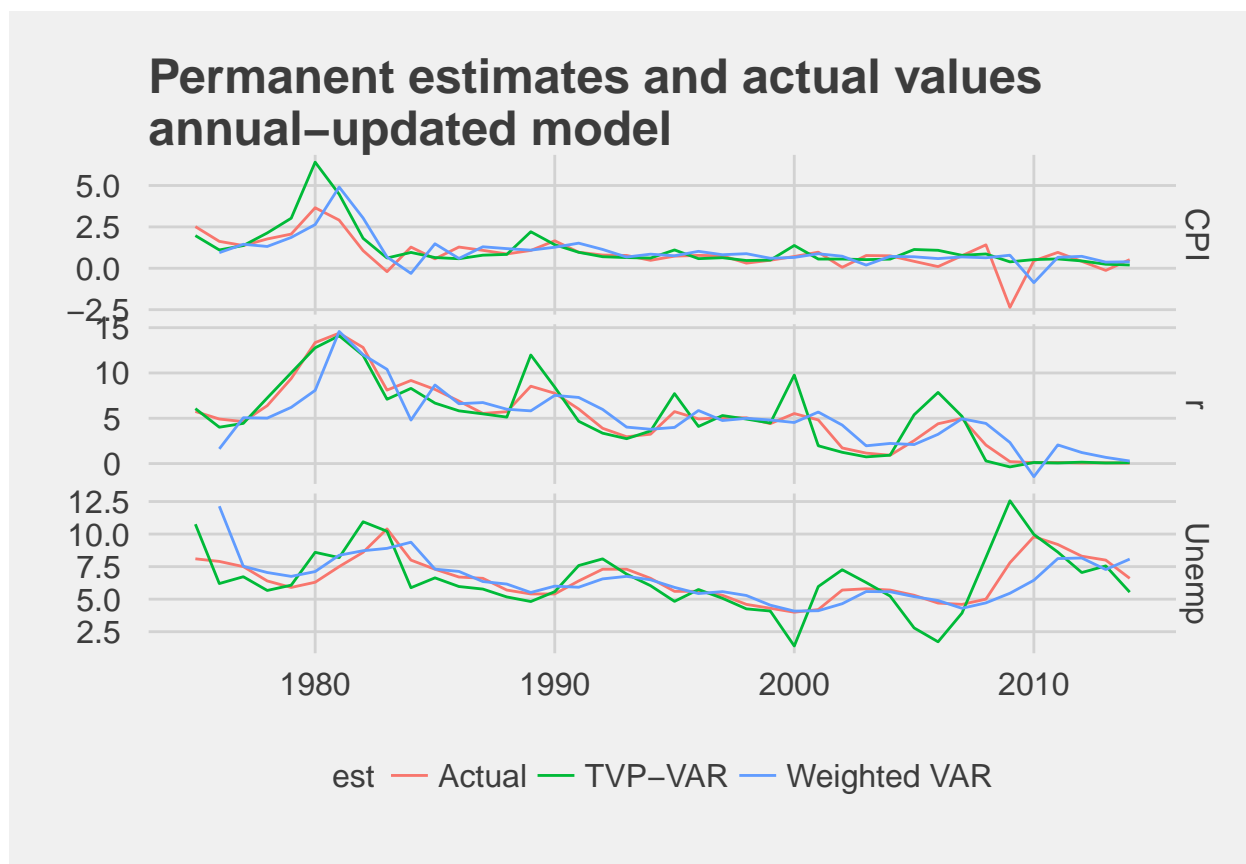
Core inflation and the Fed’s response

The model comparison above suggests that both the TVP-VAR and analogy-weighted VAR were good candidates for a forecasting model, and may therefore be useful for estimating core inflation. Yet the two models have quite different interpretations. Under the TVP-VAR model, parameters vary over time as new information comes to light, in a form of “Bayesian learning”. In the weighted VAR model, parameter variation occurs due to analogising with similar historical episodes; it therefore summarises “historical learning”. An interesting question is whether the Federal Reserve’s determination of interest rates is more consistent with the Bayesian learning paradigm than the historical learning.

