

EN.553.732 (STAT 732) Homework 4

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December 18, 2017

Problem 1. *Problem 9.2 from Hoff*

Part a

Using a g-prior with $g = n$, $\nu_0 = 2$ and $\sigma_0^2 = 1$, a regression model was fit and the posterior confidence intervals for the parameters were computed (provided in Table 1 below). The plots of the posterior distributions were also computed (blue lines are 95% confidence intervals)

R Code

```
#Problem 1 Part a
set.seed(1)
t=10000
azdiabetes = read.table("azdiabetes.dat", header = TRUE);
y=as.matrix(azdiabetes[,2])
g=length(y)
X=as.matrix(cbind(rep(1,g),azdiabetes[,-c(2,8)]))
n=nrow(X)
m=ncol(X)
nu_0=2
sigma2_0=1
Hg=(g/(g+1))*X%%solve(t(X)%X)%t(X)
SS=t(y)%%(diag(1,nrow=n)-Hg)%y
sigma2=1/rgamma(t,(nu_0+n)/2,(nu_0*sigma2_0+SS)/2)
Vb=g*solve(t(X)%X)/(g+1)
Eb=Vb%%t(X)%y
E=matrix(rnorm(t*m,0,sqrt(sigma2)),t,m)
b=t(t(E%%chol(Vb))+c(Eb))
for (i in 1:m)
{
  print(quantile(b[,i],c(0.025, 0.975)))
}
plot(density(b[,2]),col=1, xlab=names(azdiabetes)[1],
main ="Problem 1 Part a")
abline(v=mean(b[,2]),col=2)
abline(v=quantile(b[,2],0.025),col=4)
```

```

abline(v=quantile(b[,2],0.975),col=4)
for (i in 2:6)
{
  plot(density(b[,i+1]),col=1, xlab=names(azdiabetes)[i+1],
  main ="Problem 1 Part a")
  abline(v=mean(b[,i+1]),col=2)
  abline(v=quantile(b[,i+1],0.025),col=4)
  abline(v=quantile(b[,i+1],0.975),col=4)
}

```

Table 1: Posterior Confidence Intervals

	95% CI
Intercept	(35.083, 69.244)
npreg	(-1.625, 0.304)
bp	(-0.0207, 0.429)
skin	(-0.122, 0.504)
bmi	(0.161, 1.141)
ped	(3.159, 17.923)
age	(0.456, 1.080)

Posterior distribution plots *Blue lines correspond to the 95% confidence intervals*

1a1.jpeg

1a2.jpeg

1a3.jpeg	1a4.jpeg
1a5.jpeg	1a6.jpeg

Part b

Now, we perform the model selection process and the averaging procedure using the same g-prior. Table 2 below summarizes the results for $Pr(\beta_j \neq 0|y)$ and the posterior confidence intervals for all parameters. The plots below are the posterior distributions for the parameters, and the blue lines are the 95% confidence intervals. Comparing the results from part a and b, we see that for the **npreg**, **bp**, and **skin** parameters, their respective 95% posterior confidence intervals include 0 for in part b. The value of $Pr(\beta_j \neq 0|y)$ for these three parameters are all close to 0, and their posterior distributions are more narrow than in part a. For the other parameters, where $Pr(\beta_j \neq 0|y)$ is close to 1, the confidence intervals are not significantly different from their confidence intervals in part a.

R Code

```
#Problem 1 Part b
lpy.X=function(y, X)
{
  n=nrow(X)
  m=ncol(X)
  g=length(y)
  nu_0=1
  sigma2_0=try(summary(lm(y ~ -1+X))$sigma^2, silent = TRUE)
  if (m == 0)
  {
    Hg=0
    sigma2_0=mean(y^2)
  }
  else if (m > 0)
  {
    Hg=(g/(g+1))*X%%solve(t(X)%X)%t(X)
  }
  SS=t(y)%%(diag(1, nrow = n)-Hg)%y
  -(1/2)*(n*log(pi)+m*log(1+g)+(nu_0+n)*log(nu_0*sigma2_0+SS)-
  nu_0*log(nu_0*sigma2_0))+lgamma((nu_0+n)/2)-lgamma(nu_0/2)
}
t=1000
A=matrix(NA, t, m)
B=matrix(0, t, m)
a=rep(1, m)
lpy.c=lpy.X(y, X[, a == 1, drop = FALSE])
for(i in 1:t)
{
  for(j in sample(1:m))
  {
    temp=a
    temp[j]=1-temp[j]
    lpy.m=lpy.X(y, X[, temp == 1, drop = FALSE])
    r=(lpy.m-lpy.c)*(-1)^(temp[j] == 0)
    a[j]=rbinom(1, 1, 1/(1+exp(-r)))
    if (a[j] == temp[j])
    {
      lpy.c=lpy.m
    }
    A[i, ]=a
    Hg=(g/(g+1))*X[, A[i,] == 1, drop = FALSE]%%solve(t(X[,
    A[i,] == 1, drop = FALSE])%X[, A[i,] == 1,
    drop = FALSE])%t(X[, A[i,] == 1, drop = FALSE])
    SS=t(y)%%(diag(1, nrow = n)-Hg)%y
    sigma2=1/rgamma(t, (nu_0+n)/2, (nu_0*sigma2_0+SS)/2)
```

```

    Vb=g*solve(t(X[, A[i,] == 1, drop = FALSE])%*%X[, A[i,] == 1,
drop = FALSE])/(g+1)
    Eb=Vb%*%t(X[, A[i,] == 1, drop = FALSE])%*%y
    E=matrix(rnorm(sum(A[i, ]), 0, sqrt(sigma2)), 1, sum(A[i, ]))
    B[i, A[i,] == 1]=t(t(E%*%chol(Vb))+c(Eb))
  }
}
for (i in 1:m)
{
  print(sum(B[,i]!= 0)/t)
}
for (i in 1:m)
{
  print(quantile(B[,i],c(0.025, 0.975)))
}
plot(density(B[,2]),col=1, xlab=names(azdiabetes)[1],
main ="Problem 1 Part b")
abline(v=mean(B[,2]),col=2)
abline(v=quantile(B[,2],0.025),col=4)
abline(v=quantile(B[,2],0.975),col=4)
for (i in 2:6)
{
  plot(density(B[,i+1]),col=1, xlab=names(azdiabetes)[i+1],
main ="Problem 1 Part b")
  abline(v=mean(B[,i+1]),col=2)
  abline(v=quantile(B[,i+1],0.025),col=4)
  abline(v=quantile(B[,i+1],0.975),col=4)
}

