Description

Mixed-effects models are characterized as containing both fixed effects and random effects. The fixed effects are analogous to standard regression coefficients and are estimated directly. The random effects are not directly estimated (although they may be obtained postestimation) but are summarized according to their estimated variances and covariances. Random effects may take the form of either random intercepts or random coefficients, and the grouping structure of the data may consist of multiple levels of nested groups. As such, mixed-effects models are also known in the literature as multilevel models and hierarchical models. Mixed-effects commands fit mixed-effects models for a variety of distributions of the response conditional on normally distributed random effects.

Mixed-effects linear regression

mixed  Multilevel mixed-effects linear regression

Mixed-effects generalized linear model

meglm  Multilevel mixed-effects generalized linear model

Mixed-effects censored regression

metobit  Multilevel mixed-effects tobit regression
meintreg  Multilevel mixed-effects interval regression

Mixed-effects binary regression

melogit  Multilevel mixed-effects logistic regression
meprobit  Multilevel mixed-effects probit regression
mecloglog  Multilevel mixed-effects complementary log-log regression

Mixed-effects ordinal regression

meologit  Multilevel mixed-effects ordered logistic regression
meoprobit  Multilevel mixed-effects ordered probit regression

Mixed-effects count-data regression

mepoisson  Multilevel mixed-effects Poisson regression
menbreg  Multilevel mixed-effects negative binomial regression
Mixed-effects multinomial regression
Although there is no `memlogit` command, multilevel mixed-effects multinomial logistic models can be fit using `gsem`; see [SEM] Example 41g.

Mixed-effects survival model
`mestreg` Multilevel mixed-effects parametric survival models

Nonlinear mixed-effects regression
`menl` Nonlinear mixed-effects regression

Postestimation tools specific to mixed-effects commands
- `estat df` Calculate and display degrees of freedom for fixed effects
- `estat group` Summarize the composition of the nested groups
- `estat icc` Estimate intraclass correlations
- `estat recovariance` Display the estimated random-effects covariance matrices
- `estat sd` Display variance components as standard deviations and correlations
- `estat wcorrelation` Display within-cluster correlations and standard deviations

Quick start

**Linear mixed-effects models**

Linear model of y on x with random intercepts by id

\[
\text{mixed } y \text{ x } || \text{ id: }
\]

Three-level linear model of y on x with random intercepts by doctor and patient

\[
\text{mixed } y \text{ x } || \text{ doctor: } || \text{ patient: }
\]

Linear model of y on x with random intercepts and coefficients on x by id

\[
\text{mixed } y \text{ x } || \text{ id: } x
\]

Same model with covariance between the random slope and intercept

\[
\text{mixed } y \text{ x } || \text{ id: } x, \text{ covariance(unstructured)}
\]

Linear model of y on x with crossed random effects for id and week

\[
\text{mixed } y \text{ x } || \text{ _all: R.id } || \text{ _all: R.week}
\]

Same model specified to be more computationally efficient

\[
\text{mixed } y \text{ x } || \text{ _all: R.id } || \text{ week: }
\]

Full factorial repeated-measures ANOVA of y on a and b with random effects by field

\[
\text{mixed } y \text{ a##b } || \text{ field: }
\]
Generalized linear mixed-effects models

Logistic model of $y$ on $x$ with random intercepts by id, reporting odds ratios

```
melogit y x || id: , or
```

Same model specified as a GLM

```
meglm y x || id:, family(bernoulli) link(logit)
```

Three-level ordered probit model of $y$ on $x$ with random intercepts by doctor and patient

```
meoprobit y x || doctor: || patient:
```

Nonlinear mixed-effects models

Nonlinear mixed-effects regression of $y$ on $x_1$ and $x_2$ with parameters $\{b0\}, \{b1\}, \{b2\},$ and $\{b3\}$ and random intercepts $U0$ by id

```
menl y = \{(b0)\+\{(b1)\+x1+\{U0[id]\})/(1+exp(-(x2-\{b2\})/\{b3\}))\}
```