The figure below shows the quarterly estimates of the permanent components of (quarterly) CPI inflation, the nominal interest rate and unemployment from the TVP-VAR model. These estimates are the limiting forecast from the TVP-VAR, as discussed above. Of interest is the high variability of the permanent component of the nominal interest rate relative to actual nominal interest rates; these large increases correspond strongly with periods of high estimated core inflation. The turning points of unemployment also precede the actual turning points.



The results from the static-parameter VAR models are somewhat more muted. Due to the amount of computing time it would take to update these models at each quarter, these values were calculated annually, and so don't bear direct comparison with the chart above. The chart below illustrates estimates of the permanent components of the TVP-VAR and analogy-weighted VAR models.



Does the Federal Reserve adjust interest rates due to views about core inflation the emerges from a Bayesian learning model or from a historical learning model? To help understand, I regress the interest rate on both the estimated core inflation series. The results of this regression are in the table below. The model suggests that both estimates of core inflation are important in the conduct of monetary policy, suggesting that monetary policymakers do analogue with similar historical episodes.

Do both historical analogy and Bayesian learning estimated inform current interest rates?

	Model 1	Model 2	Model 3
(Intercept)	2.068**	2.436***	1.804**
	(0.612)	(0.566)	(0.565)
AW_core_inf	2.911***		1.669**
	(0.450)		(0.590)
BL_core_inf		2.271***	1.339**
		(0.347)	(0.458)
R-squared	0.531	0.537	0.621
F	41.915	42.922	29.510
p	0.000	0.000	0.000
N	39	39	39

4b: An application to portfolio risk management

Portfolio risk managers are tasked with holding sufficient cash to insure that asset purchases and withdrawals from the portfolio can be paid for even when asset prices fall. This requires generating estimates of Value At Risk (VaR), a measure defined as the minimum loss expected in some event that should occur with a given frequency over a period. For example, if a portfolio has \$5M VaR over 1 day at 95%, this would mean that not more than one in twenty days should involve losses greater than \$5M.

Due to volatility clustering (Mandelbrot 1967), the VaR of a portfolio will change over time. More volatile periods will imply a greater chance of loss, and so VaR will increase. A good volatility model can help managers hold liquidity at times that are most likely to be most volatile. Generalized Auto-Regressive Conditional Heteroskedasticity (GARCH)-style models (due to Bollerslev (1986)) are a common choice for this task.

For a stationary univariate time series y_t , a GARCH(p,q) model (for simplicity with no lags

of y) has the following specification:

$$y_t = \mu + h_t \eta_t$$

Where μ is the unconditional mean value of y , η is i.i.d. around 0, and h_t is the conditional standard deviation of residuals. h_t evolves according to:

$$h_t^2 = \alpha_0 + \sum_{i=1}^p \alpha_i (y_{t-i} - \mu)^2 + \sum_{j=1}^q \beta_j h_{t-j}^2$$

For a given starting value h_0 , estimating a GARCH model requires estimating only μ , a $p+1$ -dimensional vector α , and a q -dimensional vector β . This is most commonly done using maximum likelihood, and can be estimated using Bayesian techniques also.

While univariate GARCH models offer a simple method of modelling conditional volatility, often a risk manager may be concerned about the correlation between assets' returns. The Constant Conditional Correlation (multivariate) GARCH model of Bollerslev (1990) provides a very simple method of modelling this correlation. A vector of stationary variables y_t with unconditional mean μ and covariance matrix H_t behaves according to

$$y_t = \mu + H_t^{1/2} \eta_t$$

Where η is iid with a unit diagonal covariance matrix. The covariance matrix is modelled as being the diagonal quadratic form of a scale vector that varies over time τ_t and a correlation matrix that is fixed, Ω .

$$H_t = \text{diag}(\tau_t) \Omega \text{diag}(\tau_t)$$

And the scale parameters themselves are modelled as coming from a GARCH process.

$$\tau_t \odot \tau_t = \omega + \sum_{i=1}^p A_i (y_{t-i} - \mu) \odot (y_{t-i} - \mu) + \sum_{j=1}^q B_j \tau_{t-1-j} \odot \tau_{t-1-j}$$

Here, \odot is elementwise multiplication. ω is a vector of intercept terms, while A and B are diagonal matrices of parameters.