```

Output:

Table 2: Problem 1 Part b

	$Pr(\beta_j \neq 0 y)$	95% CI
(Intercept)	1	(42.650, 76.874)
npreg	0.149	(-1.1151, 0.000)
bp	0.291	(0.000, 0.369)
skin	0.158	(0.000, 0.411)
bmi	0.989	(0.469, 1.347)
ped	0.899	(0.000, 17.561)
age	1	(0.480, 0.993)

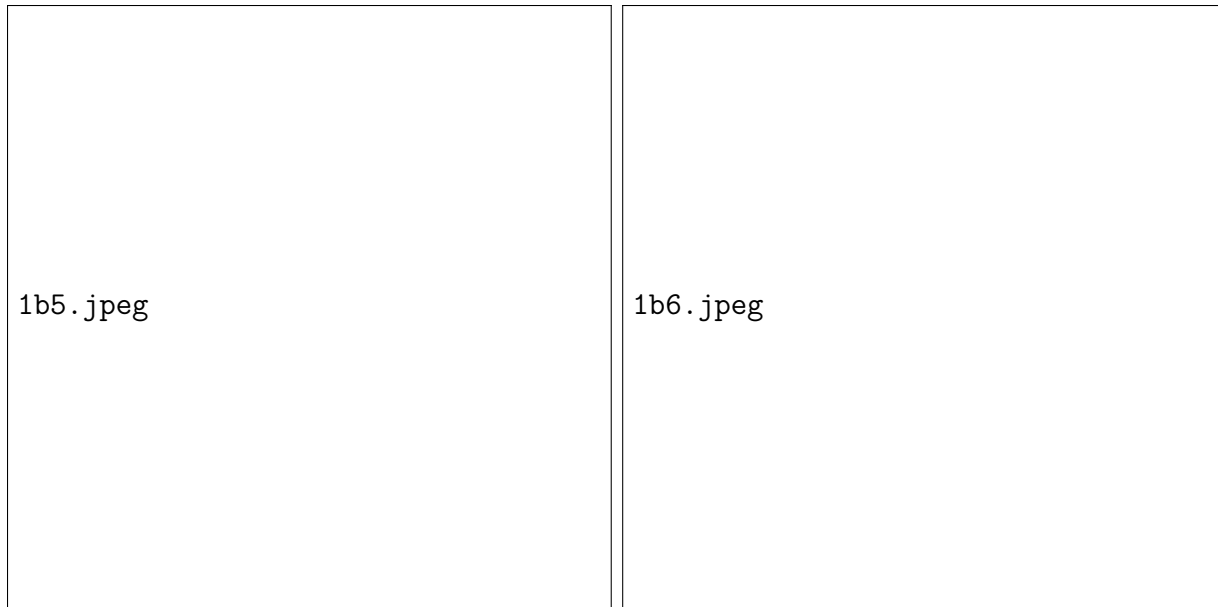
Posterior distribution plots *Blue lines correspond to the 95% confidence intervals*

1b1.jpeg

1b2.jpeg

1b3.jpeg

1b4.jpeg



Problem 2. *Problem 9.3 from Hoff*

Part a

We fit the regression model using the g-prior with $g = 47$, $v_0 = 2$, and $\sigma_0^2 = 1$. The left side of the table below gives the marginal posterior means and confidence intervals for all the variables. The right side gives the summary of the least squares estimates for each variable, with the mean, confidence interval, and p-values. Comparing the two sides of the table, we can see that the confidence intervals for the least squares approach is generally wider than that for the Bayesian approach. But their means are rather similar. The plots for the posterior distributions for each variable is also give below, with the means (blue) and 95% confidence intervals (green). The R code is also shown below

There are a couple of ways to determine whether a variable is strongly predictive of crime rates. If 0 is not contained in its 95% confidence interval, by the Bayesian approach, then it implies that it has a noticeable effect on crime rates. Another indicator of significant effect on crime rate is the p-value. From the least squares approach, if the p-value is small, it has a significant effect on the crime rate. Both methods were used (denoted by * in Table 3).