As above, but using the more efficient specification of the linear combination

```
menl y = \{(lc: x1 U0[id])\}/(1+exp(-(x2-\{b2\})/\{b3\}))
```

As above, but using `define()` to specify the linear combination

```
menl y = \{(lc:)/(1+exp(-(x2-\{b2\})/\{b3\}))\}, define(lc: x1 U0[id])
```

Include a random slope on continuous variable $x_1$ in the `define()` option, and allow correlation between random slopes $U1$ and intercepts $U0$

```
menl y = \{(lc:)/(1+exp(-(x2-\{b2\})/\{b3\}))\},
    define(lc: x1 U0[id] c.x1#U1[id])
```

Specify a heteroskedastic within-subject error structure that varies as a power of predicted mean values $\_yhat$

```
menl y = \{(lc:)/(1+exp(-(x2-\{b2\})/\{b3\}))\},
    define(lc: x1 U0[id] c.x1#U1[id])
```

Three-level nonlinear regression of $y$ on $x_1$ with random intercepts $W0$ and slopes $W1$ on continuous $x_1$ by lev2 and with random intercepts $S0$ and slopes $S1$ on $x_1$ by lev3, with lev2 nested within lev3, using unstructured covariance for $W0$ and $W1$ and exchangeable covariance for $S0$ and $S1$

```
menl y = \{phi1:}\+\{(b1)\+cos(\{b2\}\*x1)\},
```

```
Syntax

**Linear mixed-effects models**

\[
\text{mixed depvar fe\_equation [|| re\_equation] [|| re\_equation ...] [, options]}
\]

where the syntax of the fixed-effects equation, \(fe\_equation\), is

\[
\begin{align*}
\text{[indepvars] [if] [in] [weight] [, fe\_options]}
\end{align*}
\]

and the syntax of a random-effects equation, \(re\_equation\), is the same as below for a generalized linear mixed-effects model.

**Generalized linear mixed-effects models**

\[
\text{mecmd depvar fe\_equation [|| re\_equation] [|| re\_equation ...] [, options]}
\]

where the syntax of the fixed-effects equation, \(fe\_equation\), is

\[
\begin{align*}
\text{[indepvars] [if] [in] [, fe\_options]}
\end{align*}
\]

and the syntax of a random-effects equation, \(re\_equation\), is one of the following:

- for random coefficients and intercepts
  \[
  \text{levelvar: [varlist] [, re\_options]}
  \]

- for random effects among the values of a factor variable
  \[
  \text{levelvar: R.varname}
  \]

\(levelvar\) is a variable identifying the group structure for the random effects at that level or is \(_\text{all}\) representing one group comprising all observations.

**Nonlinear mixed-effects models**

\[
\text{menl depvar = <menlexpr> [if] [in] [, options]}
\]

\(<\text{menlexpr}>\) defines a nonlinear regression function as a substitutable expression that contains model parameters and random effects specified in braces \{\}, as in \(\exp({b}+{U[id]})\); see *Random-effects substitutable expressions* in [ME] \texttt{menl} for details.
Introduction

Multilevel models have been used extensively in diverse fields, from the health and social sciences to econometrics. Mixed-effects models for binary outcomes have been used, for example, to analyze the effectiveness of toenail infection treatments (Lesaffre and Spiessens 2001) and to model union membership of young males (Vella and Verbeek 1998). Ordered outcomes have been studied by, for example, Tutz and Hennevogl (1996), who analyzed data on wine bitterness, and De Boeck and Wilson (2004), who studied verbal aggressiveness. For applications of mixed-effects models for count responses, see, for example, the study on police stops in New York City (Gelman and Hill 2007) and the analysis of the number of patents (Hall, Griliches, and Hausman 1986). Rabe-Hesketh and Skrondal (2012) provide more examples of linear and generalized linear mixed-effects models. Nonlinear mixed-effects (NLME) models are popular in, for example, population pharmacokinetics, bioassays, and studies of biological and agricultural growth processes.

For a comprehensive treatment of mixed-effects models, see, for example, Searle, Casella, and McCulloch (1992); Verbeke and Molenberghs (2000); Raudenbush and Bryk (2002); Hedeker and Gibbons (2006); McCulloch, Searle, and Neuhaus (2008); and Rabe-Hesketh and Skrondal (2012). For NLME models, see, for example, Davidian and Giltinan (1995); Vonesh and Chinchilli (1997); Demidenko (2013); Pinheiro and Bates (2000); and Davidian and Giltinan (2003).
Shayle R. Searle (1928–2013) was born in New Zealand. He obtained his PhD in animal breeding from Cornell University in 1958, with a minor in statistics. Prior to moving to New York, he worked as a research statistician for the New Zealand Dairy Board, which provided the data that he would analyze for his thesis. After completing his doctoral degree, he worked as a research associate and published several articles. He later returned to his post as a statistician in New Zealand, a position which would have a lasting influence on his career.

Through his analysis of dairy production data, Searle made advancements in estimation methods for unbalanced data and published a book on this topic. He later returned to Cornell University, teaching courses in matrix algebra, linear regression models, and estimation of variance components. Searle was one of the first few statisticians to use matrices in statistics, and he wrote a couple of books applying matrix algebra to economics and statistics. In 2001, he published a book on mixed models, which proved to be a significant contribution considering that not many statisticians were well acquainted with random effects in the 1950s. His contributions did not go unnoticed: he was awarded the Alexander von Humboldt U.S. Senior Scientist Award and was elected a fellow of the Royal Statistical Society and of the American Statistical Association.