CCC-GARCH models offer an improvement beyond simple GARCH models of volatility in that they allow correlation between series. Yet for a range of tasks, the fixed correlation structure is too restrictive. As outlined in Engle (2002), a wide range of tasks in financial management require updated estimates of the correlation between variables. This necessity has led to a large number of GARCH-style models that have time-varying correlation matrices, including popular methods like VC-GARCH of Tse and Tsui (2002) and DCC-GARCH of Engle (2002), in which correlations react to recent comovements. The latter of these is used to benchmark the predictive performance of an analogy-weighted volatility model below. A detailed description of the DCC-GARCH model can be found in Engle (2002).

The analogy-weighted CCC-GARCH model described below improves on Bollerslev (1990) model by making two changes. First, it allows for analogy weighting. The practical impact of allowing analogy weights is to give time-variation to the correlation matrix, so that the model will have similar correlations between asset returns across similar states of history.

The second big improvement is an LKJ prior for the correlation matrix with a small degrees of freedom parameter. This has the effect of “switching off” the model when the world alters states to a new (unprecedented) state, imposing large correlations between returns, increasing VaR. It does this because an unprecedented state has little in common with historical observations, and so the model draws more inference from the prior. Because the LKJ prior is a uniform correlation matrix when the degrees of freedom parameter is 1, relying more on the prior implies assuming more correlation between observations.

A modern CCC-GARCH model and the analogy augmentation

The functional form of the augmented CCC-GARCH model is identical to the one as described above—all that differs is the estimation technique. To provide Bayesian inference about the model, prior distributions are given to the parameters. Elements of μ , $\text{diag}(A)$ and $\text{diag}(B)$ are all given univariate normal priors. For stability, it is necessary for elements

of $\text{diag}(A)$ to be greater than 1, and $\text{diag}(B) \leq I - \text{diag}(A)$. The initial values of the scale parameter, τ_0 , are given half Cauchy priors, which allow for large initial variances. Finally, the correlation matrix Ω is given an LKJ prior with a degrees of freedom parameter close to 1; this has the effect of increasing modelled correlations during unprecedented phases. Weights are incorporated as described above—by weighting each observation’s contribution to the log posterior.

To demonstrate the performance of the two techniques, an equal-weighted portfolio is constructed from common stocks in Apple (AAPL), IBM (IBM), Google (GOOGL), Ford (FORD), Bank of America (BAC), and Johnson and Johnson (JNJ). Two versions of a CCC-GARCH(1,1) model are estimated on weekly log returns in successive 1-week increments: an unweighted version, and one using analogy weights. Both models are then tasked with generating estimates of 1-week ahead 95% VaR for the period July 2007 through June 2013 using true out-of-sample prediction. For example, the predicted volatility for the week of July 6 2008 would involve estimating using data from August 22, 2004 through June 29, 2008 only. For the next iteration, models are estimated from the same start-date up to July 6, and make another 1-week-ahead forecast. This is continued until the end of June 2013; the test period covers the most turbulent period of the Global Financial Crisis of 2008-9.