It is seen that the strong predictors of crime rates are **M** (percentage of males aged 14–24), **Ed** (mean years of schooling), **U2** (unemployment rate of urban males 3539), **Ineq** (income inequality), and **Prob** (probability of imprisonment). Two of the variables with the lowest p-values are **Ed** and **Ineq**. They are the two most significant positively-correlated variables with crime rates. The most significant negatively-correlated variable is **Prob**, which is intuitively clear. Indeed, a higher probability of imprisonment usually leads to lower crime rates.

R Code

```

#Problem 2 Part a
crime = read.table("crime.dat", header = TRUE)
y = as.matrix(crime[,1])
X = as.matrix(cbind(rep(1, dim(y)[1], 1), crime[, -1]))
g = length(y)
nu_0 = 2
sigma2_0 = 1
n = dim(X)[1]
m = dim(X)[2]
t = 10000
Hg = (g/(g+1))*X%*%solve(t(X)%*%X)%*%t(X)
SSR = t(y)%*%(diag(1, nrow = n)-Hg)%*%y
sigma2 = 1/rgamma(t, (nu_0+n)/2, (nu_0*sigma2_0+SSR)/2)
Vb = g*solve(t(X)%*%X)/(g+1)
Eb = Vb%*%t(X)%*%y
E = matrix(rnorm(t*m, 0, sqrt(sigma2)), t, m)
beta = t(t(E%*%chol(Vb))+c(Eb));
par(mfrow = c(2, 2))
beta_mean=numeric(m)
beta_l=numeric(m)
beta_u=numeric(m)
for (i in 1:m) {
  plot(density(beta[,i]), col = 2, lwd = 1,
       xlab = names(crime)[i], main = "Problem 2 part a");
  abline(v = quantile(beta[,i], c(0.025, 0.975))[1],
        col = 3, lwd = 2);
  abline(v = quantile(beta[,i], c(0.025, 0.975))[2],
        col = 3, lwd = 2);
  abline(v = mean(beta[,i]), col = 4, lwd = 2);
  beta_mean[i]=mean(beta[,i])
  beta_l[i]=quantile(beta[,i], c(0.025, 0.975))[1]
  beta_u[i]=quantile(beta[,i], c(0.025, 0.975))[2]
}
beta_mean
beta_l
beta_u
# Least squares estimates
LS = lm(y ~ M+So+Ed+Po1+Po2+LF+M.F+Pop+NW+U1+U2+GDP+Ineq+Prob+Time,
data = crime)
summary(LS)
MSE = sum(LS$residuals^2)/(n-16);
#upper 95% bound
LS$coefficients+qt(1-0.05/2, df=n-16)*sqrt(diag(MSE*solve(t(X)%*%X)))
#lower 95% bound
LS$coefficients-qt(1-0.05/2, df=n-16)*sqrt(diag(MSE*solve(t(X)%*%X)))

```

Output:

Table 3: The marginal posterior means and 95% confidence intervals for all regressors

Regressor	Bayes		Least Square		
	Mean	95% confidence interval	Mean	95% confidence interval	p-value
Intercept	-0.000166	[-0.1420, 0.1426]	-0.0005	[-0.1614, 0.1605]	0.9954
M	0.2818	[0.0363, 0.52665]	0.2865	[0.0096, 0.5635]	0.0430*
So	0.00142	[-0.3329, 0.3323]	-0.0001	[-0.3755, 0.3753]	0.9995
Ed	0.5330	[0.2063, 0.8529]	0.5445	[0.1782, 0.9108]	0.0049**
Po1	1.4417	[-0.0403, 2.8997]	1.4716	[-0.1939, 3.1372]	0.0813
Po2	-0.7647	[-2.2790, 0.7943]	-0.7818	[-2.5186, 0.9551]	0.3657
LF	-0.0628	[-0.3365, 0.2185]	-0.0660	[-0.3789, 0.2470]	0.6703
M.F	0.1277	[-0.1560, 0.4041]	0.1313	[-0.1852, 0.4478]	0.4040
Pop	-0.0696	[-0.2947, 0.1600]	-0.0703	[-0.3291, 0.1886]	0.5837
NW	0.1061	[-0.2090, 0.4189]	0.1091	[-0.2416, 0.4597]	0.5051
U1	-0.2644	[-0.6155, 0.0947]	-0.2705	[-0.6716, 0.1305]	0.1787
U2	0.3617	[0.0386, 0.6896]	0.3687	[0.0019, 0.7356]	0.0489*
GDP	0.2303	[-0.2397, 0.7015]	0.2381	[-0.2901, 0.7662]	0.3650
Ineq	0.7096	[0.2949, 1.1350]	0.7263	[0.2487, 1.2038]	0.0041**
Prob	-0.2796	[-0.5195, -0.0362]	-0.2852	[-0.5581, -0.124]	0.0411*
Time	-0.0598	[-0.2984, 0.1729]	-0.0616	[-0.3288, 0.2056]	0.6417

