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Modeling Transport Mode Decisions Using Hierarchical Logistic Regression Models with Spatial and Cluster Effects

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Abstract

This work is motivated by a mobility study conducted in the city of Munich, Germany. The variable of interest is a binary response, which indicates whether public transport has been utilized or not. One of the central questions is to identify areas of low/high utilization of public transport after adjusting for explanatory factors such as trip, individual and household attributes. For the spatial effects a modification of a class of Markov Random Fields (MRF) models with proper joint distributions introduced by Pettitt et al. (2002) is developed. It contains the intrinsic MRF in the limit and allows for efficient Markov Chain Monte Carlo (MCMC) algorithms. Further cluster effects using group and individual approaches are taken into consideration. The first one models heterogeneity between clusters, while the second one models heterogeneity within clusters. A naive approach to include individual cluster effects results in an unidentifiable model. It is shown how a reparametrization gives identifiable parameters. This provides a new approach for modeling heterogeneity within clusters. Finally the proposed model classes are applied to the mobility study.

Key words: binary regression, spatial effects, group and individual cluster effects, MCMC, transport mode decisions

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1 Introduction

This work has been motivated by a German mobility study investigating the usage of public transport. Discrete choice models based on maximization of random utilities (McFadden (1984)) have been often used in investigating such travel mode decisions (Ben-Akiva and Lerman 1985, McFadden 2001 and Bhat (2006)) giving rise to the familiar multinomial logit model. For a binary choice this gives a logit model. Since the variable of interest is a binary indicator, whether public transport has been utilized or not, we base our models on a logit formulation. Early papers (McGillivray 1972 and McFadden 1974) on travel mode decisions also consider binary choices, but they do not include spatial components. The central question is to identify areas of low/high utilization of public transport after adjusting for trip, individual and household attributes. The goal is to develop flexible statistical models for a binary response with covariate, spatial and cluster effects. A large number of models are discussed in the literature which incorporate covariates together with spatial information. We provide now a short overview. In the context of general additive models, the simplest possibility to account for spatial information would be to use an additional nominal covariate indicating the region if there are multiple responses per region. But such an approach does not give a model for spatial dependence. This property is especially desired if the data volume is not large with respect to the number of covariates.

There are two approaches to incorporate spatial effects. The first one is appropriate for data collected at specified locations, while the second one uses data regions. The first approach is known as *generalized linear kriging* (Diggle et al. 1998). It is based on generalized linear mixed models (Breslow and Clayton 1993), where spatial random effects are modeled as realizations of a stationary Gaussian process with zero mean and a parameterized covariance structure. Diggle et al. (1998) use Markov Chain Monte Carlo (MCMC) methods for large data. Updating the covariance parameters is difficult since high dimensional matrix inversion and determinants are required at each iteration. Therefore Heagerty and Lele (1998) use restrictions on the dependence structure, while Gelfand et al. (2000) use importance sampling to avoid matrix inversions. The second approach is appropriate when spatial effects are associated with data regions. These do not need to be on a regular lattice. Now data are assumed to be aggregated over regions and spatial effects are modeled for each region instead for each observation. Here linear predictors are given as

$$\eta_i = \mathbf{x}_i'\boldsymbol{\alpha} + b_{j(i)}, \quad i = 1, \ldots, n,\ j = 1, \ldots, J\ ,$$

where J denotes the number of regions, $\mathbf{b} = (b_1, \cdots, b_J)'$ are spatial effects and $j(i)$ indicates the region associated with the i^{th} observation. Spatial effects are modeled as a realization from a Gaussian *Markov random field* (MRF) (Besag and Green (1993), Banerjee et al. (2004) and Rue and Held (2005)). The name *Gaussian conditional autoregression* (Gaussian CAR) is also used, since such a distribution is typically given through its full conditionals. This allows fast individual updating of $J << n$ spatial effects in a Gibbs sampler. This approach requires a spatial neighborhood structure. It is appropriate for our mobility application, since data are aggregated over postal codes of Munich, Germany. We consider two postal codes as neighbors if they have a joint border.

In contrast to stationary Gaussian processes used in kriging, in Gaussian CAR models the explicit form of its precision matrix (inverse covariance matrix) is available. Moreover this precision matrix is usually sparse, which allows to compute its determinant much faster than in the kriging approach. Pettitt et al. (2002) use this fact and propose a specific dependence structure which provides even an analytical computation of its determinant. Some Gaussian CAR models possess an improper joint density. The simplest example is the intrinsic CAR model (Besag and Green 1993), whose precision matrix is only semi positive. Fahrmeir and Lang (2001) used improper intrinsic CAR models as a prior in a semi parametric regression model for multi categorical time-space data, while Knorr-Held and Rue (2002) applied intrinsic CAR priors for Poisson models used in disease mapping. We study more advanced proper Gaussian CAR models with a parameterized correlation matrix. In particular, we develop a modification of the Pettitt's CAR model, which includes in the limit the intrinsic CAR model in contrast to the formulation by Pettitt et al. (2002). This modification still has all nice properties of the Pettitt et al. (2002) CAR models: proper joint distributions, a similar interpretation of parameters, the same conditional correlations and more important allows for fast computation of the determinant of the precision matrix, providing fast Gibbs sampling. An alternative proper Gaussian CAR model was also discussed in Sun et al. (2000). It also includes the intrinsic CAR model in the limit and allows for fast computation of the determinant of the precision matrix. It has been used to develop hierarchical spatio-temporal Poisson models for disease mapping data, but not for binary spatial responses. Gaussian CAR models will be considered in more detail in Section 2.

Another approach to regionally aggregated data is based on specifying the joint distribution of the spatial effects directly yielding simultaneous autoregressive (SAR) Gaussian models as introduced by Whittle (1954) and later studied by Cressie (1993) and Anselin (1988). Especially economists prefer the simultaneous approach for the analysis of spatial regional data (Anselin and Florax 1995 and Anselin and Florax 2004). Pinkse and Slade (1998) and McMillen (1992) consider a probit formulation with latent spatial regression following a SAR specification. While Pinkse and Slade (1998) use the generalized method of moments for estimation, McMillen (1992) employs the EM algorithm. Beron and Vijverberg (2004) use a simulator for multivariate normal probabilities to approximate the likelihood in a probit model with latent spatial SAR formulation to facilitate maximum likelihood estimation. LeSage (2000) gives a Bayesian analysis of probit and tobit models. A summary of these estimation methods are provided by Fleming (2004). In a recent paper Wall (2004) points out difficulties in interpreting the spatial dependence parameter for CAR and SAR models as spatial correlation parameters in nonregular lattices. We agree with N. Cressie (1993) and H. Rue and L. Held (2005) that CAR models are easier to be interpreted and do not consider SAR models in the following.

Finally, we mention auto logistic regression models (Besag 1974). Modeling the distribution of plant species Huffer and Wu (1998) propose to extend the auto logistic modeling of the success probability by incorporating a fixed effect term $\mathbf{x}_i'\boldsymbol{\alpha}$. They work on a regular rectangular lattice and one-observation-per-site data. But in spite of this simplicity Huffer and Wu (1998) note that exact MLE is not tractable, except when the number of sites is quite small, while two other estimation methods, namely the coding method (Besag (1974)) and the maximum pseudo-likelihood method (Besag (1975)), seem to be not sufficiently efficient. Huffer and Wu (1998) investigate a MCMC MLE approach, which produces the likelihood function via Monte Carlo simulations. They do not give any idea, how to account for possible interactions between species. For the Gaussian CAR approach Pettitt et al. (2002) solve this problem by modeling the correlation between several Gaussian CAR models for each species applied to tree biodiversity data. Also Carlin and Banerjee (2003) develop this approach for multiple cancer survival data. Further a multivariate extension of the proper Gaussian CAR model developed in Sun et al. (1999) and Sun et al. (2000) is considered by Gelfand and Vounatsou (2003) for multivariate continuous and multinomial response data. An overview of multivariate CAR models is provided by Jin et al. (2005). A different autologistic model for binary data in space and time was developed by Dubin (1995), who explored this model in a very small simulation set up. Here spatial interactions are modeled by a parameterized distance based weight matrix, while time dependence is captured by an autologistic formulation.

In addition to spatial effects we extend our modeling of the linear predictor η_i by cluster random effects. It allows us to take into account possible overdispersion caused by unobserved heterogeneity. We consider two approaches, namely group and individual cluster effects. The first one, which models heterogeneity between clusters, follows the usual idea of having the same random effect within a cluster. The second approach allows for heterogeneity within a cluster, i.e. we model cluster effects within a cluster as independent normally distributed random variables with zero mean and a cluster specific variance. For K clusters we have to estimate K cluster specific variances instead of K cluster effects as before. We will show how an unidentifiability problem occurring in the second case can be overcome. Efficient MCMC algorithms will be developed. In this paper we restrict our analysis to logit models with spatial and cluster effects. However Prokopenko (2004) also develops MCMC algorithms for probit formulations using a latent variable representation (Albert and Chib 1993).

2 Modeling of Spatial Effects Using Gaussian CAR Models

The most popular kind of Markov random fields (MRF) are Gaussian MRF's (Besag and Green 1993), or Gaussian CAR models (Pettitt et al. 2002), where a random vector $\mathbf{b} \in R^J$ is defined through its full conditionals as follows:

$$b_j|\mathbf{b_{-j}} \sim N\left(\mu_j + \sum_{j' \neq j} c_{jj'}(b_{j'} - \mu_{j'}),\, \kappa_j\right), \quad j = 1, \ldots, J\,.$$

Here $\mathbf{b_{-j}} = (b_1, \ldots, b_{j-1}, b_{j+1}, \ldots, b_J)^t$ and $N(\mu, \sigma^2)$ denotes a normal distribution with mean μ and variance σ^2. Besag and Green (1993) show that the joint distribution of a zero-mean Gaussian CAR is given by $\mathbf{b} \sim N_J\,(0,\,(I_J - C)^{-1}M)$, where $C = (c_{jj'})$ with $c_{jj} = 0$, $j = 1, \ldots, J$, and $M = \text{diag}(\kappa_1, \ldots, \kappa_J)$. Here $N_J(\mu, \Sigma)$ denotes a J-dimensional normal distribution with mean vector μ and covariance matrix Σ. The precision matrix is given by $Q = M^{-1}(I_J - C)$. Below we present examples of Gaussian CAR models. Further we assume that the neighborhood structure has no isolated regions or groups of regions.

Example 1: The *intrinsic Gaussian CAR* (Besag and Green 1993) is defined by:

$$b_j|\mathbf{b_{-j}} \sim N(\overline{b_j}, \frac{\tau^2}{N_j})\,, \quad j = 1, \ldots, J, \quad \text{and} \quad \overline{b_j} = \frac{\sum_{j \sim j'} b_{j'}}{N_j}\,, \tag{2.1}$$

where $N_j = \#$ of neighbors of the region j, and "$j \sim j'$" denotes contiguous regions. In particular, we have $j \nsim j$. The corresponding precision matrix is positive semi-definite with rank $= J - 1$, therefore $\mathbf{b}$ has an improper density, but can be characterized (see Prokopenko 2004).

Example 2: Pettitt et al. (2002) use a particular Gaussian CAR, where

$$b_j|\mathbf{b_{-j}} \sim N\left(\frac{\phi}{1 + |\phi| N_j} \sum_{j \sim j'} b_{j'}\,,\, \frac{\tau^2}{1 + |\phi| N_j}\right)\,. \tag{2.2}$$

The parameter ϕ measures the strength of the spatial dependency. There is no spatial dependency, if $\phi = 0$. Since maximum likelihood estimation is intractable for this model MCMC methods have been used to estimate ϕ and τ^2. Pettitt et al. (2002) show that a fast and simple update of ϕ for a Gibbs Step given the vector $\mathbf{b}$ and τ^2 is available. In contrast to the intrinsic CAR, the joint distribution of $\mathbf{b}$ based on conditionals specified in (2.2) is a proper distribution, which leads to a proper posterior when used as a prior distribution. This will circumvent any problems in the Gibbs sampler arising from using an improper prior.