Week-ahead VaR estimates are generated by generating 1-week ahead posterior predictive distributions for all stocks in the portfolio. The posterior predictive distribution is similar to a classical forecast, but allows for uncertainty in the model parameters. In each HMC draw, the projections for each stock are aggregated to construct a portfolio return; the VaR measure is then the bottom 5th percentile of the resulting distribution. For the benchmark DCC-GARCH model, 1-week-ahead forecasts of the mean vector and covariance matrix are taken. These parameters inform a multivariate normal distribution, from which 1000 draws are taken, and the portfolio return calculated for each draw. The 5th percentile of portfolio returns is then stored. For a correctly specified model, losses should not exceed VaR more than one in twenty weeks; the evaluation period involves the Global Financial Crisis, a period over which Value at Risk models notoriously underperformed (Jorion 2009).

The analogy weights are constructed using the technique described above, using random

forests to predict the weekly squared returns of each of the stocks in question, using their own (squared) first and second lags, and the (squared) first and second lags of the other stocks.

The unconditional mean prior is $\mu \sim \mathcal{N}(0, 0.01)$, indicating that we expect the unconditional weekly return to be close to 0 over long periods. The volatility scale prior is $\tau \sim \text{Cauchy}_+(0, 2.5)$, while the prior on elements of μ , A , and B are all $\mathcal{N}(0, 1)$.

Predictive performance

This section assesses the predictive performance of the analogy-weighted CCC-GARCH model against two other models. The back-testing occurs over the Global Financial Crisis of 2008-9. The first benchmark model is an unweighted CCC-GARCH model of an identical specification to the weighted CCC-GARCH model, though with no weighting scheme. It illustrates the effect of including analogy weights. The second comparison model is the DCC-GARCH model of Engle (2002), implemented using the `rmgarch` package of Ghalanos (2014). The second comparison is useful as DCC-GARCH explicitly models the time variation in correlation structures—a major limitation in CCC-GARCH which analogy weighting should help overcome.

Assessing the performance of volatility models is not straightforward. For one, over-predicting volatility has an opportunity cost (as it implies holding more low-yielding assets), while under-predicting volatility can lead to insufficient liquidity being held. Consequently a loss function needs to be specified, as discussed in Hou and Suardi (2012). Common choices include a mean-absolute-error of the volatility forecasts $MAE = T^{-1} \sum_t |\hat{\sigma}_t^2 - \epsilon_t^2|$, mean squared error of volatility forecasts, $MSE = T^{-1} \sum_t (\hat{\sigma}_t - \epsilon_t)^2$, the heteroskedasticity-adjusted mean absolute error of Bollerslev and Ghysels (1996), which is defined as $HMAE = T^{-1} \sum_t \left| \frac{\epsilon_t^2}{\hat{\sigma}_t^2} - 1 \right|$ and the quadratic version: $HMSE = T^{-1} \sum_t \left(\frac{\epsilon_t^2}{\hat{\sigma}_t^2} - 1 \right)^2$, where ϵ_t^2 are the squared errors of returns and $\hat{\sigma}_t^2$ is the volatility forecast for $t \in (1, T)$. Another loss function due to Pagan and Schwert (1990) is the log loss function $LL = T^{-1} \sum_t \log \left(\frac{\epsilon_t^2}{\hat{\sigma}_t^2} \right)^2$. All five loss functions are used to compare the performance of the three models. Results for the loss functions evaluated over the full back-test (2007-07-01 through 2013-06-30) are in the table below.