George Casella (1951–2012) was born in Bronx, New York. After obtaining a PhD in statistics from Purdue University, he went on to join the faculty at Rutgers University, and later Cornell University, where he taught for 19 years, and the University of Florida. He published on topics such as confidence estimation, Bayesian analysis, and empirical Bayes methods. In general, his work was motivated by applications to science, and in particular, his work on variable selection and clustering was motivated by genetics. Casella coauthored a book with Roger Berger that introduced many graduate students to mathematical statistics. He coauthored another book with Christian P. Robert on Monte Carlo methods. In addition to his own published work, Casella was an editor for three journals: *Statistical Science*, *Journal of the American Statistical Society*, and *Journal of the Royal Statistical Society*.

Casella’s many contributions are reflected in his election to fellowship on behalf of four different associations and institutes and being made a foreign member of the Spanish Royal Academy of Sciences. He acquired the Spanish language during a year he spent in Spain for sabbatical and even gave talks on Monte Carlo methods in Spanish. Aside from his academic accomplishments, Casella completed 13 marathons and spent time as a volunteer firefighter.

Using mixed-effects commands

Below we summarize general capabilities of the mixed-effects commands. We let *mecmd* stand for any mixed-effects command, such as *mixed*, *melogit*, or *meprobit*, except *menl*. *menl* models the mean function nonlinearly and thus has a different syntax; see [ME] *menl*.

1. Fit a two-level random-intercept model with *levelvar* defining the second level:

2. Fit a two-level random-coefficients model containing the random-effects covariates `revars` at the level `levelvar`:

   . `mecmd depvar [indepvars] ... || levelvar: revars, ...`

   This model assumes an independent covariance structure between the random effects; that is, all covariances are assumed to be 0. There is no statistical justification, however, for imposing any particular covariance structure between random effects at the onset of the analysis. In practice, models with an unstructured random-effects covariance matrix, which allows for distinct variances and covariances between all random-effects covariates (`revars`) at the same level, must be explored first; see Other covariance structures and example 3 in [ME] `melogit` for details.

   Stata’s commands use the default independent covariance structure for computational feasibility. Numerical methods for fitting mixed-effects models are computationally intensive—computation time increases significantly as the number of parameters increases; see Computation time and the Laplacian approximation for details. The unstructured covariance is the most general and contains many parameters, which may result in an unreasonable computation time even for relatively simple random-effects models. Whenever feasible, however, you should start your statistical analysis by fitting mixed-effects models with an unstructured covariance between random effects, as we show next.

3. Specify the unstructured covariance between the random effects in the above:

   . `mecmd depvar [indepvars] ... || levelvar: revars, covariance(unstructured) ...`

4. Fit a three-level nested model with `levelvar1` defining the third level and `levelvar2` defining the second level:

   . `mecmd depvar [indepvars] ... || levelvar1: || levelvar2: ..., ...`

5. Fit the above three-level nested model as a two-level model with exchangeable covariance structure at the second level (mixed only):

   . `mecmd depvar [indepvars] ... || levelvar1: R.levelvar2, cov(exchangeable) ...`

   See example 11 in [ME] `mixed` for details about this equivalent specification. This specification may be useful for a more efficient fitting of random-effects models with a mixture of crossed and nested effects.

6. Fit higher-level nested models:

   . `mecmd depvar [indepvars] ... || levelvar1: || levelvar2: || levelvar3: || ...`

7. Fit a two-way crossed-effects model with the `_all:` notation for each random-effects equation:

   . `mecmd depvar [indepvars] ... || _all: R.factor1 || _all: R.factor2 ...`

   When you use the `_all:` notation for each random-effects equation, the total dimension of the random-effects design equals $r_1 + r_2$, where $r_1$ and $r_2$ are the numbers of levels in `factor1` and `factor2`, respectively. This specification may be infeasible for some mixed-effects models; see item 8 below for a more efficient specification of this model.

8. Fit a two-way crossed-effects model with the `_all:` notation for the first random-effects equation only:

   . `mecmd depvar [indepvars] ... || _all: R.factor1 || factor2:, ...`

   Compared with the specification in item 7, this specification requires only $r_1 + 1$ parameters and is thus more efficient; see Crossed-effects models for details.
9. Fit a two-way full-factorial random-effects model:

```
.mecmd depvar [indepvars] ... || _all: R.factor1 || factor2: || factor1: ...
```

10. Fit a two-level mixed-effects model with a blocked-diagonal covariance structure between `revars1` and `revars2`:

```
.mecmd depvar [indepvars] ... || levelvar: revars1, noconstant ///
|| levelvar: revars2, noconstant ...
```

11. Fit a linear mixed-effects model where the correlation between the residual errors follows an autoregressive process of order 1:

```
.mixed depvar [indepvars] ... || levelvar:, residuals(ar 1, t(time)) ...
```

More residual error structures are available; see `[ME] mixed` for details.

12. Fit a two-level linear mixed-effects model accounting for sampling weights `expr1` at the first (residual) level and for sampling weights `expr2` at the level of `levelvar`:

```
mixed depvar [indepvars] [pweight=expr1] ... || levelvar:, pweight(expr2) ...
```

Mixed-effects commands—with the exception of `mixed`—allow constraints on both fixed-effects and random-effects parameters. We provide several examples below of imposing constraints on variance components.

13. Fit a mixed-effects model with the variance of the random intercept on `levelvar` constrained to be 16:

```
constraint 1 _b[var(_cons[levelvar]):_cons]=16
.mecmd depvar [indepvars] ... || levelvar:, constraints(1) ...
```

14. Fit a mixed-effects model with the variance of the random intercept on `levelvar` and the variance of the random slope on `revar` to be equal:

```
constraint 1 _b[var(revar[levelvar]):_cons] = _b[var(_cons[levelvar]):_cons]
mecmd depvar [indepvars] ... || levelvar: revar, constraints(1) ...
```

Note that the constraints above are equivalent to imposing an identity covariance structure for the random-effects equation:

```
mecmd depvar [indepvars] ... || levelvar: revar, cov(identity) ...
```

15. Assuming four random slopes `revars`, fit a mixed-effects model with the variance components at the level of `levelvar` constrained to have a banded structure:

```
.mat p = (1,.,.,. \ 2,1,.,. \ 3,2,1,. \ 4,3,2,1)
mecmd depvar [indepvars] ... || levelvar: revars, noconstant ///
covariance(pattern(p)) ...
```

16. Assuming four random slopes `revars`, fit a mixed-effects model with the variance components at the level of `levelvar` constrained to the specified numbers, and with all the covariances constrained to be 0:

```
.mat f = diag((1,2,3,4))
mecmd depvar [indepvars] ... || levelvar: revars, noconstant ///
covariance(fixed(f)) ...
```

The variance components in models in items 15 and 16 can also be constrained by using the `constraints()` option, but using `covariance(pattern())` or `covariance(fixed())` is more convenient.
**Mixed-effects models**

**Linear mixed-effects models**

Linear mixed-effects (LME) models for continuous responses are a generalization of linear regression allowing for the inclusion of random deviations (effects) other than those associated with the overall error term. In matrix notation,

\[ y = X\beta + Zu + \epsilon \]  

(1)

where \( y \) is the \( n \times 1 \) vector of responses, \( X \) is an \( n \times p \) design/covariate matrix for the fixed effects \( \beta \), and \( Z \) is the \( n \times q \) design/covariate matrix for the random effects \( u \). The \( n \times 1 \) vector of errors \( \epsilon \) is assumed to be multivariate normal with mean 0 and variance matrix \( \sigma^2_\epsilon R \).

The fixed portion of (1), \( X\beta \), is analogous to the linear predictor from a standard OLS regression model with \( \beta \) being the regression coefficients to be estimated. For the random portion of (1), \( Zu + \epsilon \), we assume that \( u \) has variance–covariance matrix \( G \) and that \( u \) is orthogonal to \( \epsilon \) so that

\[
\text{Var} \left[ \begin{bmatrix} u \\ \epsilon \end{bmatrix} \right] = \begin{bmatrix} G & 0 \\ 0 & \sigma^2_\epsilon R \end{bmatrix}
\]

The random effects \( u \) are not directly estimated (although they may be predicted) but instead are characterized by the elements of \( G \), known as variance components, that are estimated along with the error-covariance parameters that include the overall error variance \( \sigma^2_\epsilon \) and the parameters that are contained within \( R \).

The general forms of the design matrices \( X \) and \( Z \) allow estimation for a broad class of linear models: blocked designs, split-plot designs, growth curves, multilevel or hierarchical designs, etc. They also allow a flexible method of modeling within-cluster correlation. Subjects within the same cluster can be correlated as a result of a shared random intercept, or through a shared random slope on age (for example), or both. The general specification of \( G \) also provides additional flexibility: the random intercept and random slope could themselves be modeled as independent, or correlated, or independent with equal variances, and so forth. The general structure of \( R \) also allows for within-cluster errors to be heteroskedastic and correlated and allows flexibility in exactly how these characteristics can be modeled.

In clustered-data situations, it is convenient not to consider all \( n \) observations at once but instead to organize the mixed model as a series of \( M \) independent groups (or clusters)

\[ y_j = X_j\beta + Z_j u_j + \epsilon_j \]  

(2)

for \( j = 1, \ldots, M \), with cluster \( j \) consisting of \( n_j \) observations. The response \( y_j \) comprises the rows of \( y \) corresponding with the \( j \)th cluster, with \( X_j \) and \( \epsilon_j \) defined analogously. The random effects \( u_j \) can now be thought of as \( M \) realizations of a \( q \times 1 \) vector that is normally distributed with mean 0 and \( q \times q \) variance matrix \( \Sigma \). The matrix \( Z_j \) is the \( n_j \times q \) design matrix for the \( j \)th cluster random effects. Relating this to (1),