Example 3: We introduce now a modified Pettitt's CAR model, where the full conditionals for $\mathbf{b}$ are given as follows:

$$b_j|\mathbf{b_{-j}} \sim N\left(\frac{\phi}{1+|\phi|N_j}\sum_{j\sim j'} b_{j'}\,,\,\frac{(1+|\phi|)\tau^2}{1+|\phi|N_j}\right). \tag{2.3}$$

This (also proper) distribution differs from Pettitt's CAR (2.2) by the additional term $1+|\phi|$ in the numerator of the conditional variance. This allows us to have the intrinsic CAR (2.1) in the limit, when $\phi \to \infty$. Note that the conditional variance of $b_j|\mathbf{b_{-j}}$ decreases to τ^2/N_j as $|\phi|$ increases to infinity, while in the original model (2.2) this quantity decreases to zero, which is a very restrictive assumption. Further, this model has the same behavior as Pettitt's CAR (2.2) when ϕ goes to zero (no spatial dependency), and also all partial correlations between b_j and b_i given all the other sites are the same. In the modified Pettitt's model we can also achieve a simple update for ϕ. We write now $\tau^{-2} \times Q^{m.P}(\phi)$ for the precision matrix of the modified Pettitt's model (2.3). In particular $Q^{m.P}(\phi) = M^{-1}(\phi)(I_J - C(\phi))$, where $M(\phi) = diag(\frac{(1+|\phi|)}{1+|\phi|N_1}, \cdots, \frac{(1+|\phi|)}{1+|\phi|N_J})$ and $C(\phi) = (c_{jj'}(\phi))_{jj'=1,\cdots,J}$ with

$$c_{jj'}(\phi) = \begin{cases} \frac{\phi}{1+|\phi|N_j}, & \text{if } j \sim j' \\ 0, & \text{if } j \not\sim j',\, j = j' \end{cases}.$$

Each update of ϕ requires the computation of the determinant of $Q^{m.P}(\phi)$. With the reparametrization $\psi = \frac{\phi}{1+|\phi|}$ we follow a similar approach as in Pettitt et al. (2002). More precisely, if we define the diagonal matrix

$$D = \text{diag}(N_1 - 1, \ldots, N_J - 1) \quad \text{and} \quad \Gamma = (\gamma_{jj'})_{j,j'=1,\ldots,J} = \begin{cases} 1, & \text{if } j \sim j' \\ 0, & \text{if } j \not\sim j',\, j = j' \end{cases},$$

then $Q^{m.P}(\phi)$ can be written in the form $Q^{m.P}(\psi) = I_J + |\psi| D - \psi\Gamma$. If $(\lambda_1, \ldots, \lambda_J)$ are the eigenvalues of $\Gamma - D$ and $(\nu_1, \ldots, \nu_J)$ are the eigenvalues of $\Gamma + D$, then the determinant of $Q^{m.P}(\psi)$ is equal to

$$|Q^{m.P}(\psi)| = \begin{cases} \prod_j (1 - \psi\lambda_j), & \text{if } \psi > 0 \\ 1, & \text{if } \psi = 0 \\ \prod_j (1 - \psi\nu_j), & \text{if } \psi < 0 \end{cases}. \tag{2.4}$$

and can be computed quickly for any value of ψ. Finally we like to note that the conditional variance of $b_j|\mathbf{b_{-j}}$ is independent of the spatial dependence parameter for the proper Gaussian CAR model considered by Sun et al. (2000) in contrast to the modified Pettitt's CAR model (2.3). It is more reasonable to assume that this conditional variance increases as dependence among the spatial effects decreases. If $\phi = 0$, then the conditional variance in Sun et al. (2000) still depends on N_j, while this is not the case for the modified Pettitt's CAR model. Therefore we prefer the modified proper Pettitt's CAR model over the proper CAR model studied by Sun et al. (2000) for modeling spatial effects.

3 Spatial Logistic Regression Models with Group Cluster Effects

For the mobility study we use a binary response vector $\mathbf{Y} = (Y_1, \ldots, Y_n)^t$ with

$$Y_i = \begin{cases} 1 & \text{if trip i used individual transport} \\ 0 & \text{if trip i used public transport} \end{cases}, \quad i = 1, \cdots, n, \tag{3.1}$$

where Y_i's is Bernoulli with the success probabilities p_i and assume that Y_i given p_i are independent for $i = 1, \ldots, n$. We specify p_i through their logits :

$$\theta_i := \log\left(\frac{p_i}{1-p_i}\right) = \underbrace{\mathbf{x_i^t}\boldsymbol{\alpha}}_{\textit{fixed effect}} + \underbrace{b_{j(i)}}_{\textit{random spatial effect}} + \underbrace{c_{m(i)}}_{\textit{random group cluster effect}} . \tag{3.2}$$

Here the design vector $\mathbf{x_i}$ multiplied with the regression parameter vector $\boldsymbol{\alpha} \in \mathbb{R}^p$ represents fixed effects. With the vector $\mathbf{b} = (b_1, \ldots, b_J)$ we allow for random spatial effects. As sites we take $J = 74$ postal code areas of the city of Munich. Therefore, the index $j(i)$ denotes the residence postal code of the person who takes trip i. In order to be able to take into account possible spatial smoothness we assume, that b_j's arise from the modified Pettitt's CAR (2.3).

To model heterogeneity between clusters we allow for random cluster effects represented by the vector $\mathbf{c} = (c_1, \ldots, c_M)$. Each of the M clusters (say age groups or household types) induces a group specific random effect, which we denote by $c_m, m = 1, \ldots, M$, respectively. The index $m(i)$ denotes the cluster of trip i. We assume that $c_m \sim N(0, \sigma_c^2)$ *i.i.d.* for $m = 1, \ldots, M$.

Note that the likelihood of the response vector $\mathbf{Y}$ is proportional to

$$[\mathbf{Y}|\ \boldsymbol{\alpha},\ \mathbf{b},\ \mathbf{c}] \propto \prod_{i=1}^{n} \frac{\exp(Y_i(\mathbf{x_i}^t\boldsymbol{\alpha} + b_{j(i)} + c_{m(i)}))}{1 + \exp(\mathbf{x_i}^t\boldsymbol{\alpha} + b_{j(i)} + c_{m(i)})} .$$

Parameter	Prior specification
Regression	$\alpha_l \sim N(0, \sigma^2_{\alpha_l}),\ l = 1, \dots, p\ ind.\quad \sigma^2_{\alpha_l}$ large
Spatial	$b_j \vert \mathbf{b_{-j}} \sim N\left(\frac{\phi}{1+\vert\phi\vert N_j} \sum_{j\sim j'} b_{j'}\ ,\ \frac{(1+\vert\phi\vert)\tau^2}{1+\vert\phi\vert N_j}\right) j+1, \cdots J$
Spatial dependence	$\psi := \frac{\phi}{1+\vert\phi\vert} \in (-1, 1) \quad \pi(\psi) \propto \frac{1}{(1-\vert\psi\vert)^{1-a}},\ a > 0$
Spatial variance	$\pi(\tau^2) \propto 1$ or $\pi(\tau^2) = IG(a_\tau, b_\tau)$
Cluster	$c_m \vert \sigma^2_c \sim N(0, \sigma^2_c)\ i.i.d.$
Cluster variance	$\sigma^2_c \sim IG(a_c, b_c$ or $\pi(\sigma^2_c) \propto 1)$

Table 3.1: Prior distributions utilized in the spatial logistic regression model with group cluster effects

Since we will follow a Bayesian approach we need to complete the model specification by providing the prior specifications for $\boldsymbol{\alpha}, \phi, \tau^2$ and σ^2_c. For this we denote the density of a random variable X by $[X]$ and the conditional density of X given Y by $[X|Y]$, respectively. We assume independent prior distributions, i.e we assume $[\boldsymbol{\alpha},\, \mathbf{b},\, \mathbf{c},\, \phi, \tau^2, \sigma^2_c] \;=\; [\boldsymbol{\alpha}] \times [\mathbf{b}|\phi, \tau^2] \times [\phi] \times [\tau^2] \times [\mathbf{c}|\sigma^2_c] \times [\sigma^2_c]$. The specific priors for all parameters considered are given in Table 3.1. Here $IG(a, b)$ denotes the inverse gamma density given by $[x] = \frac{1}{b^a \Gamma(a) x^{a+1}} \exp(-\frac{1}{bx})$ for $x > 0$. MCMC methods allow us to draw an arbitrary large number of joint samples from the posterior distribution $[\boldsymbol{\alpha}, \mathbf{b}, \mathbf{c}, \phi, \tau^2, \sigma^2_c|\mathbf{Y}]$ approximately. With these samples we can make parameter inference using for example estimated posterior means or density estimates of the marginal posterior. Readers unfamiliar with MCMC methods can consult Chib (2001) for an introduction and Gilks et al. (1996) for applications of MCMC methods.

Individual Metropolis Hastings (MH) updates are used for the regression α_l, the spatial b_j and the cluster c_m parameters, since good joint proposal distributions are difficult to find. As individual proposal distributions we use a normal distribution with mean equal to the previous value and a fixed value for the standard deviation. This standard deviation is determined by pilot runs which resulted in an acceptance rate between 30-60% (as proposed in Bennett et al. (1996) or Besag et al. (1995)). They also serve as burnin phase. The reparameterized spatial hyperparameter $\psi = \frac{\phi}{1+|\phi|}$ also requires an MH Update.

The variance hyperparameters τ^2 and σ^2_c can be updated in a Gibbs step. For the full conditional of τ^2 we have $[\tau^2|\mathbf{Y}, \boldsymbol{\alpha}, \mathbf{b}, \mathbf{c}, \phi, \boldsymbol{\sigma^2_c}] = [\boldsymbol{\tau^2}|\mathbf{b}, \phi] \propto [\mathbf{b}|\phi, \boldsymbol{\tau^2}] \times [\boldsymbol{\tau^2}]$. Using an $IG(a_\tau, b_\tau)$ prior for τ^2 it is easy to see that $[\tau^2|\mathbf{Y}, \boldsymbol{\alpha}, \mathbf{b}, \mathbf{c}, \phi, \boldsymbol{\sigma^2_c}]$ is again $IG(a^*_\tau, b^*_\tau)$ with $a^*_\tau = a_\tau + \frac{J}{2}$ and $b^*_\tau = \left\{\frac{1}{b_\tau} + \frac{\mathbf{b}'\, Q(\phi)\, \mathbf{b}}{2}\right\}^{-1}$. When a flat improper prior for τ^2 is used (as we have chosen), the posterior $[\tau^2|\mathbf{b}, \phi]$ is $IG(a^*_\tau, b^*_\tau)$ with $a^*_\tau = \frac{J}{2} - 1$ and $b^*_\tau = \left\{\frac{1}{2}\, \mathbf{b}'\, Q(\phi)\, \mathbf{b}\right\}^{-1}$. Finally for the cluster variance σ^2_c if $\sigma^2_c \sim IG(a_c, b_c)$,

Parameter	Update
$\alpha_l,\ l = 1, \ldots, p$	Individual MH with normal RW proposal
$b_j,\ j = 1, \ldots, J$	Individual MH with normal RW proposal
$\psi = \frac{\phi}{1+\lvert\phi\rvert}$	MH Update with $uniform(-1, 1)$ proposal
τ^2	Gibbs Update, FC = IG(a^*_τ, b^*_τ)
$c_m,\ m = 1, \ldots, M$	Individual MH with normal RW proposal
σ^2_c	Gibbs Update, FC = IG(a^*_c, b^*_c)

Table 3.2: Updating Schemes of the MCMC algorithm for a Spatial Logistic Regression Model with Group Cluster Effects (MH = Metropolis Hastings step, RW = random walk, FC = full conditional)

the full conditional $\sigma^2_c|\mathbf{Y}, \boldsymbol{\alpha}, \mathbf{b}, \mathbf{c}, \phi, \tau^2$ is $IG(a^*_c, b^*_c)$ with $a^*_c = a_c + \frac{M}{2}$ and $b^*_c = \left\{\frac{1}{b_c} + \frac{\mathbf{c}'\mathbf{c}}{2}\right\}^{-1}$. For an improper prior the full conditional density for σ^2_c is a $IG(a^*_c, b^*_c)$ density with $a^*_c = \frac{M}{2} - 1$ and $b^*_c = \left\{\frac{1}{2}\,\mathbf{c}'\mathbf{c}\right\}^{-1}$. This density has a finite expectation for $M \geq 5$ and a finite variance for $M \geq 7$. A summary of these update schemes is given in Table 3.2.