Posterior distribution plots *Green lines correspond to the 95% confidence intervals and Blue line corresponds to the mean*

Part b (i)

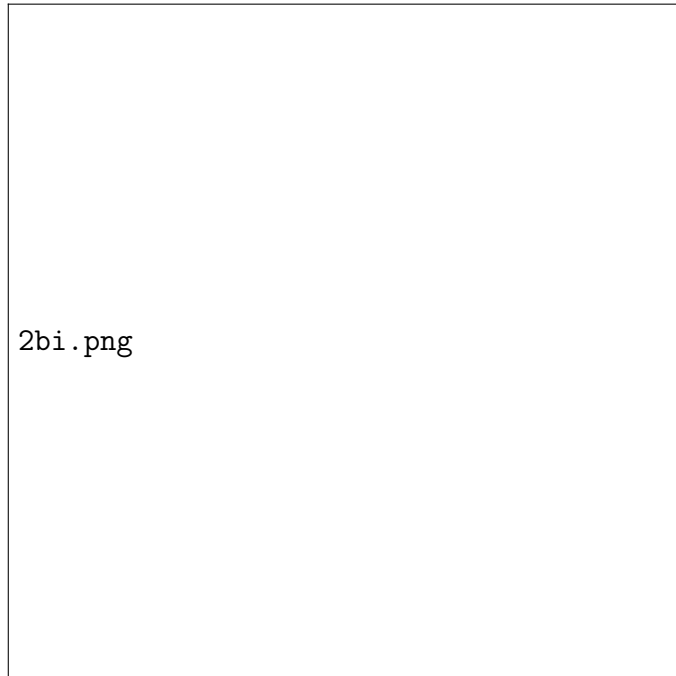
After randomly dividing the data into training and test data sets, we compute \hat{y}_{ols} and subsequently plot it against y_{te} (below). The prediction error was found to be 0.5562.

R code and output on the next page

R Code

```
> #Problem 2 Part b (i)
>
> set.seed(1)
> tr = sample(1:n, n/2)
> te = (1:n)[-tr]
> X_tr = X[tr,]
> X_te = X[te,]
> y_tr = y[tr,]
> y_te = y[te,]
> beta_ols = solve(t(X_tr)%*%X_tr)%*%t(X_tr)%*%y_tr
> y_ols = X_te%*%beta_ols
> dev.off()
> plot(y_te, y_ols, xlab = "y_te", ylab = "y_ols",
main='Problem 2 part b (i)', xlim = c(-2, 4), ylim = c(-2, 4),
asp = 1);
> abline(c(0, 0), c(1, 1), col = "red")
> sum((y_ols-y_te)^2)/length(y_te)
[1] 0.5562224
```

Plot:



Part b (ii)

Now we are going to use the g-prior from part a and use only the training set to compute $y_{Bayes}^{\hat{}}$. Then we plot versus the test data (below). The prediction error was found to be

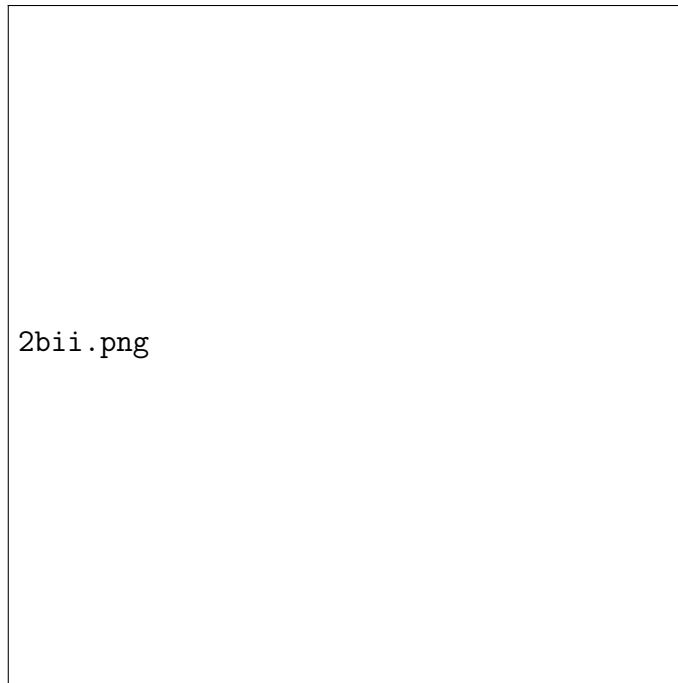
0.5219, which is smaller than the OLS prediction error determined in part a.