\[
Z = \begin{bmatrix} Z_1 & 0 & \cdots & 0 \\ 0 & Z_2 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & Z_M \end{bmatrix}; \quad u = \begin{bmatrix} u_1 \\ \vdots \\ u_M \end{bmatrix}; \quad G = I_M \otimes \Sigma; \quad R = I_M \otimes \Lambda
\]

where \( \Lambda \) denotes the variance matrix of the level-1 errors and \( \otimes \) is the Kronecker product.
The mixed-model formulation (2) is from Laird and Ware (1982) and offers two key advantages. First, it makes specifications of random-effects terms easier. If the clusters are schools, you can simply specify a random effect at the school level, as opposed to thinking of what a school-level random effect would mean when all the data are considered as a whole (if it helps, think Kronecker products). Second, representing a mixed-model with (2) generalizes easily to more than one set of random effects. For example, if classes are nested within schools, then (2) can be generalized to allow random effects at both the school and the class-within-school levels.

By our convention on counting and ordering model levels, (2) is a two-level model, with extensions to three, four, or any number of levels. The observation $y_{ij}$ is for individual $i$ within cluster $j$, and the individuals compose the first level, whereas the clusters compose the second level of the model. In a hypothetical three-level model with classes nested within schools, the observations within classes (the students, presumably) would constitute the first level, the classes would constitute the second level, and the schools would constitute the third level. This differs from certain citations in the classical ANOVA literature and texts such as Pinheiro and Bates (2000) but is the standard in the vast literature on hierarchical models, for example, Skrondal and Rabe-Hesketh (2004).

In Stata, you can use `mixed` to fit linear mixed-effects models; see `[ME] mixed` for a detailed discussion and examples. Various predictions, statistics, and diagnostic measures are available after fitting an LME model with `mixed`. For the most part, calculation centers around obtaining estimates of random effects; see `[ME] mixed postestimation` for a detailed discussion and examples.

### Generalized linear mixed-effects models

Generalized linear mixed-effects (GLME) models, also known as generalized linear mixed models (GLMMs), are extensions of generalized linear models allowing for the inclusion of random deviations (effects). In matrix notation,

$$g\left\{E(y|X,u)\right\} = X\beta + Zu, \quad y \sim F$$

(3)

where $y$ is the $n \times 1$ vector of responses from the distributional family $F$, $X$ is an $n \times p$ design/covariate matrix for the fixed effects $\beta$, and $Z$ is an $n \times q$ design/covariate matrix for the random effects $u$. The $X\beta + Zu$ part is called the linear predictor and is often denoted as $\eta$. $g(\cdot)$ is called the link function and is assumed to be invertible such that

$$E(y|u) = g^{-1}(X\beta + Zu) = H(\eta) = \mu$$

For notational convenience here and throughout this manual entry, we suppress the dependence of $y$ on $X$. Substituting various definitions for $g(\cdot)$ and $F$ results in a wide array of models. For instance, if $g(\cdot)$ is the logit function and $y$ is distributed as Bernoulli, we have

$$\text{logit}\left\{E(y|u)\right\} = X\beta + Zu, \quad y \sim \text{Bernoulli}$$

or mixed-effects logistic regression. If $g(\cdot)$ is the natural log function and $y$ is distributed as Poisson, we have

$$\ln\left\{E(y|u)\right\} = X\beta + Zu, \quad y \sim \text{Poisson}$$

or mixed-effects Poisson regression.

For the random portion of (3), $Zu$, we assume that $u$ has variance–covariance matrix $G$ such that

$$\text{Var}(u) = G$$
The random effects \( u \) are not directly estimated (although they may be predicted) but instead are characterized by the elements of \( G \), known as variance components.