4 Spatial Logistic Regression Models with Individual Cluster Effects

We consider now a more advanced model where individual cluster effects are modeled by a normal distribution with fixed variance inside each cluster given by:

$$\begin{aligned}
&Y_i|p_i \sim \text{Bernoulli}(p_i) \text{ conditionally independent with}\\
&\theta_i := \log\left(\frac{p_i}{1-p_i}\right) = \underbrace{\mathbf{x_i^t}\boldsymbol{\alpha}}_{\textit{fixed effect}} + \underbrace{b_{j(i)}}_{\textit{random spatial effect}} + \underbrace{c_{m(i),k(i)}}_{\textit{random individual cluster effect}},
\end{aligned} \tag{4.1}$$

where for fixed $m = 1, \ldots, M$, $c_{m,k} \sim N(0, \sigma^2_m)$, $k = 1, \ldots, K_m$, $i.i.d.$ As in Model (3.2), M denotes the number of clusters and $m(i)$ denotes the cluster of trip i. K_m stands for the number of trips, which belong to cluster m (i.e. $K_1 + \ldots + K_M = n$) and $k(i)$ gives the number of trip i in its cluster. The specification of the fixed effects $\boldsymbol{\alpha}$ and the spatial effects $\mathbf{b}$ remain as before. In contrast to (3.2), the cluster effects are now not the same for each trip in cluster m, namely c_m, but random realizations $c_{m,k},\ k = 1, \ldots, K_m$ from the same cluster distribution $N(0, \sigma^2_m)$. This allows for heterogeneity within each cluster.

In Model (4.1) we have to estimate in addition to the parameters $\boldsymbol{\alpha}, \mathbf{b}$ the cluster effect variances $\boldsymbol{\sigma^2} = (\sigma^2_1, \ldots, \sigma^2_M)^t$ instead of the cluster effects $\mathbf{c} = (c_1, \ldots, c_M)^t$ and their variance σ^2_c for Model (3.2). One problem with Model (4.1) is that even without an intercept term α_0 the model is unidentifiable. To understand this unidentifiability we first substitute in (4.1) the logit link function with the probit link function, i.e. we assume for $i = 1, \ldots, n$:

$$\begin{aligned} &Y_i|p_i \sim \text{Bernoulli}(p_i) \text{ conditionally independent with} \\ &p_i = \mathbf{P}\{Y_i = 1|\mathbf{x_i}, \boldsymbol{\alpha}, b_{j(i)}, c_{m(i),k(i)}\} = \Phi(\mathbf{x_i^t}\boldsymbol{\alpha} + b_{j(i)} + c_{m(i),k(i)}), \end{aligned} \tag{4.2}$$

where $\Phi(\cdot)$ is the standard normal distribution function. This allows for the latent variable representation (compare to Albert and Chib 1993):

$$\begin{aligned} Y_i = 1|\mathbf{x_i}, \boldsymbol{\alpha}, b_{j(i)}, \sigma^2_{m(i)} &\Leftrightarrow Z_i \le 0, \quad \text{where} \\ Z_i = -\eta_i + \epsilon_i^*,\ \epsilon_i^* &\sim N(0, 1 + \sigma^2_{m(i)}) \text{ independent and } \eta_i = \mathbf{x_i^t}\boldsymbol{\alpha} + b_{j(i)}\ . \end{aligned} \tag{4.3}$$

We have for $i = 1, \ldots, n$

$$\mathbf{P}\{Y_i = 1|\mathbf{x_i}, \boldsymbol{\alpha}, b_{j(i)}, \sigma^2_{m(i)}\} = \mathbf{P}\{Z_i \le 0|\mathbf{x_i}, \boldsymbol{\alpha}, b_{j(i)}, \sigma^2_{m(i)}\} = \Phi\left(\frac{\mathbf{x_i^t}\boldsymbol{\alpha} + b_{j(i)}}{\sqrt{1 + \sigma^2_{m(i)}}}\right). \tag{4.4}$$

Equation (4.4) shows that the parameters $\boldsymbol{\alpha}, \mathbf{b}$ and $\boldsymbol{\sigma^2}$ are not jointly identifiable in Model (4.2), since it is invariant with respect to the parameter vectors $\left\{k \times (\boldsymbol{\alpha}^t, \mathbf{b}^t, \sqrt{1+\sigma_1^2}, \ldots, \sqrt{1+\sigma_M^2}\,)^t,\ k \in \mathbb{R}\right\}$. If we define now

$$\boldsymbol{\alpha}' := \frac{\boldsymbol{\alpha}}{\sqrt{1+\sigma_1^2}}, \quad \mathbf{b}' := \frac{\mathbf{b}}{\sqrt{1+\sigma_1^2}}, \quad \sigma'^2_m := \frac{1+\sigma_m^2}{1+\sigma_1^2},\ m = 2, \ldots, M,\ \sigma'^2_1 = 1, \tag{4.5}$$

then the marginal distributions (4.4) of $Y_i|\mathbf{x_i}, \boldsymbol{\alpha}, b_{j(i)}, \sigma^2_{m(i)}$ from Model (4.2) will coincide with the marginal distributions from the following model:

$$\begin{aligned} &Y_i|p_i \sim \text{Bernoulli}(p_i) \text{ conditionally independent with} \\ &p_i = \mathbf{P}\{Y_i = 1|\mathbf{x_i}, \boldsymbol{\alpha}', b'_{j(i)}, \sigma'^2_{m(i)}\} = \begin{cases} \Phi\left(\mathbf{x_i^t}\boldsymbol{\alpha}' + b'_{j(i)}\right) & \text{if } m(i) = 1 \\ \Phi\left(\frac{\mathbf{x_i^t}\boldsymbol{\alpha}' + b'_{j(i)}}{\sigma'_{m(i)}}\right) & \text{if } m(i) = 2, \ldots, M\ . \end{cases} \end{aligned} \tag{4.6}$$

Using (4.3) it follows, that also the joint distribution of $\mathbf{Y}$ in both Models (4.2) and (4.6) are equal. Therefore Model (4.6) is an equivalent reparametrization of Model (4.2). But this representation (4.6) has one parameter less and is therefore identifiable. The above discussion helps us to understand the unidentifiability of logit Model (4.1), since the behavior of both probit and logit link functions is quite similar and they differ only significantly in the tails. So we use the same idea to construct an identifiable logit model. In particular we assume for $i = 1, \ldots, n$

$$\begin{aligned} &Y_i|p_i \sim \text{Bernoulli}(p_i) \text{ conditionally independent with} \\ &\log\left(\frac{p_i}{1-p_i}\right) = \begin{cases} \mathbf{x_i^t}\boldsymbol{\alpha}' + b'_{j(i)} & \text{if } m(i) = 1 \\ \frac{\mathbf{x_i^t}\boldsymbol{\alpha}' + b'_{j(i)}}{\sigma'_{m(i)}} & \text{if } m(i) = 2, \ldots, M \end{cases}, \end{aligned} \tag{4.7}$$

where $\boldsymbol{\alpha}', \mathbf{b}', \boldsymbol{\sigma}^{\mathbf{2}'} := (\sigma_1^{2'}, \ldots, \sigma_M^{2'})^t$ are defined as in (4.5). From (4.7) it follows that the likelihood of the response vector $\mathbf{Y}$ is proportional to

$$[\mathbf{Y}|\ \boldsymbol{\alpha}',\ \mathbf{b}',\ \boldsymbol{\sigma}'] \propto \prod_{i=1}^{n} \frac{\exp(Y_i \frac{\mathbf{x_i}^t \boldsymbol{\alpha}' + b'_{j(i)}}{\sigma'_{m(i)}})}{1 + \exp(\frac{\mathbf{x_i}^t \boldsymbol{\alpha}' + b'_{j(i)}}{\sigma'_{m(i)}})},$$

where $\boldsymbol{\sigma}' := (1, \sigma'_2, \ldots, \sigma'_M)^t$, $\sigma'_m := \sqrt{\sigma_m^{2'}}$, $m = 2, \ldots, M$. We assume independent prior distributions for the fixed effect $\boldsymbol{\alpha}'$, the spatial parameters $\mathbf{b}'$ given their dependence parameter ϕ' and the variance scalar $\tau^{2'}$ and the cluster parameters $\boldsymbol{\sigma}'$. Finally we assume independence between the hyperparameters $\phi' := \phi$ and $\tau^{2'} := \frac{\tau^2}{\sqrt{1+\sigma_1'^2}}$. Therefore the joint prior distribution is given by $[\boldsymbol{\alpha}',\ \mathbf{b}',\ \boldsymbol{\sigma}',\ \phi', \tau^{2'}] = [\boldsymbol{\alpha}'] \times [\mathbf{b}'|\phi', \tau^{2'}] \times [\phi'] \times [\tau^{2'}] \times [\boldsymbol{\sigma}']$.

According to (4.5) large deviations from 1 for some σ'_m, $m = 2, \ldots, M$, correspond to large values for some σ_m^2, $m = 1, \ldots, M$, in the primary Model (4.1), which corresponds to insignificance of the regression and spatial effects in these clusters. This implies that one should use a prior for σ'_m which is relatively concentrated around 1. One such choice is a normal distribution $N(1, 4)$ truncated on the interval $[0.2, +\infty)$, which we have chosen as prior for σ'_m, $m = 2, \ldots, M$. Even though σ'_1 is fixed to 1, a value of $\sigma'_m \geq 1 (\leq 1)$ corresponds to $\sigma_m^2 \geq \sigma_1^2 (\sigma_m^2 \leq \sigma_1^2)$. This shows that our prior choice can support high and low variability of cluster m compared to cluster 1.

The parameters $\boldsymbol{\alpha}', \mathbf{b}'$ and $\boldsymbol{\sigma}'$ will be updated using individual MH steps. Since the full conditionals of $\tau^{2'}$ and ϕ' depend only on the spatial effects $\mathbf{b}'$, their MCMC updates have the same form as described in Table 3.2.

5 Simulation Studies

We conducted two simulation studies for spatial logistic regression one with group cluster and the other with individual cluster effects. The first study is based on the Logit Model (3.2) with the following mean structure:

$$\theta_i := \log\left(\frac{p_i}{1 - p_i}\right) = x_{1i}\alpha_1 + x_{2i}\alpha_2 + b_{j(i)} + c_{m(i)}$$

for $i = 1, \ldots, n, j = 1, \ldots, J, m = 1, \ldots, M$. Adapted to our mobility study we simulated $n = 2100$ binary responses residing in $J = 70$ regions arranged on a 7×10 regular lattice and in $M = 5$ clusters so, that each cluster is represented in each region with 6 responses. More precisely, we chose x_{i1} as categorical covariate with possible values 0 or 1 and x_{i2} as continuous covariate taking cycled integer values between 1 and 23 with $\alpha_1 = -1$ and $\alpha_2 = 0.05$. With this choice we achieved good data mixing inside regions and clusters. Spatial effects $\mathbf{b}$ are simulated from the modified Pettitt's Model (2.3) with $\phi = 2$ giving significant spatial smoothing. We chose $\tau^2 = 0.64$ which gives a similar range of the observed spatial effects in the mobility data. A first order neighborhood dependence defined by joint borders was selected. We simulated group cluster effects from $\mathbf{c} \sim N_5(0, \sigma_c^2)$ with $\sigma_c^2 = 1$. As priors we chose $\alpha_1 \sim N(0, 100^2)$, $\alpha_2 \sim N(0, 10^2)$ and $\tau^2 \propto 1$ reflecting a diffuse prior choice. For $\psi = \frac{\phi}{1+|\phi|}$, $\psi \in (-1, 1)$ the $J = 70$ regions may be too few to provide enough information for estimating ψ. Therefore we chose an informative prior density $[\psi] \sim \frac{1}{(1-|\psi|)^{1-a}}$ with $a = 1.25$. From the same reasons we took an informative prior density for σ_c^2, namely $IG(5, 1/6)$ with mean 1.5, variance 3/4 and mode 1.