R code and output on the next page

R Code

```
> #Problem 2 Part b (ii)
>
> g = length(y_tr);
> nu_0 = 2;
> sigma2_0 = 1;
> n = dim(X_tr)[1];
> m = dim(X_tr)[2];
> t = 10000;
> Hg = (g/(g+1))*X_tr%%solve(t(X_tr)%X_tr)%t(X_tr);
> SSR = t(y_tr)%%(diag(1, nrow = n)-Hg)%y_tr;
> sigma2 = 1/rgamma(t, (nu_0+n)/2, (nu_0*sigma2_0+SSR)/2);
> Vb = g*solve(t(X_tr)%X_tr)/(g+1);
> Eb = Vb%%t(X_tr)%y_tr;
> E = matrix(rnorm(t*m, 0, sqrt(sigma2)), t, m);
> beta_bayes = colMeans(t(t(E%%chol(Vb))+c(Eb)));
> y_bayes = X_te%%beta_bayes;
> plot(y_te, y_bayes, xlab = "y_te", ylab = "y_bayes",
main='Problem 2 part b (ii)', xlim = c(-2, 4), ylim = c(-2, 4),
asp = 1);
> abline(c(0, 0), c(1, 1), col = "red");
> sum((y_bayes-y_te)^2)/length(y_te);
[1] 0.5218922
```

Plot:



2bii.png

Part c

The above procedures in part b were repeated 10,000 times and an average prediction error was subsequently computed for the OLS method and was found to be 1.2162 . Using Bayesian methods, the average prediction error was found to be 1.1456. Again, we see that the average prediction error when implementing Bayesian methods is lower throughout 10000 randomly generated test sets.

R code and output below

R Code

```
> #Problem 2 part c
> t1 = 10000;
> error_ols = rep(NA, 1, t1);
> error_bayes = rep(NA, 1, t1);
> for (i in 1:t1) {
+   set.seed(i);
+   n = 47;
+   tr = sample(1:n, n/2);
+   te = (1:n)[-tr];
+   X_tr = X[tr,];
+   X_te = X[te,];
+   y_tr = y[tr,];
+   y_te = y[te,];
+   beta_ols = solve(t(X_tr)%*%X_tr)%*%t(X_tr)%*%y_tr;
+   y_ols = X_te%*%beta_ols;
+   error_ols[i] = sum((y_ols-y_te)^2)/length(y_te)
+   g = length(y_tr);
+   nu_0 = 2;
+   sigma2_0 = 1;
+   n = dim(X_tr)[1];
+   m = dim(X_tr)[2];
+   t = 10000;
+   Hg = (g/(g+1))*X_tr%*%solve(t(X_tr)%*%X_tr)%*%t(X_tr);
+   SSR = t(y_tr)%*%(diag(1, nrow = n)-Hg)%*%y_tr;
+   sigma2 = 1/rgamma(t, (nu_0+n)/2, (nu_0*sigma2_0+SSR)/2);
+   Vb = g*solve(t(X_tr)%*%X_tr)/(g+1);
+   Eb = Vb%*%t(X_tr)%*%y_tr;
+   E = matrix(rnorm(t*m, 0, sqrt(sigma2)), t, m);
+   beta_bayes = colMeans(t(t(E%*%chol(Vb))+c(Eb)));
+   y_bayes = X_te%*%beta_bayes;
+   error_bayes[i] = sum((y_bayes-y_te)^2)/length(y_te);
+ }
> mean(error_ols)
[1] 1.216266
> mean(error_bayes)
[1] 1.145592
```

Problem 3. *Problem 10.2 from Hoff*

Part a

We know that Y_i is a binary indicator r.v. corresponding to whether or not the i^{th} sparrow successfully nests, so we have that

$$p(y_i|\alpha, \beta, x_i) = p_i^{y_i} (1 - p_i)^{1-y_i}$$

The joint sampling distribution is then

$$\begin{aligned} \prod_{i=1}^n p(y_i|\alpha, \beta, x_i) &= \prod_{i=1}^n p_i^{y_i} (1 - p_i)^{1-y_i} = \prod_{i=1}^n \exp \left\{ y_i \log\left(\frac{p_i}{1 - p_i}\right) + \log(1 - p_i) \right\} \\ &= \exp \left\{ \sum_{i=1}^n y_i \log\left(\frac{p_i}{1 - p_i}\right) + \sum_{i=1}^n \log(1 - p_i) \right\} \end{aligned}$$

Since we are using the logit link, $\log\left(\frac{p_i}{1-p_i}\right) = \alpha + \beta x_i$, we have

$$\prod_{i=1}^n p(y_i|\alpha, \beta, x_i) = \exp \left\{ \sum_{i=1}^n y_i (\alpha + \beta x_i) - \sum_{i=1}^n \log(1 + \exp(\alpha + \beta x_i)) \right\}$$

Part b

Using logistic regression, we can determine the prior distributions for α and β as follows
R code below

$$\begin{aligned} \alpha &\sim N(\mu_\alpha, \sigma_\alpha^2), \text{ where } \mu_\alpha = -9.78681, \sigma_\alpha = 4.620189 \\ \beta &\sim N(\mu_\beta, \sigma_\beta^2), \text{ where } \mu_\beta = 0.7757703, \sigma_\beta = 0.3578308 \end{aligned}$$