Loss functions evaluated for week-ahead forecasts, July 2007–June 2013

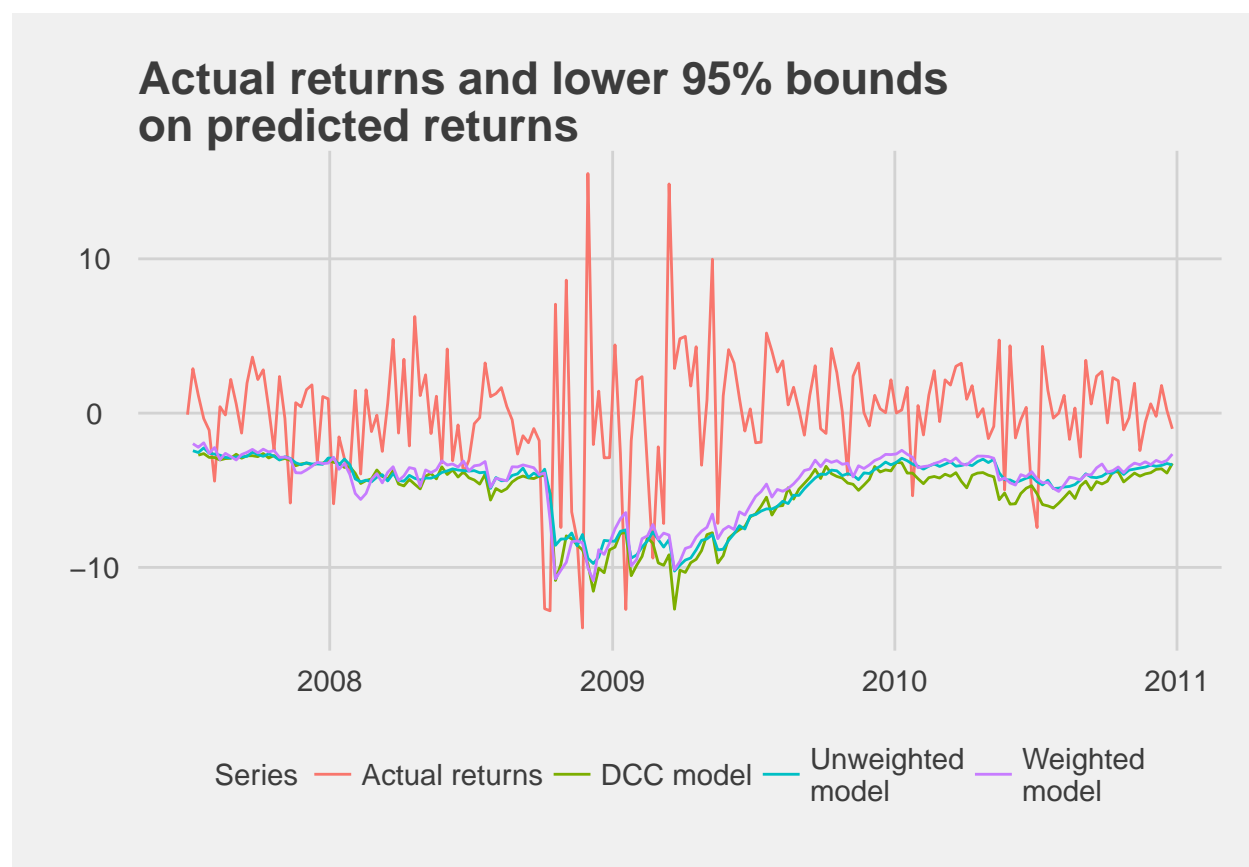
Loss function	Unweighted	Weighted	DCC
Quadratic loss	34.31	32.52	38.54
Mean absolute value	18.39	17.20	21.56
Heteroskedasticity-adjusted absolute value	0.85	0.86	0.83
Heteroskedasticity-adjusted quadratic loss	1.34	1.32	1.09
Log loss	9.93	9.50	10.79

As can be seen, over the whole period, the weighted model outperforms in three of the five loss measures, with DCC-GARCH outperforming in the other two. The two measures on which the DCC-GARCH model outperformed especially penalises large negative errors that occur during low-volatility periods. Note that for the loss metrics in which the weighted model outperforms DCC-GARCH, the CCC-GARCH model also outperforms DCC-GARCH. As the DCC-GARCH model provides consistently higher volatility forecasts during the second half of the sample (which has lower volatility), its negative errors during this period are smaller in magnitude than for the other models. Another assessment can be made on the period during the global financial crisis of 2008-2009, when Value at Risk models performed quite poorly. When the assessment is restricted to this period, the rank ordering of models appears to be the same, though the advantage of the DCC-GARCH model diminishes.