The MCMC algorithm of Section 3.2 was implemented in MATLAB and was run for 20,000 iterations with every 10th iteration recorded. As "burn in" phase served 5 pilot runs with 300 iterations per each pilot run giving an acceptance rate of $30 - 60$ % for the MH step. The resulting trace plots (not shown) show that such a length of "burn in" phase is enough. The autocorrelation plots (not shown) indicate, that the autocorrelations between recorded iterations are below 0.1. Figure 5.1 shows marginal posterior density estimates of the parameters $\alpha_0, \alpha_1, \psi = \frac{\phi}{1+|\phi|}, \tau^2$ and σ_c^2 from four independent simulated data sets, where the vertical fat dashed lines correspond to the true parameter value. For each density curve its mode is also marked by a thin vertical line. We see that in all four cases the true values are well inside 90% credible intervals. Although the estimation of ψ is somewhat dispersed, posterior mode estimates of the spatial and cluster effects (not shown) are quite precise. This fact indicates the dominance of the observed information over the prior information provided by CAR prior choice.

The second simulation is based on the Logit Model (4.7) with mean structure:

$$\theta_i' := \log\left(\frac{p_i}{1-p_i}\right) = \frac{x_{1i}\alpha_1' + x_{2i}\alpha_2' + b_{j(i)}'}{\sigma_{m(i)}'}, i = 1, \dots, n, \ j = 1, \dots, J, \ m = 1, \dots, M.$$

We used the same spatial and fixed effect structure as in the first simulation study, in particular we set $\alpha_1' = -1$, $\alpha_2' = 0.05$, $\tau^{2'} = 0.64$, $\phi' = 2$. As true values for the cluster parameters σ_m', $m = 2, \dots, M$, we take values $0.5, 1.25, 1.5$ and 2.5, respectively. According to Model (4.7) we set $\sigma_1' = 1$. Prior choices for $\boldsymbol{\alpha}', \mathbf{b}'$ and $\tau^{2'}$ remain the same. For the prior distribution of $\psi' = \frac{\phi'}{1+|\phi|'}$ we used now $Uni(-1, 1)$, since a similar prior choice as for the group cluster case causes slight underestimation of ψ'. For the cluster parameters σ_m', $m = 1, \dots, M$, we used $N(1, 4)$ distribution truncated to the interval $[0.2, +\infty]$ as prior distribution. Figure 5.2 gives posterior density estimates of the cluster variance parameters $\boldsymbol{\sigma}'$ using the MCMC algorithm of Section 4.2 based on 20000 iterations with every 10th iteration recorded indicating

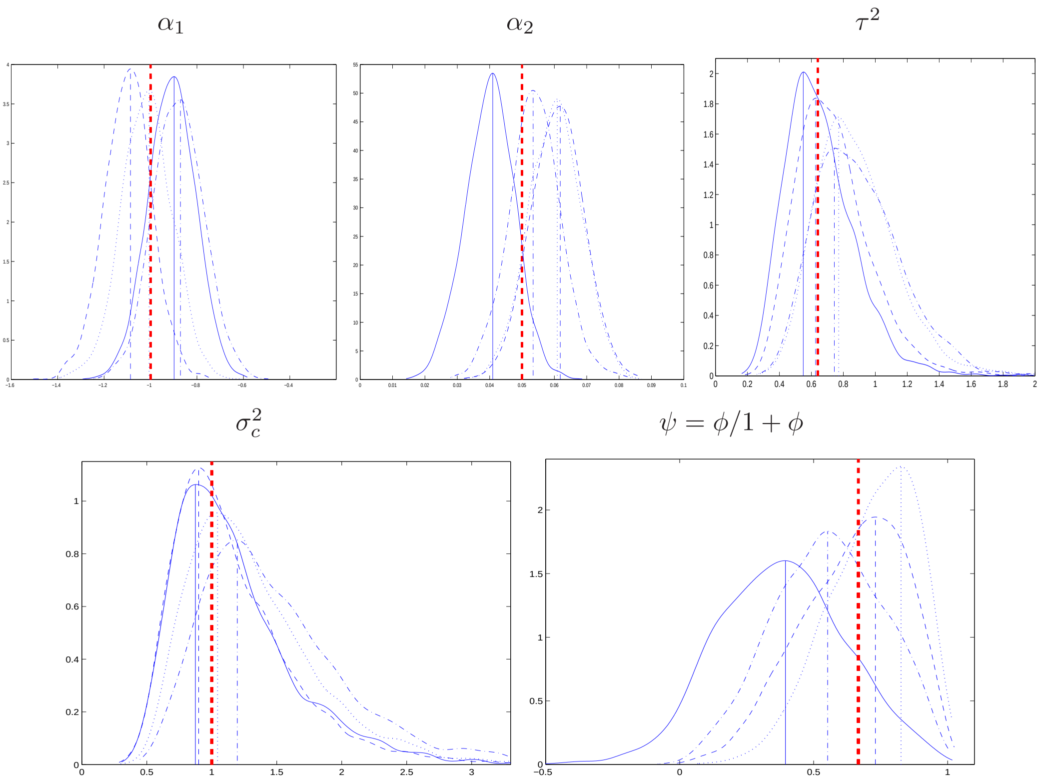

Figure 5.1: Estimated Marginal Posterior Densities for Parameters $\alpha_1, \alpha_2, \tau^2, \psi, \sigma_c^2$ in Model (3.2) (solid for Data Set 1, dashed for Data Set 2, dash-dot for Data Set 3, dotted for Data Set 4)

a satisfactory behavior. The remaining parameters show a similar behavior as the corresponding parameters in Logit Model (4.7) (not shown).

Summarizing we see that all posterior estimates for the main parameters ($\boldsymbol{\alpha}$, $\mathbf{b}$, $\mathbf{c}$ for Model (3.2) and $\boldsymbol{\alpha}$, $\mathbf{b}$, $\boldsymbol{\sigma}'$ for Model (4.7)) lie quite closely around the corresponding true values. With regard to the spatial hyperparameters τ^2 ($\tau^{2'}$) and ψ (ψ') for the group cluster Model (3.2) (individual cluster Model (4.7)) we draw the following conclusions. The number of regions $J = 70$ seems to be enough for τ^2 ($\tau^{2'}$). This is not the case for ψ (ψ'). Further, a simulation study with a large number of regions showed good precision for estimation of ψ, as well as robustness of the posterior with respect to prior choice already for $J = 500$. Finally we note, that in contrast to individual cluster Model (4.7), in group cluster Model (3.2) the small number of clusters M causes lack of information for estimating the cluster variance σ_c^2. If we want to avoid an informative prior choice, the number of clusters must be greater equal 7 to assure a finite variance of the posterior.

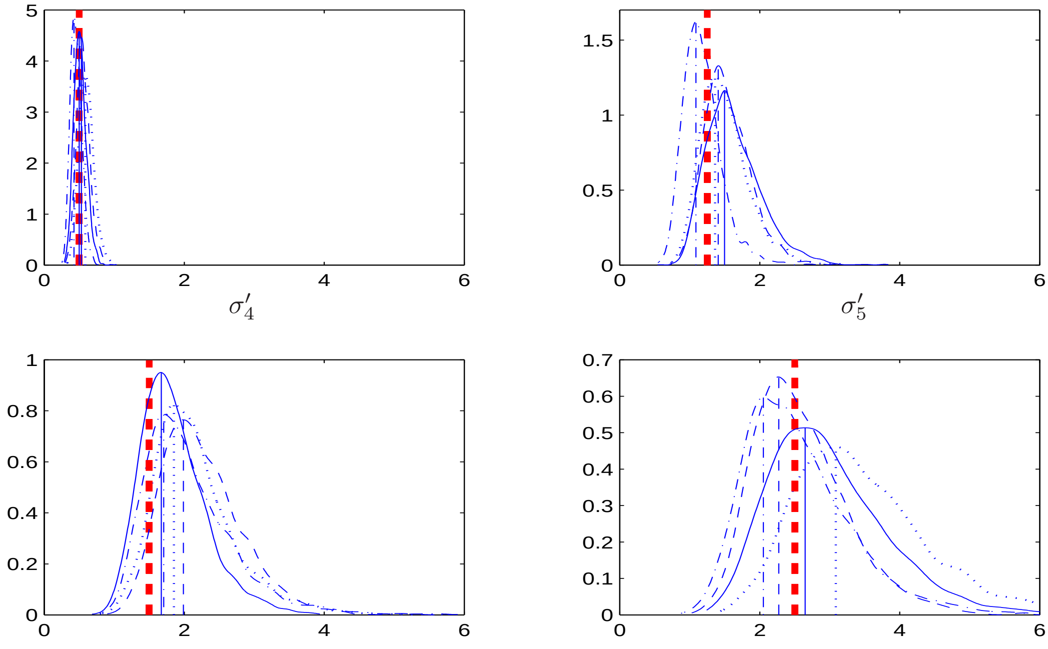

Figure 5.2: Estimated Marginal Posterior Densities for $\boldsymbol{\sigma}'$ in Model (4.7) (solid for Data Set 1, dashed for Data Set 2, dash-dot for Data Set 3, dotted for Data Set 4)

6 Application: Mobility Data

6.1 Data Description

We analyze a data set studying mobility behavior of private households in Munich. One central question is to identify areas of low/high utilization of public transport after adjusting for trip, individual and household related attributes. The goal is to find flexible statistical models which incorporate covariates together with spatial and cluster information. The data was collected within the study "Mobility 97" (Zängler 2000). The participants are German-speaking persons not younger then 10 years, which live in a private household in the state of Bavaria. In order to take into consideration seasonal fluctuations in mobility behavior the survey was carried out in three waves in March, June and October of 1997 with different participants for each wave. Each participant reported all his or her trips conducted by public or individual transport during a period of two or three days. We consider part of the data which includes 1375 trips taken by 296 persons in 167 households living in 74 postal code areas of Munich. For each trip Y has value 1, if individual transport was used and value 0, if public transport was used. Person, household and trip related covari-

Covariable type	Variable	Levels	Number of Trips Using Individual Transport	Public Transport	Total	Most frequently obs. value
PERSON related	**PERSONAL INCOME**	*NO INCOME* ($<$ *200 DM*)	24	31	55	0
		MIDDLE ($200-3000$ *DM*)	475	193	668	1
		HIGH ($>$ *3000 DM*)	521	131	652	0
	USAGE	*MAIN USER*	731	100	831	1
		SECONDARY USER	213	99	312	0
		NOT USER	76	156	232	0
	NET CARD	*YES*	235	247	482	0
		NO	785	108	893	1
	SEX	*MALE*	549	172	721	1
		FEMALE	471	183	654	0
	AGE	metric (quadratic, normalized with Splus function *poly(age,2)*)				**median** 42 years
HOUSE-HOLD related	**HOUSEHOLD TYPE**	*SINGLE*	156	125	281	0
		SINGLE PARENT	84	10	94	0
		NOT SINGLE	780	220	1000	1
TRIP related	**DAY TYPE**	*WORK DAY*	595	297	892	1
		WEEKEND	425	58	483	0
	DISTANCE	*SHORT* ($\leq$ *3.5 km*)	294	71	365	0
		MIDDLE ($3.6-21.5$ *km*)	571	257	828	1
		FAR ($>$ *21.5 km*)	155	27	182	0
	WAY ALONE	*ALONE*	507	267	774	1
		NOT ALONE	513	88	601	0
	DAY TIME	*DAY (6 a.m. - 9 p.m.)*	905	336	1241	1
		NIGHT (9 p.m. - 6 a.m.)	115	19	134	0
T O T A L			1020	355	1375	

Table 6.1: Significant covariates identified in logistic regression model selection without spatial and cluster effects

ates were recorded. Neglecting spatial and cluster effects standard model selection techniques for logistic regression selected the following covariates. Person related covariates are age (metric), sex, personal income, car usage (main, secondary or not user) and whether the person possesses or not a public transport net card. We retain only one household related covariate, namely household type (single, single parent or not single). Trip related covariates are day type (work day or weekend), day time (day or night), distance and whether the person took the trip alone or not alone. Table 6.1 shows the chosen covariates. For the covariate **USAGE**, note that both main and secondary users must be not younger than 18 years and must have a driver license and a car available in the household. The following significant interactions were identified: WAY ALONE:NET CARD, USAGE:SEX, WAY ALONE:USAGE, DIS-

TANCE:USAGE, DAY TYPE:NET CARD, USAGE:DAY TIME, SEX:DAY TIME, PERSONAL INCOME:NET CARD, DISTANCE:AGE and DAY TYPE:AGE. We used this model as a starting model for our analysis. We like to note that a seasonal effect measured by temperature is not significant. Trips which have been taken together by let us say k persons are treated as k trips each associated with the specific covariates of these persons. The fact that these trips were conducted together is taken into account by the covariate WAY ALONE defined in Table 6.1.