R Code

```
#Problem 3 Part b
msparrownest <- read.table("msparrownest.dat", quote="\\"",
  comment.char="")
y=msparrownest$V1
x=msparrownest$V2
f1=glm(formula = y~x,family = binomial, data=msparrownest)
mu.alpha=summary(f1)$coefficients[1,1]
mu.alpha
sd.alpha=summary(f1)$coefficients[1,2]
sd.alpha
mu.beta=summary(f1)$coefficients[2,1]
mu.beta
sd.beta=summary(f1)$coefficients[2,2]
sd.beta
```

Part c

A Metropolis algorithm was implemented using 10^7 repetitions using the R code below. We achieved an acceptance rate of 0.3003839. The effective sample size for α is 1,055 and the effective sample size for β is 1,107, so both are greater than 1,000, as desired.

R Code

```
#Problem 3 Part c
posterior=function(alpha,beta,y,x){
  b=sum(log(1+exp(alpha+beta*x)))
  theta=alpha+beta*x
  likelihood=exp(sum(y*theta)-b)
  prior.alpha=dnorm(alpha, mean=mu.alpha, sd=sd.alpha)
  prior.beta=dnorm(beta, mean=mu.beta, sd=sd.beta)
  return(likelihood*prior.alpha*prior.beta)
}

#Metropolis
B=20000
n=10000000
alpha.m=rep(NA,B+n)
beta.m=rep(NA,B+n)
alpha.m[1]=mu.alpha
beta.m[1]=mu.beta
acc=0
for ( t in 2:(B+n)){
  alpha.new=rnorm(1,mean=alpha.m[t-1],sd=.1)
  beta.new=rnorm(1,mean=beta.m[t-1],sd=.1)
  pi.new=posterior(alpha.new, beta.new, y,x)
  pi.t=posterior(alpha.m[t-1], beta.m[t-1], y,x)
  rho=min(c(pi.new/pi.t,1))
  U=runif(1,min=0,max=1)
  if (U<rho){
    alpha.m[t]=alpha.new
    beta.m[t]=beta.new
    acc=acc+1
  } else{
    alpha.m[t]=alpha.m[t-1]
    beta.m[t]=beta.m[t-1]
  }
}

print(acc/(B+n))
effectiveSize(alpha.m)
effectiveSize(beta.m[1:((B+n)/2)]) +
  effectiveSize(beta.m[((B+n)/2+1):(B+n)])
```

Part d

The posterior densities (black curve) and prior densities (blue curve) for α and β were plotted in R (Plots below). Referring to the plots, we can see that the posterior and prior densities for both α and β have significant differences. The posterior distributions have higher peaks and thinner tails compared to the prior distributions (as seen below), which was to be expected. However, it appears that they have roughly the same mean, for both α and β .

R Code

```
#Problem 3 Part d
grapha=density(alpha.m)
graphb=density(beta.m)
prior.alpha=dnorm(grapha$x, mean=mu.alpha, sd=sd.alpha)
prior.beta=dnorm(graphb$x, mean=mu.beta, sd=sd.beta)
plot(grapha$x,grapha$y,lwd=2,typ="l", xlab = "alpha",
      ylab = "density", main = "Prior and Posterior for alpha")
lines(grapha$x,prior.alpha,col="blue",lwd=2)
legend("topright",legend = c("prior","posterior"),lwd=c(1,1),
      col=c("blue","black"))
plot(graphb$x,graphb$y,typ="l",xlab = "beta",ylab = "density",
      main = "Prior and Posterior for beta")
lines(graphb$x,prior.beta,col="blue")
legend("topright",legend = c("prior","posterior"),lwd=c(1,1),
      col=c("blue","black"))
```


Part e

A 95% confidence band for the given function $f_{\alpha\beta}(x)$ of wingspan is given below

R Code

```
#Problem 3 Part e
x.seq=seq(10,15,by=0.02)
conf.band=matrix(NA,nrow=length(x.seq),ncol=2)
for (i in 1:length(x.seq)) {
  fun=exp(alpha.m+beta.m*x.seq[i])
  fun1=fun/(1+fun)
  conf.band[i,]=quantile(fun1,probs=c(0.025, 0.975), na.rm = TRUE)
}
plot(x.seq,conf.band[,1],ylim=c(0,1),lwd=2,col="black",type="l",
     main="Confidence Band",xlab="x",ylab="Percentile")
points(x.seq,conf.band[,2],col="red",type="l")
legend("topleft", col=c("black","red"),
      legend=c("lower quantile","upper quantile"), lwd = c(1,1))
```

Problem 4.