Loss functions evaluated for week-ahead forecasts, January 2008–December 2009

Loss function	Unweighted	Weighted	DCC
Quadratic loss	60.87	57.33	65.66
Mean absolute value	33.85	31.71	37.84
Heteroskedasticity-adjusted absolute value	0.94	0.92	0.91
Heteroskedasticity-adjusted quadratic loss	2.25	1.85	1.75
Log loss	8.13	7.45	8.53

The performance of the three models can be seen in the figure below. Two things stand out: first, the lower bound on both the DCC-GARCH and weighted model respond more sharply to the first couple of shocks during the crisis; throughout the crisis their predicted volatility is higher than for the CCC-GARCH. Second, during the return to a low-volatility state after the crisis, the weighted model drops far more quickly than either the DCC-GARCH or unweighted model. This suggests that the weighting scheme may have a similar effect to the regime-switching of Klaassen (2002).



5: Conclusion

Training time-series models on the most relevant data improves their predictive performance, suggesting that the estimated parameters are more relevant to the present. Yet existing methods to explicitly model the time-variation in parameters are both difficult to implement and richly parameterised, necessitating strong modelling assumptions. The method presented

in this paper helps overcome these shortcomings, by building a weighting vector that measures similarity between historical periods in terms of the variables that help to predict the dependent variable. Once the weighting vector is estimated, it can be incorporated into simple time-series regression models in a straightforward manner.

There are two reasons one may wish to use analogy weights. The first is that it tends to improve models' predictive performance at a low cost. The second is that because analogy weights shrink as the present value of economic variables shifts away from historical observations, the method generates more prediction uncertainty during unprecedented phases. These two benefits come at very little cognitive cost. Implementing the method requires training random forests on the dependent variables and saving the (last row of the) proximity matrix; the average of this row across runs becomes the analogy weights.

The main problem with the method is that the weights do not come from a generative model; it is therefore impossible to verify whether a model estimated with analogy weights provides better estimates of “relevant” parameters than any other method. The attempt at verifying the approach in this paper is to subject the model to out-of-sample forecasting checks. Success at this step is a necessary, but not sufficient condition for the approach to be valid.

The conjecture made (but not shown) in this paper is that *relevant periods*—relevance defined on observed variables—should help more in estimating present parameter values. Underlying this is an assumption that the parameter values themselves are in part determined by the values of the observed variables used to construct analogy scores. A sufficient test of the method in this paper would be to simulate data from a generative model whose (reduced form) parameters are themselves functions of the variables, and check whether analogy score weighting succeeds at estimating these parameters. This is left to future research.

References

- Alloway, Nasiripour; Masters; 2013. “Basel Eyes Set Periods for Banks’ Risk Models.” *Financial Times*.
- Athey, Susan, and Guido Imbens. 2015. “Lectures on Machine Learning.”