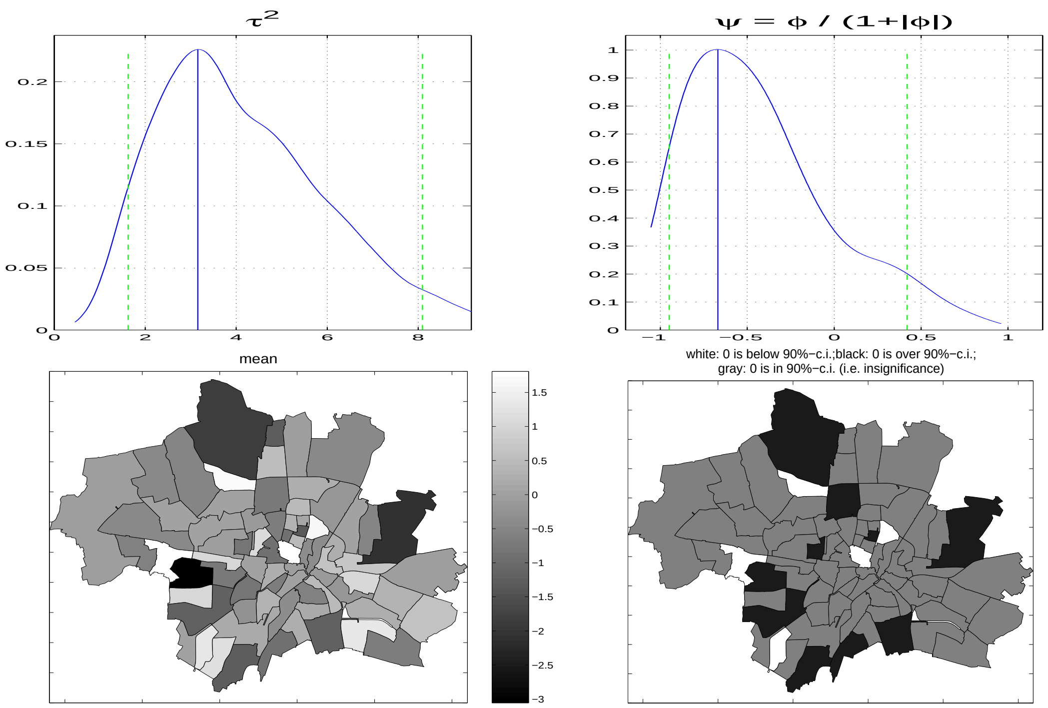

Figure 6.1: Results for Model 1: Top: Estimated Posterior Densities of Spatial Hyperparameters (Solid Line = Estimated Posterior Mode, Dashed Line = 90% CI). Bottom: Estimated Posterior Mean Spatial Effects $\hat{b}_j$, $j = 1, \ldots, 74$ and 90% CI

6.2 Results

We present the results for 8 different model specifications. Model 1 is a spatial logit regression model with no cluster effects, while Models 2 — 5 are spatial logit models with group cluster effects. Finally Models 6 — 8 are logit spatial models with individual cluster effects. For all models 25000 MCMC iterations were run and every 25^{th} iteration was recorded, giving acceptable low autocorrelations (not shown). We found, that 10 pilot runs with 300 iterations per pilot run are sufficient as “burn in”.

As a starting point for the choice of fixed effects we used the covariates identified in Table 6.1 involving a total of 36 regression parameters. The intercept effect is modeled within the spatial and cluster part. As prior for $\alpha_1, \ldots, \alpha_{36}$ we chose independent normal distributions with zero mean and standard deviation equal to 5. We consider an interaction as insignificant when the corresponding estimated 90% credible interval (CI) contains the zero value for all interaction terms. If an interaction is found to be insignificant, then the corresponding terms were removed and the model reestimated. Continuing with this procedure we arrive at a model where all interactions are significant.

For Model 1 we chose a uniform prior for $\psi = \frac{\phi}{1+|\phi|}$ on $(-1, 1)$ and for $[\tau^2] \propto 1$. The top row of Figure 6.1 presents estimated posterior densities for τ^2 and ψ. The parameter ψ is negative, which indicates that positive spatial effects in an area can be surrounded by negative spatial effects and vice versa. This is seen in Figure 6.1 (bottom row), where posterior means and their 90% CI of the spatial effects are given.

In Models 2 — 5 we used group cluster specifications. First we considered in Model 2 group clusters formed by the 74 postal codes. Since single trips are taken by individuals and households we ideally would like to allow for person or household specific effects. This would require the estimation of 296 person and 167 household specific cluster variances. The data is too sparse to accommodate these models. Therefore we form cluster groups by the number of trips taken by individuals or households. Since it is unclear how many cluster groups should be considered, we investigated several specifications. To avoid unbalanced cluster groups we chose the cut points in such a way that the corresponding cluster groups consists of about equal number of trips. For example for Model 3 we used 5 clusters with 1st cluster group defined with ≥ 23 trips, the 2nd one with $16-22$ trips, the 3rd one with $12-15$ trips, the 4th one with $8-11$ trips and the 5th one with ≤ 7 trips per household. Model 4 and 5 use 12 and 5 cluster groups formed by the number of trips a household has taken, respectively.

For σ_c^2 we choose $\sigma_c^2 \sim IG(3, 0.5)$, while prior choices for fixed and spatial parameters remain the same as in Model 1. Only in Model 2, in order to avoid numerical problems (clustering around border values -1 and 1) we chose $[\psi] \propto (1-|\psi|)^{0.5}$ instead of $[\psi] \propto 1$ on the interval $(-1, 1)$. The posterior centrality estimates of the hyperparameters and their 90% CI are given in Table 6.2. In Model 2 we have as cluster groups the 74 postal codes. Therefore both structured $(b_j,\ j = 1, \ldots, 74)$ and unstructured $(c_j,\ j = 1, \ldots, 74)$ spatial effects are included in Model 2. Figure 6.2 presents spatial maps with estimated posterior means for the structured spatial effects b_j (top left) and unstructured spatial effects c_j (top middle). On the top right map we present estimated posterior means of the sum $b_j + c_j$ of structured and unstructured spatial effects. Corresponding 90% CI are given in the middle row of Figure 6.2. Both structured and unstructured effects are insignificant, while their sum is, and form a similar spatial pattern as in Model 1. Therefore it is not surprising that the posterior density of ψ, are also similar (see bottom row of Figure 6.2).

Model	Number of Clusters	Parameter	Mode	Mean	Median	90%	CI
						5%	95%
	74 formed	ψ	−0.500	−0.271	−0.372	−0.857	0.646
2	by postal	τ^2	3.628	4.777	4.313	0.981	10.335
	codes	σ_c^2	0.554	0.836	0.678	0.315	1.912
	5 formed	ψ	−0.541	−0.422	−0.446	−0.930	0.149
3	by # of trips	τ^2	6.262	9.124	8.233	3.358	18.417
	per household	σ_c^2	0.802	1.270	1.076	0.486	2.797
	12 formed	ψ	−0.507	−0.516	−0.538	−0.954	0.031
4	by # of trips	τ^2	6.293	8.299	7.452	3.194	16.067
	per household	σ_c^2	0.880	1.272	1.122	0.589	2.398
	5 formed	ψ	−0.874	−0.543	−0.594	−0.956	0.058
5	by # of trips	τ^2	4.025	5.298	4.777	2.020	9.685
	per person	σ_c^2	0.526	0.753	0.646	0.324	1.585
6	3 formed	ψ	−0.468	−0.396	−0.418	−0.870	0.181
	by household	$\tau^{2'}$	4.861	6.854	5.931	2.553	14.196
	type	σ_2'	0.277	0.484	0.430	0.226	0.921
		σ_3'	1.439	1.461	1.443	1.068	1.943
7	5 formed	ψ	−0.410	−0.413	−0.422	−0.865	0.075
	by # of trips	$\tau^{2'}$	10.769	17.101	14.799	6.002	36.512
	per household	σ_2'	0.922	1.010	0.973	0.648	1.464
		σ_3'	2.842	2.951	2.913	2.240	3.734
		σ_4'	1.430	1.486	1.459	1.078	2.019
		σ_5'	1.822	1.797	1.789	1.313	2.343
8	5 formed	ψ	−0.476	−0.403	−0.439	−0.876	0.199
	by # of trips	$\tau^{2'}$	7.538	9.468	8.232	3.167	19.895
	per person	σ_2'	1.027	1.058	1.041	0.752	1.430
		σ_3'	1.168	1.180	1.166	0.797	1.610
		σ_4'	1.271	1.300	1.287	0.897	1.768
		σ_5'	1.553	1.681	1.642	1.196	2.255

Table 6.2: Point and Interval Posterior Estimates for the Hyperparameters in Models 2 - 5 (with Group Cluster Effects) and Models 6 - 8 (with Individual Cluster Effects)

In Figure 6.3 we present for Model 3 estimated posterior densities of the group cluster effects c_m, $m = 1, \ldots, 5$. A cluster effect is significant (marked with *), if its 90% CI does not include zero. Note that cluster effects for households with large numbers of trips are positive and cluster effects for households with few numbers of trips are negative. Finally the maps on the bottom row of Figure 6.4 give estimated spatial effects.

Also in Models 4–5 only the higher cluster effects (i.e. with fewest numbers of trips) are significant. For brevity we omit the corresponding density plots. For Models 4–5 the spatial patterns are similar to the ones of Models 1 or 3 and Model 2 when the joint effect of structured and unstructured spatial components is considered. The posterior density of ψ also remains similar (not shown for Models 3, 4 and 5). Table 6.2 gives posterior centrality estimates and 90% CI's for the hyperparameters.