Part 1

We first compute the conditional distributions as follows:

$$\beta|rest \sim MVN \left(\left(\frac{I}{100} + \frac{X^T X}{\sigma^2} \right)^{-1} \frac{X^T z}{\sigma^2}, \left(\frac{I}{100} + \frac{X^T X}{\sigma^2} \right)^{-1} \right)$$

$$p(z_i|rest) \propto \exp[y_i z_i - \log(1 + e^{z_i}) - \frac{1}{2\sigma^2}(z_i - x_i^T \beta)^2] \quad i = 1, 2, \dots, n$$

We now set $\sigma^2 = 1$ and apply Metropolis Hastings algorithm. The R code and the output/results are given below.

```
#Problem 4 Part 1
library(LearnBayes)
library(MASS)
data(donner)

y=as.matrix(donner$survival,ncol=1)
X=cbind(1,donner$age,donner$male)
n=dim(X)[1]
p=dim(X)[2]
X=as.matrix(X,nrow=n,ncol=p)
Sigma=solve(diag(x=1,nrow=p,ncol=p)/100+t(X) %*% X)
#Gibbs
B=1000
```

```

nmc=10000
beta.mc=matrix(NA,nrow=p,ncol=nmc+B)
z.mc=matrix(NA,nrow=n,ncol=nmc+B)
fit.freq=glm(donner$survival~X-1,family=binomial)
beta.mc[,1]=fit.freq$coefficients
z.mc[,1]=X%%beta.mc[,1]
zi.cond<-function(zi,xi,beta,sigma,yi){
  a=dnorm(zi,mean=t(xi)%%beta,sd=sigma)*exp(yi*zi-log(1+exp(zi)))
  return(a)
}

for (t in 2:(B+nmc)){
  mu.beta=Sigma %*% t(X) %*% as.matrix(z.mc[,t-1],ncol=1)
  beta.mc[,t]=mvrnorm(1,mu=mu.beta,Sigma)
  pz.s=mvrnorm(1,mu=z.mc[,t-1],Sigma=diag(x=1,nrow=n,ncol=n))
  for (i in 1:n){
    xi=as.matrix(X[i,],ncol=1)
    beta=as.matrix(beta.mc[,t],ncol=1)
    yi=y[i]
    pz.s=zi.cond(pz.s[i],xi,beta,sigma=1,yi)
    pz.t=zi.cond(z.mc[i,t-1],xi,beta,sigma=1,yi)
    rho=min(c(1,pz.s/pz.t))
    if (runif(1,min=0,max=1)<rho)
    { z.mc[i,t]=pz.s[i]
    }else{
      z.mc[i,t]=z.mc[i,t-1]
    }
  }
}

#Posterior results for logistic model
beta.mc=beta.mc[, (B+1):(B+nmc)]
beta.mc1.mean=apply(beta.mc,1,mean)
beta.mc1.lower=apply(beta.mc,1,quantile,.025)
beta.mc1.upper=apply(beta.mc,1,quantile,.975)
print(beta.mc1.mean)
print(beta.mc1.upper)
print(beta.mc1.lower)

#Probit model Results
beta.probit=bayes.probit(y,X,m=(nmc+B),prior=list(beta=c(0,0,0),
P=diag(x=1,nrow=p,ncol=p)/100))$beta
beta.probit=beta.probit[(B+1):(B+nmc),]
beta.probit.mean=apply(beta.probit,2,mean)
beta.probit.lower=apply(beta.mc,1,quantile,.025)
beta.probit.upper=apply(beta.mc,1,quantile,.975)
print(beta.probit.mean)

```

```
print(beta.probit.upper)
print(beta.probit.lower)
```

The results of logistic posterior means and credible intervals for β are given in the table below:

	Mean	Lower 2.5% quantiles	Upper 2.5% quantiles
Intercept	4.1656533	1.2234501	8.05297357
Age	-0.1039775	-0.2035014	-0.02211799
Male	-1.9720865	-4.0257651	-0.25426965

The results of probit regression are summarized in the table below:

	mean	lower 2.5% quantiles	upper 2.5% quantiles
Intercept	2.05396256	1.2234501	8.05297357
Age	-0.04961277	-0.2035014	-0.02211799
Male	-0.98946242	-4.0257651	-0.25426965

It is seen that the posterior means for β are significantly different between the two models.

Part 2

First consider all of the conditionals, including the new conditional for σ^2 :

$$\beta|rest \sim MVN\left(\left(\frac{I}{100} + \frac{X^T X}{\sigma^2}\right)^{-1} \frac{X^T z}{\sigma^2}, \left(\frac{I}{100} + \frac{X^T X}{\sigma^2}\right)^{-1}\right)$$

$$p(z_i|rest) \propto \exp[y_i z_i - \log(1 + e^{z_i}) - \frac{1}{2\sigma^2}(z_i - x_i^T \beta)^2] \quad i = 1, 2, \dots, n$$

$$\sigma^2|rest \sim Inv - Gamma(a + \frac{n}{2}, b + \frac{1}{2}(z - X\beta)^T(z - X\beta)), \quad \text{where } z = (z_1, \dots, z_n)^T$$

The R code and the results of logistic posterior means and credible intervals for β are given below.