- Auerbach, Alan J, and Yuriy Gorodnichenko. 2012. “Fiscal Multipliers in Recession and Expansion.” In *Fiscal Policy After the Financial Crisis*, 63–98. University of Chicago press.
- Betancourt, Michael, and Mark Girolami. 2015. “Hamiltonian Monte Carlo for Hierarchical Models.” *Current Trends in Bayesian Methodology with Applications*. CRC Press, 79.
- Bianchi, Javier. 2010. “Credit Externalities: Macroeconomic Effects and Policy Implications.” *The American Economic Review*. JSTOR, 398–402.
- Bollerslev, Tim. 1986. “Generalized Autoregressive Conditional Heteroskedasticity.” *Journal of Econometrics* 31 (3). Elsevier: 307–27.
- . 1990. “Modelling the Coherence in Short-Run Nominal Exchange Rates: A Multivariate Generalized ARCH Model.” *The Review of Economics and Statistics*. JSTOR, 498–505.
- Bollerslev, Tim, and Eric Ghysels. 1996. “Periodic Autoregressive Conditional Heteroscedasticity.” *Journal of Business & Economic Statistics* 14 (2). Taylor & Francis Group: 139–51.
- Breiman, Leo. 2001. “Random Forests.” *Machine Learning* 45 (1). Springer: 5–32.
- Breiman, Leo, Jerome Friedman, Charles J Stone, and Richard A Olshen. 1984. *Classification and Regression Trees*. CRC press.
- Brunnermeier, Markus K, and Yuliy Sannikov. 2012. “A Macroeconomic Model with a Financial Sector.” *National Bank of Belgium Working Paper*, no. 236.
- Caruana, Rich, and Alexandru Niculescu-Mizil. 2006. “An Empirical Comparison of Supervised Learning Algorithms.” In *Proceedings of the 23rd International Conference on Machine Learning*, 161–68. ACM.
- Chib, Siddhartha, and Edward Greenberg. 1995. “Understanding the Metropolis-Hastings Algorithm.” *The American Statistician* 49 (4). Taylor & Francis Group: 327–35.
- Cogley, Timothy, and Thomas J Sargent. 2002. “Evolving Post-World War II US Inflation Dynamics.” In *NBER Macroeconomics Annual 2001, Volume 16*, 331–88. MIT Press.
- . 2005. “Drifts and Volatilities: Monetary Policies and Outcomes in the Post WWII US.” *Review of Economic Dynamics* 8 (2). Elsevier: 262–302.

- Deng, Kan, Andrew W Moore, and Michael C Nechyba. 1997. "Learning to Recognize Time Series: Combining Arma Models with Memory-Based Learning." In *Computational Intelligence in Robotics and Automation, 1997. CIRA'97., Proceedings., 1997 IEEE International Symposium on*, 246–51. IEEE.
- Engle, Robert. 2002. "Dynamic Conditional Correlation: A Simple Class of Multivariate Generalized Autoregressive Conditional Heteroskedasticity Models." *Journal of Business & Economic Statistics* 20 (3). Taylor & Francis: 339–50.
- Fabozzi, Frank J, and Steven V Mann. 2012. *The Handbook of Fixed Income Securities*. McGraw Hill Professional.
- Gelfand, Alan E, and Adrian FM Smith. 1990. "Sampling-Based Approaches to Calculating Marginal Densities." *Journal of the American Statistical Association* 85 (410). Taylor & Francis Group: 398–409.
- Ghalanos, Alexios. 2014. *rmgarch: Multivariate GARCH Models*.
- Hamilton, James D. 1988. "Rational-Expectations Econometric Analysis of Changes in Regime: An Investigation of the Term Structure of Interest Rates." *Journal of Economic Dynamics and Control* 12 (2). Elsevier: 385–423.
- Hamilton, James D, and Baldev Raj. 2013. *Advances in Markov-Switching Models: Applications in Business Cycle Research and Finance*. Springer Science & Business Media.
- Homan, Matthew D, and Andrew Gelman. 2014. "The No-U-Turn Sampler: Adaptively Setting Path Lengths in Hamiltonian Monte Carlo." *The Journal of Machine Learning Research* 15 (1). JMLR. org: 1593–1623.
- Hou, Aijun, and Sandy Suardi. 2012. "A Nonparametric GARCH Model of Crude Oil Price Return Volatility." *Energy Economics* 34 (2). Elsevier: 618–26.
- Jorion, Philippe. 2009. "Risk Management Lessons from the Credit Crisis." *European Financial Management* 15 (5). Wiley Online Library: 923–33.
- Kapetanios, George. 2004. "A Note on Modelling Core Inflation for the UK Using a New Dynamic Factor Estimation Method and a Large Disaggregated Price Index Dataset." *Economics Letters* 85 (1). Elsevier: 63–69.