Analogously to (2), in clustered-data situations, we can write

\[
E(y_j | u_j) = g^{-1}(X_j \beta + Z_j u_j) \quad y_j \sim F
\]

with all the elements defined as before. In terms of the whole dataset, we now have

\[
Z = \begin{bmatrix}
Z_1 & 0 & \cdots & 0 \\
0 & Z_2 & \cdots & 0 \\
\vdots & \vdots & \ddots & \vdots \\
0 & 0 & \cdots & Z_M
\end{bmatrix}; \quad u = \begin{bmatrix}
u_1 \\
\vdots \\
u_M
\end{bmatrix}; \quad G = I_M \otimes \Sigma
\]

In Stata, you can use \texttt{meglm} to fit mixed-effects models for nonlinear responses. Some combinations of families and links are so common that we implemented them as separate commands in terms of \texttt{meglm}.

<table>
<thead>
<tr>
<th>Command</th>
<th>\texttt{meglm} equivalent</th>
</tr>
</thead>
<tbody>
<tr>
<td>melogit</td>
<td>family(bernoulli) link(logit)</td>
</tr>
<tr>
<td>meprobit</td>
<td>family(bernoulli) link(probit)</td>
</tr>
<tr>
<td>mecloglog</td>
<td>family(bernoulli) link(cloglog)</td>
</tr>
<tr>
<td>meologit</td>
<td>family(ordinal) link(logit)</td>
</tr>
<tr>
<td>meoprobit</td>
<td>family(ordinal) link(probit)</td>
</tr>
<tr>
<td>meopoisson</td>
<td>family(poisson) link(log)</td>
</tr>
<tr>
<td>menybreg</td>
<td>family(nbinomial) link(log)</td>
</tr>
</tbody>
</table>

When no family–link combination is specified, \texttt{meglm} defaults to a Gaussian family with an identity link. Thus \texttt{meglm} can be used to fit linear mixed-effects models; however, for those models we recommend using the more specialized \texttt{mixed}, which, in addition to \texttt{meglm} capabilities, allows for modeling of the structure of the within-cluster errors; see \cite{me} for details.

Various predictions, statistics, and diagnostic measures are available after fitting a GLME model with \texttt{meglm} and other \texttt{me} commands. For the most part, calculation centers around obtaining estimates of random effects; see \cite{me} \texttt{meglm postestimation} for a detailed discussion and examples.

### Survival mixed-effects models

Parametric survival mixed-effects models use a trivariate response variable \((t_0, t, d)\), where each response corresponds to a period under observation \((t_0, t]\) and results in either failure \((d = 1)\) or right-censoring \((d = 0)\) at time \(t\). See \cite{st} \texttt{streg} for background information on parametric survival models. Two often-used models for adjusting survivor functions for the effects of covariates are the accelerated failure-time (AFT) model and the multiplicative or proportional hazards (PH) model.

In the AFT parameterization, the natural logarithm of the survival time, \(\log t\), is expressed as a linear function of the covariates. When we incorporate random effects, this yields the model

\[
\log(t_j) = X_j \beta + Z_j u_j + v_j
\]
where \( \log(\cdot) \) is an elementwise function, and \( v_j \) is a vector of observation-level errors. The distributional form of the error term determines the regression model.

In the PH model, the covariates have a multiplicative effect on the hazard function

\[
h(t_j) = h_0(t_j) \exp(X_j \beta + Z_j u_j)
\]

where all the functions are elementwise, and \( h_0(\cdot) \) is a baseline hazard function. The functional form of \( h_0(\cdot) \) determines the regression model.

In Stata, you can use \texttt{mestreg} to fit multilevel mixed-effects parametric survival models for the following distributions and parameterizations.

<table>
<thead>
<tr>
<th>Distribution</th>
<th>Parameterization</th>
</tr>
</thead>
<tbody>
<tr>
<td>exponential</td>
<td>PH, AFT</td>
</tr>
<tr>
<td>loglogistic</td>
<td>AFT</td>
</tr>
<tr>
<td>weibull</td>
<td>PH, AFT</td>
</tr>
<tr>
<td>lognormal</td>
<td>AFT</td>
</tr>
<tr>
<td>gamma</td>
<td>AFT</td>
</tr>
</tbody>
</table>

\texttt{mestreg} is suitable only for data that have been set using the \texttt{stset} command. By using \texttt{stset} on your data, you define the variables \_t0, \_t, and \_d, which serve as the trivariate response. See [ME] \texttt{mestreg} for more details about the command. Various predictions, statistics, and diagnostic measures are available after fitting a mixed-effects survival model with \texttt{mestreg}; see [ME] \texttt{mestreg postestimation} for a detailed discussion and examples.