R Code

```
#Problem 4 Part 2
B=2000
nmc=20000
beta.mc=matrix(NA,nrow=p,ncol=nmc+B)
z.mc=matrix(NA,nrow=n,ncol=nmc+B)
beta.mc[,1]=fit.freq$coefficients
z.mc[,1]=X%*%beta.mc[,1]
sigma2.mc=rep(NA,B+nmc)
sigma2.mc[1]=1
count=rep(0,n)

#Gibbs Sampling:
for (t in 2:(B+nmc)){
```

```

Sigma.beta=solve(diag(x=1,nrow=p,ncol=p)/100+t(X) %*%
X/sigma2.mc[t-1])
mu.beta=Sigma.beta %*% t(X) %*% as.matrix(z.mc[,t-1],
ncol=1)/sigma2.mc[t-1]
beta.mc[,t]=mvrnorm(1,mu=mu.beta,Sigma.beta)
pz.s=mvrnorm(1,mu=z.mc[,t-1],Sigma=0.01*diag(1,ncol=n,nrow=n))
for (i in 1:n){
  xi=as.matrix(X[i,],ncol=1)
  beta=as.matrix(beta.mc[,t],ncol=1)
  yi=y[i]
  pz.s=zi.cond(pz.s[i],xi,beta,sigma=sqrt(sigma2.mc[t-1]),yi)
  pz.t=zi.cond(z.mc[i,t-1],xi,beta,sigma=sqrt(sigma2.mc[t-1]),yi)
  rho=min(c(1,pz.s/pz.t))
  if (runif(1,min=0,max=1)<rho){
    z.mc[i,t]=pz.s[i]
    count[i]=count[i]+1
  }
  else{z.mc[i,t]=z.mc[i,t-1]}
}
SSR.t = sum((z.mc[,t] - X %*% as.matrix(beta.mc[,t],ncol=1))^2)
sigma2.mc[t] = 1/rgamma(1,shape=0.1+n/2,rate=1+SSR.t/2)
}
beta.mc=beta.mc[, (B+1):(B+nmc)]

#Posterior results for logistic model
beta.mc2.mean=apply(beta.mc,1,mean)
beta.mc2.upper=apply(beta.mc,1,quantile,.975)
beta.mc2.lower=apply(beta.mc,1,quantile,.025)
print(beta.mc2.mean)
print(beta.mc2.upper)
print(beta.mc2.lower)

#Probit model Results
beta.probit=bayes.probit(y,X,m=(nmc+B),prior=list(beta=c(0,0,0),
P=diag(x=1,nrow=p,ncol=p)/100))$beta
beta.probit=beta.probit[(B+1):(B+nmc),]
beta.probit.mean=apply(beta.probit,2,mean)
beta.probit.upper=apply(beta.mc,1,quantile,.975)
beta.probit.lower=apply(beta.mc,1,quantile,.025)
print(beta.probit.mean)
print(beta.probit.upper)
print(beta.probit.lower)

```

	mean	lower 2.5% quantiles	upper 2.5% quantiles
Intercept	4.633564	1.721778	7.54241823
Age	-0.111619	-0.184988	-0.03621309
Male	-2.471943	-4.656946	-0.45278008

The results of probit regression:

	mean	lower 2.5% quantiles	upper 2.5% quantiles
Intercept	2.05353078	1.721778	7.54241823
Age	-0.04951136	-0.184988	-0.03621309
Male	-0.99562403	-4.656946	-0.45278008

While the results from Part 1 and Part 2 are similar, the two models produce significantly different posterior means.

Problem 5.

1. We are given that $Y_i \sim \text{Poisson}(\lambda_i)$, so

$$p(y_i|\lambda_i) = \exp(-\lambda_i) \frac{\lambda_i^{y_i}}{y_i!} = \exp\{y_i \log \lambda_i - \lambda_i - \log(y_i!)\}$$

From the above it is seen that the canonical link is $\theta_i = \log \lambda_i$.

Using this canonical link, $\theta_i = \eta_i = x_i^T \beta = \log(\lambda_i) \implies \lambda_i = \exp(x_i^T \beta)$

Plugging this into the conditional density, we get

$$p(y_i|x_i, \beta) = \exp\{y_i x_i^T \beta - \log(y_i!) - \exp(x_i^T \beta)\}$$

Then the likelihood can be computed as follows

$$p(y|x, \beta) = \prod_{i=1}^n p(y_i|x_i, \beta) = \exp\left\{\sum_{i=1}^n y_i x_i^T \beta - \sum_{i=1}^n \log(y_i!) - \sum_{i=1}^n \exp(x_i^T \beta)\right\}$$

2. Using the logit link instead, we get $\theta_i = \log(\lambda_i) = \log(\eta_i) = \log(x_i^T \beta)$
 $\implies \lambda_i = x_i^T \beta$.

Plugging this into the conditional density:

$$p(y_i|x_i, \beta) = \exp(y_i \log(x_i^T \beta) - \log(y_i!) - x_i^T \beta)$$

The likelihood is then computed as follows:

$$p(y|x, \beta) = \prod_{i=1}^n p(y_i|x_i, \beta) = \exp\left\{\sum_{i=1}^n y_i \log(x_i^T \beta) - \sum_{i=1}^n \log(y_i!) - \sum_{i=1}^n x_i^T \beta\right\}$$