- Keogh, Eamonn, and Jessica Lin. 2005. "Clustering of Time-Series Subsequences Is Meaningless: Implications for Previous and Future Research." *Knowledge and Information Systems* 8 (2). Springer: 154–77.
- Kim, Chang-Jin, and Christian J Murray. 2002. "Permanent and Transitory Components of Recessions." In *Advances in Markov-Switching Models*, 19–39. Springer.
- Kiyotaki, Nobuhiro, and John Moore. 1995. "Credit Cycles." National Bureau of Economic Research.
- Klaassen, Franc. 2002. "Improving GARCH Volatility Forecasts with Regime-Switching GARCH." In *Advances in Markov-Switching Models*, 223–54. Springer.
- Koop, Gary, and Dimitris Korobilis. 2010. *Bayesian Multivariate Time Series Methods for Empirical Macroeconomics*. Now Publishers Inc.
- . 2013. "Large Time-Varying Parameter VARs." *Journal of Econometrics* 177 (2). Elsevier: 185–98.
- Koop, Gary, Roberto Leon-Gonzalez, and Rodney W Strachan. 2009. "On the Evolution of the Monetary Policy Transmission Mechanism." *Journal of Economic Dynamics and Control* 33 (4). Elsevier: 997–1017.
- Kuan, CHUNG-MING. 2002. "Lecture on the Markov Switching Model." *Institute of Economics, Academia Sinica, Taipei* 115.
- Lewandowski, Daniel, Dorota Kurowicka, and Harry Joe. 2009. "Generating Random Correlation Matrices Based on Vines and Extended Onion Method." *Journal of Multivariate Analysis* 100 (9). Elsevier: 1989–2001.
- Mandelbrot, Benoit. 1967. "The Variation of Some Other Speculative Prices." *Journal of Business*. JSTOR, 393–413.
- Miller, David, James Savage, and Sarah Tan. Forthcoming. "Random Forest Proximity Matching for Causal Inference."
- Mishkin, Frederic S. 2001. "Monetary Policy." *NBER Reporter*. National Bureau of Economic Research, Inc., 8.

- Pagan, Adrian R, and G William Schwert. 1990. "Alternative Models for Conditional Stock Volatility." *Journal of Econometrics* 45 (1). Elsevier: 267–90.
- Quah, Danny, and Shaun P Vahey. 1995. "Measuring Core Inflation." *The Economic Journal*. JSTOR, 1130–44.
- Ribba, Antonio. 2003. "Permanent-Transitory Decompositions and Traditional Measures of Core Inflation." *Economics Letters* 81 (1). Elsevier: 109–16.
- Smith, Jeffrey, and Petra Todd. 2001. "Reconciling Conflicting Evidence on the Performance of Propensity-Score Matching Methods." *American Economic Review*. JSTOR, 112–18.
- Stan development team. 2015. *Stan Modeling Language Users Guide and Reference Manual, Version 2.8.0*. <http://mc-stan.org/>.
- Stock, James H, and Mark W Watson. 2008. "Phillips Curve Inflation Forecasts." National Bureau of Economic Research.
- Tanner, Martin A, and Wing Hung Wong. 1987. "The Calculation of Posterior Distributions by Data Augmentation." *Journal of the American Statistical Association* 82 (398). Taylor & Francis: 528–40.
- Tse, Yiu K, and Albert K C Tsui. 2002. "A Multivariate Generalized Autoregressive Conditional Heteroscedasticity Model with Time-Varying Correlations." *Journal of Business & Economic Statistics* 20 (3). Taylor & Francis: 351–62.
- Varian, Hal R. 2014. "Big Data: New Tricks for Econometrics." *The Journal of Economic Perspectives*. JSTOR, 3–27.
- Wager, Stefan, and Susan Athey. 2015. "Estimation and Inference of Heterogeneous Treatment Effects Using Random Forests." *ArXiv Preprint ArXiv:1510.04342*.
- Wang, Xiaozhe, Kate Smith, and Rob Hyndman. 2006. "Characteristic-Based Clustering for Time Series Data." *Data Mining and Knowledge Discovery* 13 (3). Springer: 335–64.