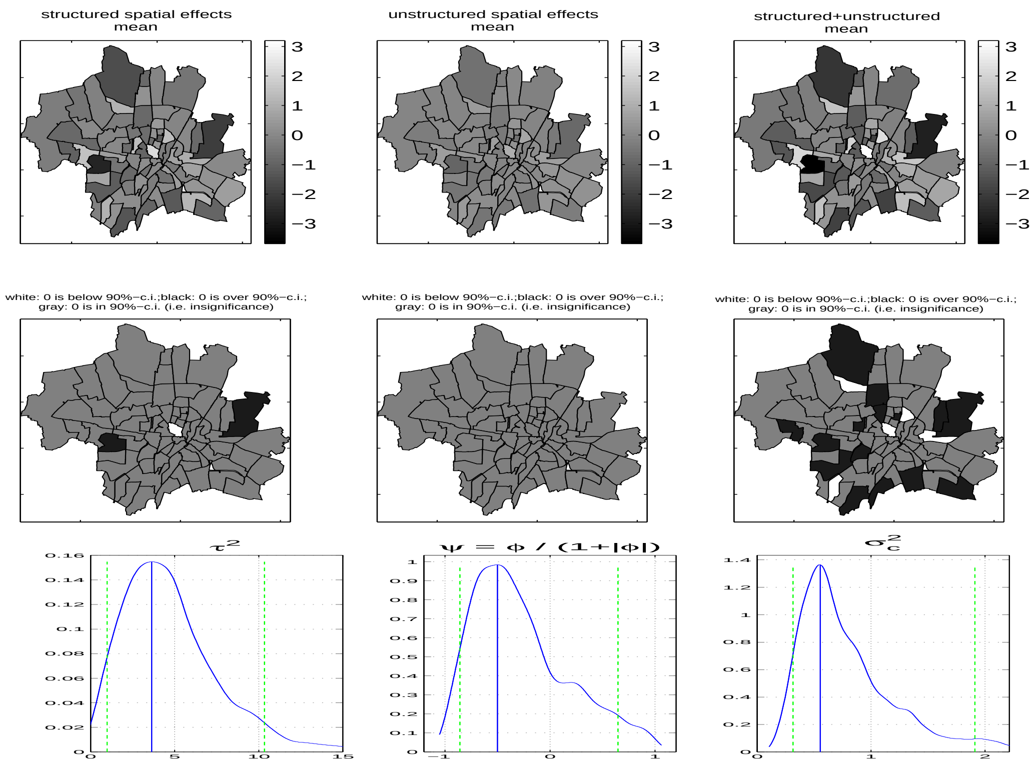

Figure 6.2: Results for Model 2: Top: Estimated Spatial Effects: Structured $\hat{b}_j, j = 1:74$ (left), Unstructured $\hat{c}_j, j = 1:74$ (middle) and their Sum $\widehat{b_j + c_j}, j = 1:74$ (right). Middle: 90% Credible Intervals for Structured Effects, Unstructured Effects and their Sum. Bottom: Estimated Posterior Densities of Hyperparameters (Solid Line = Estimated Posterior Mode, Dashed Line = 90% CI)

We consider now model specifications with individual cluster effects given in Table 6.2. As before, we chose a flat prior $[\tau^{2'}] \propto 1$ and take $[\psi'] \propto (1 - |\psi'|)^{0.5}$ to avoid numerical problems (clustering around border values -1 and 1). In Models 6–8 we assume for $\sigma'_2, \ldots, \sigma'_M$ a normal $N(1,1)$ prior truncated to $(0.2, +\infty)$.

The posterior centrality estimates and their 90% CI's of the hyperparameters for Models 6–8 are given in Table 6.2. We see that cluster components of the higher clusters are significant, i.e. $1 \notin 90\%$ CI. This shows that the heterogeneity within the group with the fewest numbers of trips per household (or per person) is the largest. Further we see, that more cluster components are significant for individual cluster effects formed by household type or number of trips per household than by the number of trips per person. In all models with individual cluster effects the spatial dependence hyperparameter ψ is negative and about the same size.

The estimates for the fixed effects $\boldsymbol{\alpha}'$ in all 8 models are given in Table 6.3. Posterior mode estimates are marked with *, when the corresponding parameter is

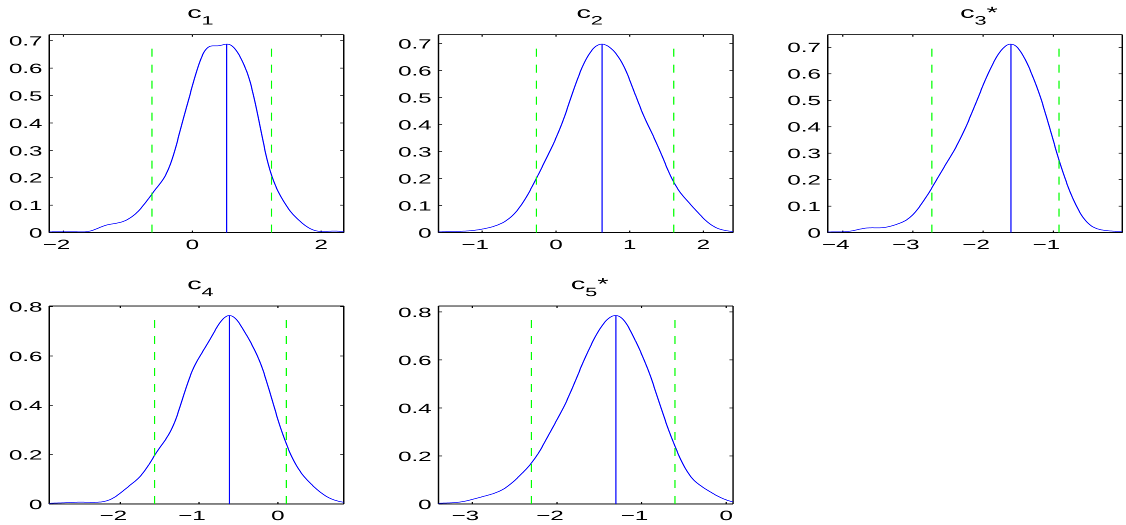

Figure 6.3: Estimated Posterior Densities of Group Cluster Effects c_m, $m = 1, \ldots, 5$ in Model 3. (Solid Line = Estimated Posterior Mode, Dashed Line = 90% CI)

insignificant, i.e. the 90% CI contains zero. If all terms of an interaction effect were insignificant, the model was reduced and reestimated. Those interactions are marked with "*n.r.*", correspond to "not represented" in the model. In particular the significant interactions PERSONAL INCOME:NET CARD and DAY TIME:AGE from the starting logistic model disappear.

Main Effect	Model 1	2	3	4	5	6	7	8
	spatial only	spatial+group cluster				spatial+individual cluster		
PERSONAL INCOME								
MIDDLE	0.41*	0.48*	1.63	1.41	0.71*	1.06	1.62	0.64*
HIGH	0.25*	0.42*	1.27	1.14	0.12*	0.76	1.46	0.24*
USAGE								
SECOND.USER	0.38*	1.09*	1.27	1.16*	0.88*	1.11	1.23*	1.51
NOT.USER	−3.87	−6.41	−6.52	−6.52	−5.90	−6.38	−9.99	−7.44
NET CARD								
NO	2.07	2.67	3.03	3.32	2.72	2.78	3.90	3.11
SEX								
FEMALE	0.28*	0.16*	−0.19*	−0.47*	0.10*	0.30*	−0.48*	0.01*
AGE								
POLY.AGE.1	16.80	8.73	11.64	11.53	9.95	6.11	7.97	9.81
POLY.AGE.2	−13.07	−8.96	−9.03	−8.64	−9.67	−8.93	−9.69	−7.63
HOUSEHOLD								
SINGLE.PARENT	1.61	3.15	3.42	2.92	3.31	n. r.	4.24	3.65
NOT.SINGLE	0.70	0.68	0.25*	0.27*	0.90	n. r.	0.85*	0.96
DAY TYPE								
WEEKEND	1.44	2.21	2.46	2.52	2.11	2.25	3.32	2.78
DISTANCE								
MIDDLE	−0.96	−1.15	−1.06	−1.17	−1.05	−1.29	−1.90	−1.16
FAR	0.32*	0.81*	0.98*	0.83*	0.97*	1.21*	0.78*	0.85*
WAY ALONE								
NOT.ALONE	1.82	2.09	2.07	2.30	1.93	2.17	3.21	2.30
DAY TIME								
NIGHT	−0.58*	−1.02	−1.12	−1.29	−1.13	−1.19	−1.99	−1.30
Interaction								
WAY ALONE:NET CARD								
NOT.ALONE:NO	−1.86	−2.39	−2.37	−2.76	−2.37	−1.54	−1.69	−2.53
USAGE:SEX								
SECOND.USER:FEMALE	−1.70	−2.13	−2.07	−1.81	−2.01	−2.30	−2.80	−2.50
NOT.USER:FEMALE	−0.20*	0.66*	0.58*	0.79*	0.40*	0.26*	1.39*	0.80*
WAY ALONE:USAGE								
NOT.ALONE:SECOND.USER	0.79	1.21	0.80	0.76*	1.22	1.09	1.20*	1.32
NOT.ALONE:NOT.USER	1.75	3.65	4.19	3.76	3.41	4.35	5.08	4.22
DISTANCE:USAGE								
MIDDLE:SECOND.USER	−0.68*	−1.03	−1.39	−0.97	−1.19	−1.31	−1.44*	−1.54
FAR:SECOND.USER	−1.02	−2.25	−2.12	−1.72	−2.22	−2.73	−2.41	−3.61
MIDDLE:NOT.USER	0.95	1.68	1.52	1.64	1.27	1.20	2.47	1.53
FAR:NOT.USER	−1.19	−1.19*	−1.55*	−2.01	−1.51*	−2.31	−2.68	−1.94
DAY TYPE:NET CARD								
WEEKEND:NO	n. r.	−0.91	−1.23	−1.23	−1.07	−0.82*	−1.51	−1.25
USAGE:DAY TIME								
SECOND.USER:NIGHT	1.32	5.01	5.22	6.63	5.71	5.07	6.17	5.67
NOT.USER:NIGHT	−0.06*	0.31*	0.45*	0.38*	0.26*	0.32*	0.72*	0.68*
SEX:DAY TIME								
FEMALE:NIGHT	1.70	2.88	3.36	3.55	3.49	3.02	2.94	3.30
DISTANCE:AGE								
MIDDLE:POLY.AGE.1	−12.93	n. r.	n. r.	n. r.	n. r.	n. r.	n. r.	n. r.
FAR:POLY.AGE.1	−0.09*	n. r.	n. r.	n. r.	n. r.	n. r.	n. r.	n. r.
MIDDLE:POLY.AGE.2	−2.41*	n. r.	n. r.	n. r.	n. r.	n. r.	n. r.	n. r.
FAR:POLY.AGE.2	0.76*	n. r.	n. r.	n. r.	n. r.	n. r.	n. r.	n. r.

Table 6.3: Posterior Mode Estimates for Main Effect and Interaction Parameters (*= 90% credible interval does not include 0, n.r.= effect was not required in model, since model with effect has a 90% credible interval which includes 0)

6.3 Model Comparison

A general method for model comparison in Bayesian models estimated by MCMC is the DIC criterion suggested by Spiegelhalter et al. (2002). It is developed for exponential family models and based on the deviance. Even though binary logit models belong to this class, Collett (2002) has shown that the residual deviance in binary regression should not be used for model assessment, while the partial deviance is valid for nested model comparison. Further, Figure 1 of Spiegelhalter et al. (2002) shows that the DIC does not perform satisfactory for binary responses. Since our binary responses cannot be grouped to binomial responses with sufficient large numbers of trials because of the complexity of the fixed, spatial and cluster effects, we decided not to use the DIC criterion. Meaningful DIC values of our models can be determined as long as the binary regression data can be grouped to binomial regression data with sufficiently large number of trials.

To facilitate model comparison we follow two alternative approaches. In the first one we focus on the spatial fit, while in the second one we focus on the overall fit. For the first focus we propose to use D_w the sum of weighted squared residuals over all postal codes of Munich defined by

$$D_w(\mathbf{Y}) := \sum_{j=1}^{74} n_j (p_j^{empir} - p_j^{estim})^2 , \tag{6.1}$$

where n_j := number of trips in the j^{th} postal code. Empirical probabilities p_j^{empir} are equal to the observed proportion of trips using individual transport in postal code area j, and posterior probability estimates p_j^{estim} are based on the MCMC run, and defined as:

$$p_j^{\text{estim}} := \frac{1}{n_j * R} \sum_{i:\, j(i)=j} \sum_{r=1}^{R} \frac{\exp(\eta_{ir})}{1+\exp(\eta_{ir})}, \tag{6.2}$$

where

$$\eta_{ir} := \begin{cases} \mathbf{x_i^t}\boldsymbol{\alpha}_r + b_{j(i),r} & \text{for Model 1} \\ \mathbf{x_i^t}\boldsymbol{\alpha}_r + b_{j(i),r} + c_{m(i),r} & \text{for Models 2-5} \\ \frac{\mathbf{x_i^t}\boldsymbol{\alpha}'_r + b'_{j(i),r}}{\sigma'_{m(i),r}} & \text{for Models 6-8 .} \end{cases}$$

Here $\boldsymbol{\alpha}_r, b_{j,r}$ and $\sigma'_{m,r}$ are the corresponding MCMC estimates in the rth recorded iteration.

In Table 6.4 we present value D_w for all 8 models and the number of parameters required in calculating D_w. The total number of parameters required for D_w will be used as a rough measure for the complexity of the model with regard to the spatial fit. This means we regard these parameters as model parameters and the

Model	1 spatial only	2	3 spatial + group cluster	4	5	6	7 spatial + individual cluster	8
fixed effects	31	28	28	28	28	26	28	28
spatial effects	74	74	74	74	74	74	74	74
cluster effects	0	74	5	12	5	2	4	4
total number of parameters for D_w	105	176	107	114	107	102	106	106
$\mathbf{D_w}$	2.35	1.23	0.95	1.02	1.44	1.9	3.25	1.84
$\sum_{i=1}^{n}(\mu_i - y_i)^2$	110.49	106.54	96.05	94.91	102.86	108.30	103.77	107.56
$\sum_{i=1}^{n}\sigma_i^2$	129.75	111.96	104.19	102.64	109.84	114.21	111.50	114.55
PMCC	240.24	218.50	200.23	197.55	212.70	222.51	215.27	222.11
BS	.0866	.0840	.0768	.0761	.0812	.0851	.0821	.0850

Table 6.4: Model Fit Comparison using D_w, PMCC and BS

spatial dependence parameter, spatial variance and the cluster variance parameters in group cluster models as hyperparameters belonging to the prior. This approach is consistent with the approach taken in Spiegelhalter et al. (2002), which point out in their discussion that complexity depends on the focus of the analysis. We want to add that in setting our focus on assessing the spatial fit, the corresponding calculations of the complexity measure p_D suggested by Spiegelhalter et al. (2002) cannot be facilitated since the corresponding deviances are not available in closed form as pointed out by S.P. Brooks in the discussion of Spiegelhalter et al. (2002). According to Table 6.4 the best fit with regard to spatial probabilities has Model 3 (with group cluster effects). We see that even though the models with individual cluster effects have a lower model complexity with regard to spatial fit, their goodness of fit as measured by D_w is worse than Model 3. Model 4 has a comparable D_w value to Model 3 but the model complexity is higher, therefore we prefer Model 3.

To complement our analysis of spatial fit we consider now also the predictive model choice criterion (PMCC) of Gelfand and Ghosh (1998) and the Brier score BS (Brier 1950) as proper scoring rule (Gneiting and Raftery 2004). The PMCC is defined as

$$PMCC = \sum_{i=1}^{n}(\mu_i - y_i)^2 + \sum_{i=1}^{n}\sigma_i^2$$

where $\mu_i := \frac{1}{R}\sum_{r=1}^{R} p_{ir}$ and $\sigma_i^2 := \frac{1}{R}\sum_{r=1}^{R} p_{ir}(1-p_{ir})$ are MCMC based estimates of the mean and variance of the posterior predictive distribution. Here $p_{ir} = \frac{\exp(\eta_{ir})}{1+\exp(\eta_{ir})}$. The second term is considered as a penalty term which will tend to be large both for poor and overfitted models. The Brier score BS for our models is given by

$$BS = \frac{1}{nR}\sum_{r=1}^{R}\sum_{i=1}^{n}(p_{ir} - y_i)^2.$$

PMCC and BS are given in Table 6.4 and again show that Models 3 and 4 are the preferred models. This substantiates that Model 3 is the preferred overall model.

For Model 3 we present a map with estimated spatial probabilities over postal codes of Munich (Figure 6.4, top right map), which coincides quite well with the map showing the empirical spatial probabilities (Figure 6.4, top left map). This indicates that Model 3 has a reasonably good fit of the data with respect to the spatial resolution.

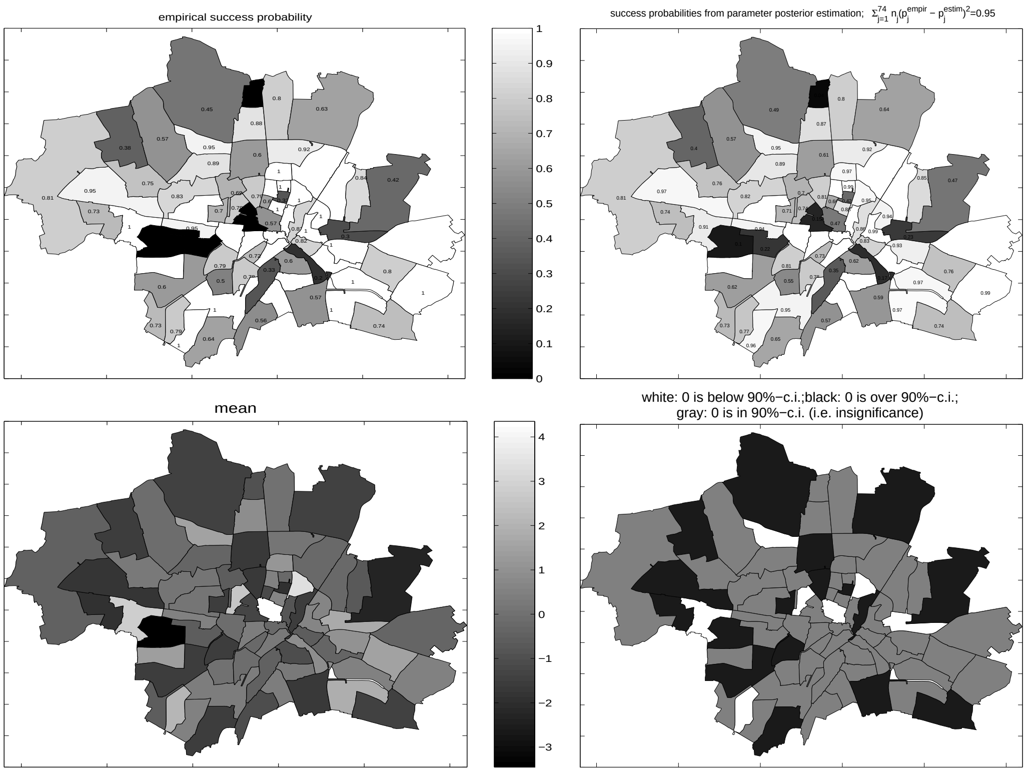

Figure 6.4: Top right map: Observed Probabilities of Individual Transport Use by Postal Codes in Munich, Germany; Top left map: Posterior Mean Probability Estimates of Individual Transport Use by Postal Codes in Munich, Germany for Model 3; Bottom maps: Estimated Spatial Effects $\hat{b}_j$, $j = 1, \ldots, 74$ in Model 3.

6.4 Model Interpretation

After model fitting and model selection one is interested in what can be learned about the travel mode decisions based on Model 3. First we estimate individual transport probabilities when one or combinations of two covariates change. The remaining covariates in the model are set to their "most usual values", corresponding to the modus for categorical covariates and median values for quantitative covariates (Table 6.1). Since Model 3 includes spatial effects we have to specify a postal code for which we estimate these probabilities. We have chosen postal code area 81377, since this postal code area has a large observed number of trips and the smallest 90% CI for its spatial effect. Finally Model 3 contains group cluster effects with regard to the number of trips a household has taken. Since each cluster group contains the similar number of individual trips, for our investigations we chose the last, i.e. the 5th cluster group corresponding to households with ≤ 7 trips, which has the smallest 90% CI for its cluster effect c_5. For "the most usual" trip associated with postal code 81377 and 5th cluster, the estimated posterior mean probability for taking individual transport is equal to 0.7.

Figure 6.5 gives the estimated posterior mean probability with 90% credible

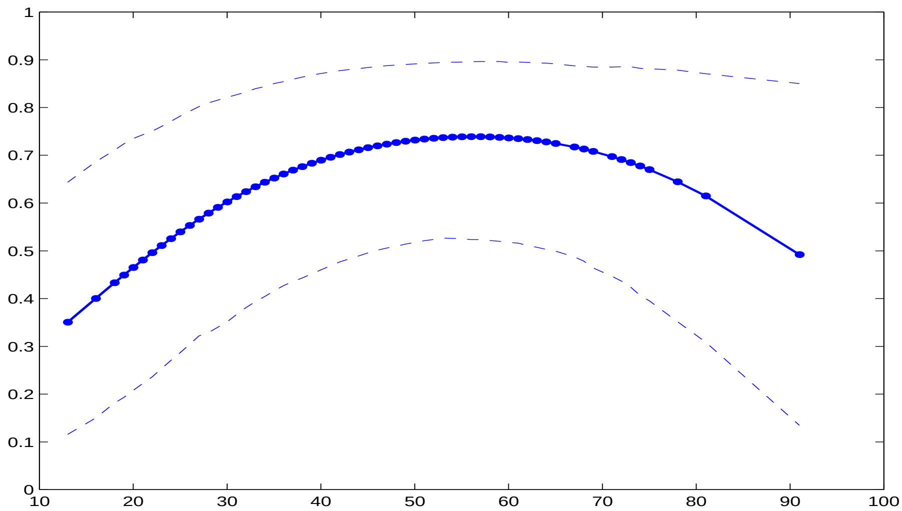

Figure 6.5: Estimated posterior mean probabilities for using an individual transport in Postal code area 81377 and 5th cluster group for different **AGE**, while other covariates are set as in Table 6.1 (dotted lines correspond to 90% credible bounds)

bounds for choosing individual transport as age changes in postal code area 81377 and trips associated with the 5th cluster when the remaining covariates are set to their "most usual value". It is not very surprising that the probability of using a car increases rapidly to an age of about 35 years, remains reasonably stable between 35 years and 65 years and decreases slowly after 65 years. Younger people have a lower probability to own a car, while older people might prefer public transport options.

We can interpret the effect of age directly, since no interaction terms include age. For almost all other covariate effects we have to consider covariate combinations corresponding to interaction terms. Note that Model 3 includes 7 interaction terms. In order to interpret effects of the categorical covariates we plot for each of the 7 interactions the estimated posterior mean probabilities for using individual transport. For brevity we interpret only 2 of the 7 interaction plots. From top left panel of Figure

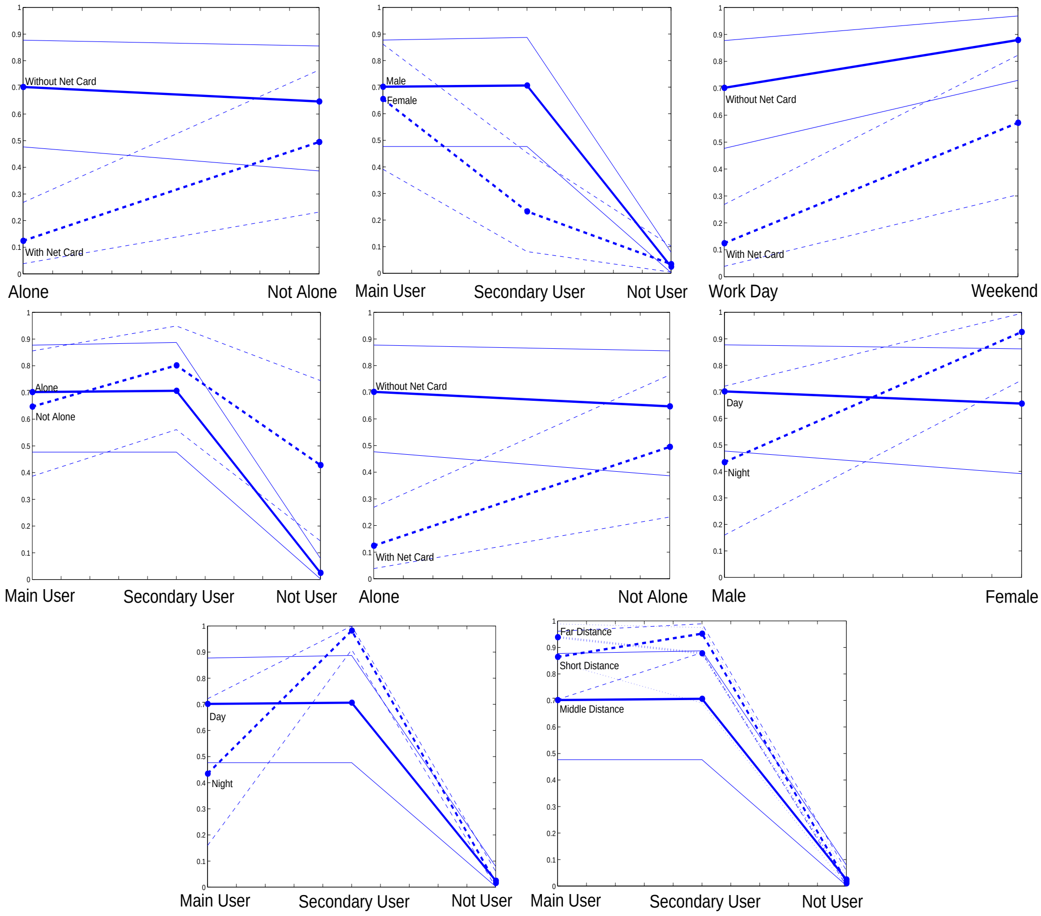

Figure 6.6: Estimated posterior mean probabilities for individual transport in Postal code area 81377 and 5th cluster group for different combinations of the covariates which form the interaction, while other covariates are set to the "most usual value" given in Table 6.1. Dotted lines correspond to 90% credible bounds.

6.6 we see that net card users prefer public transport for trips taken alone much more often than when the trip is taken with others. This is to be expected since a net card in general can only be used by a single person. In contrast users without a net card take individual transport options much more often regardless if the trip taken alone or not. The right panel in the second row shows an interesting behaviorial difference between females and males. During the day there is a little difference. However during night women nearly always use individual transport options, while males choose this option only half as often. An explanation might be that women are afraid to use public transport at night because of low usage and deserted stops, while males might prefer a car free option at night. This shows that some expected behaviorial patterns can be captured when interactions are allowed in the model. The remaining panels of Figure 6.6 are interpreted in detail in Section 6.4 of Prokopenko (2004).

We continue now with the interpretation of spatial effects. There are 24 postal codes whose 90% CI's do not include zero and therefore are significant. We expect that the interpretation of the spatial effects is related to the structure of the subway (U-Bahn) net and suburban railway (S-Bahn) net. Table 6.5 confirms our assumption in general. The left column shows the numbers of postal code areas, which have U- or S-stops inside. The right column contains the numbers of postal code areas without

	with U- or S-stops	without U- or S-stops inside PLZ
90% CI over 0	2 (80333, 81476)	5
90% CI below 0	11	6 (80999, 80634, 80797, 81243, 80689, 81373)

Table 6.5: Interpretation of spatial effects in context of presence/absence of the U-or S- stops inside of postal codes; the postal code numbers of 8 untypical postal code areas are given in parentheses.

stops. The estimated odds ratio of Table 6.5 is $\frac{2 \cdot 6}{11 \cdot 5} \approx 0.22$, which is below 1 (a 90% confidence interval is $[0.044, 1.091]$). This confirms that presence of U- and S-stops are related to significant spatial effects. While there is a general relationship between significant spatial effects and the presence of the U+S-net in these postal areas, 8 areas do not follow this pattern (see Table 6.5). These areas should therefore be of special interest to the city planners, which seek to improve the public transport net, since these areas indicate areas of low/high public transport usage even after adjustment of trip, person, household specific effects and the structure of the public transportation network. We noted that the estimate of the spatial dependency parameter $\hat{\psi} \approx -0.5$ is negative. This can be explained by the specific structure of S- and U-Bahn net of Munich, whose lines run from the center to suburbs like a star. Since the sign of the spatial effects correlates with the presence/absence of the U-or S- stops, it is not surprising, that especially far from the center the neighboring postal codes have often spatial effects with opposite signs.

Finally we mention that cluster effects for households with large numbers of trips are positive and cluster effects for households with few numbers of trips are negative (Figure 6.3). This implies that households with high mobility needs use a car more often than households with low mobility needs.

7 Model Modifications and Extensions

A possible modification of the models developed in Section 3 and 4 is to consider the problem of including interactions between cluster and spatial effects. For this we suggest to use multivariate CAR models mentioned for example by Pettitt et al. (2002), which is a model for $\mathbf{b} = (\mathbf{b_1}, \ldots, \mathbf{b_J})^t$, where the components $\mathbf{b_j} = (b_{j1}, \ldots, b_{jM})^t,\ j = 1, \ldots, J$ are $M-$dimensional vectors instead of scalars, as before. The joint distribution of the vector $\mathbf{b}$ is defined as follows:

$$\mathbf{b} = (\mathbf{b_1}, \ldots, \mathbf{b_J})^t \sim N_{J\times M}\left(\mathbf{0}, \tau^2(Q^{-1} \otimes V)\right), \quad V = \begin{pmatrix} 1 & \rho & \cdots & \rho \\ \rho & 1 & & \vdots \\ \vdots & & \ddots & \rho \\ \rho & \cdots & \rho & 1 \end{pmatrix} \in \mathbb{R}^{M\times M}, \tag{7.1}$$

where $A \otimes B$ stands for *Kronecker product* of matrices A and B. In particular for the multivariate modified Pettitt CAR, the conditional distribution is given then as follows (compare with (2.3)):

$$\mathbf{b_j}|\mathbf{b_{-j}} \sim N_M\left(\frac{\phi}{1+|\phi|N_j}\sum_{j\sim j'}\mathbf{b_{j'}}, \frac{(1+|\phi|)\tau^2}{1+|\phi|N_j}V\right).$$

The parameter ρ measures the strength of the cluster dependence. If $\rho = 0$ then all M components of vector $\mathbf{b_j}$ are iid, i.e. $\mathbf{b}$ decomposes in M iid. Gaussian CAR models. As before, the parameter ϕ measures the strength of the spatial dependence. If $\phi = 0$ then the vectors $\mathbf{b_j}, j = 1, \ldots, J$ are independent and normally distributed with mean zero and covariance matrix $\tau^2 V$. Properties of the multivariate CAR model are studied in Pettitt et al. (2002). Gelfand and Vounatsou (2003) use multivariate extensions of Sun et al. (2000) proper CAR models for multivariate continuous and multinomial spatial data. Carlin and Banerjee (2003) use multivariate Gaussian CAR models in multivariate generalized linear mixed models.

We now propose to apply the multivariate Gaussian CARs in a new way, namely for modeling spatial-cluster interactions for univariate response data. More precisely, we propose to model spatial and cluster effects jointly as some multivariate CAR. As before, J denotes the number of regions, while M stands for the number of clusters. Therefore the multivariate Gaussian CAR model is associated with a(M-categorial) covariate instead of a(M-variate) response as usually. Then logits are modeled as follows (compare with (3.2)):

$$\theta_i := \log\left(\frac{p_i}{1-p_i}\right) = \underbrace{\mathbf{x_i^t}\boldsymbol{\alpha}}_{\textit{fixed effect}} + \underbrace{b_{j(i),m(i)}}_{\textit{spatial and cluster effect}}, \tag{7.2}$$

where $\mathbf{b} = (\mathbf{b_1}, \ldots, \mathbf{b_J})^t,\ \mathbf{b_j} = (b_{j1}, \ldots, b_{jM})^t,\ j = 1, \ldots, J$ is modeled as a realization of the multivariate CAR (7.1). We have to estimate one additional parameter ρ, which measures strength of a space-cluster interaction. The absence of interaction is indicated by $\rho = 0$. In this case the M vectors $(b_{1m}, \ldots, b_{Jm})^t,\ m = 1, \ldots, M$ are independent identically distributed Gaussian CAR models. Model (7.2) has been investigated for simulated data, where it performed well. However the sparseness of the mobility data does not support an application of such models for this data. Prokopenko (2004) also shows that modeled interaction present in the multivariate CAR model (7.1) can be interpreted as a product of spatial and cluster effects.

A further extension would be to model heterogeneity within and between clusters simultaneously, which would combine group and individual cluster approaches. In particular we would assume $c_{mk} \sim N(c_m, \sigma_m^2), \quad c_m \sim N(0, \sigma_c^2) \quad$ for $m = 1, \ldots, M,\ k = 1, \ldots, K_m$. (compare with 3.2 and 4.1). Here a similar non identifiability problem has to be solved and is subject of current research.

8 Summary and Discussion

An extended version of the spatial Gaussian CAR model proposed by Pettitt et al. (2002) has been presented, which allows for spatial independence and the intrinsic CAR model as special cases. This model possesses a proper joint distribution and allows for a fast update of the spatial dependence parameter. Additionally, this modification has a more reasonable behavior of the conditional variance of a spatial effect given all other spatial effects than the model considered in Sun et al. (2000).

In a hierarchical setup this extended CAR model has been used for binary spatial regression data. To capture additional heterogeneity, cluster effects have been included. In addition to the conventional modeling of heterogeneity between groups (group cluster effects) through independent random effects, modeling of heterogeneity within groups (individual cluster effects) has also been considered. A naive approach yields an unidentifiable model. It is shown how the model can be reparametrized to overcome nonidentifiability. Parameter estimation is facilitated by an MCMC approach. Separate MCMC algorithms have been developed for the two hierarchical model classes considered: logistic regression with spatial and group cluster effects and logistic regression with spatial and individual cluster effects. Probit formulations could have used as well and have been investigated in Prokopenko (2004). There latent variables are used for probit models with individual cluster effects requiring only a single MH update. This is faster because of better mixing behavior than a corresponding MCMC algorithm based on the logit formulation. A different approach to logit models is given in Holmes and Held (2004). However logit formulations are easier to interpret and therefore more often used in practice. All MCMC algorithms presented in this paper are validated through simulation. The usefulness of these models has been demonstrated by the application to a mobility study. We show that this approach is able to detect spatial regions where public transport options are more/less often used after adjusting for explanatory factors.

For model comparison, we use the sum of weighted squared residuals as a measure of fit and the number of parameters required for estimating spatial probabilities as a rough measure of model complexity in addition to PMCC and Brier score. A more theoretical based approach is still needed and of current research interest. Alternatives such as posterior predictive p-values proposed by Gelman et al. (1996) are possible, however their calibration is difficult in such complex settings (see Hjort et al. 2006).

The mobility study also included information on trips conducted by foot and bicycle which have been ignored so far. A multionomial logit (MNL) analysis without spatial and cluster effect of this data has been performed by Ehrlich (2002). Therefore we plan to extend our analysis to MNL models with spatial and cluster effects. For point location data a MNL model with spatial effects based on spatial distances has been considered by Mohammadian and Kanaroglou (2003). However many discrete choice modelers have objected to the restrictions implied by a MNL model. In particular the MNL model assumes that the random utilities are independent identically distributed and that the responsiveness to attributes of alternatives across individuals after controlling for individual characteristics is homogenous. To relax these two restrictions the generalized extreme value (GEV) class of models and the mixed multinomial logit (MMNL) class have been proposed (see for example Bhat (2002) and Bhat (2006)). Bhat and Guo (2004) consider a mixed spatially correlated logit model based on a GEV structure to accommodate correlations between spatial units of a location point referenzed data. They use the Halton simulation method (see Train (2003)) to simulate the corresponding likelihood for parameter estimation. It would be interesting to provide alternative Bayesian estimates for these dicrete choice models for point location data. In addition one can develop models following the approach taken in this paper for spatially aggregated data allowing for a spatial CAR formulation. The addition of cluster effects would provide an alternative to the MMNL model class.

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