**Nonlinear mixed-effects models**

NLME models are models containing both fixed effects and random effects where some of, or all, the fixed and random effects enter the model nonlinearly. They can be viewed as a generalization of LME models, in which the conditional mean of the outcome given the random effects is a nonlinear function of the coefficients and random effects. Alternatively, they can be considered as an extension of nonlinear regression models for independent data (see [R] \texttt{nl}), in which coefficients may incorporate random effects, allowing them to vary across different levels of hierarchy and thus inducing correlation within observations at the same level.

Using the notation from \textit{Linear mixed-effects models} for LME models for clustered data, we can write an NLME model as

\[
y_j = \mu(A_j, \beta, u_j) + \epsilon_j
\]

where \( \mu(\cdot) \) is a real-valued vector function and \( A_j \) is an \( n_j \times l \) matrix of covariates for the \( j \)th cluster, which includes both within-subject and between-subject covariates. Do not be surprised to see the \( A_j \) matrix here instead of the more familiar fixed-effects and random-effects design matrices \( X_j \) and \( Z_j \) from previous sections. Because both covariates and parameters can enter the model nonlinearly in NLME, we cannot express the regression function as a function containing the linear term \( X_j \beta + Z_j u_j \) as we can for LME and GLME models. The distributional assumptions on \( u_j \)’s and \( \epsilon_j \)’s are the same as for the LME models.

Parameters of NLME models often have scientifically meaningful interpretations, and research questions are formed based on them. To allow parameters to reflect phenomena of interest, NLME models are often formulated by using a multistage formulation; see Alternative mixed-effects model specification below for examples.
We can formulate our previous NLME model as a two-stage hierarchical model:

**Stage 1:** Individual-level model

\[ y_{ij} = m(x_{ij}^w, \phi_j) + \epsilon_{ij}, \quad i = 1, \ldots, n \]

**Stage 2:** Group-level model

\[ \phi_j = d(x_{bj}, \beta, u_j), \quad j = 1, \ldots, M \]

In stage 1, we model the response by using a function \( m(\cdot) \), which describes within-subject behavior. This function depends on subject-specific parameters \( \phi_j \)'s, which have a natural physical interpretation, and a vector of within-subject covariates \( x_{ij}^w \). In stage 2, we use a known vector-valued function \( d(\cdot) \) to model between-subject behavior, that is, to model \( \phi_j \)'s and to explain how they vary across subjects. The \( d(\cdot) \) function incorporates random effects and, optionally, a vector of between-subject covariates \( x_{bj} \). The general idea is to specify a common functional form for each subject in stage 1 and then allow some parameters to vary randomly across subjects in stage 2.

You can use the `menl` command to fit NLME models to continuous outcomes; see [ME] menl. `menl` supports both the single-equation and multistage model formulations. It supports different covariance structures for random effects and can model heteroskedasticity and correlations within lowest-level groups. Various predictions, statistics, and diagnostic measures are available after fitting an NLME model; see [ME] menl postestimation.

For an introductory example, see Nonlinear models.

### Alternative mixed-effects model specification

In this section, we present a hierarchical or multistage formulation of mixed-effects models where each level is described by its own set of equations. This formulation is common for NLME models; see Nonlinear mixed-effects models.

Consider a random-intercept model that we write here in general terms:

\[ y_{ij} = \beta_0 + \beta_1 x_{ij} + u_j + \epsilon_{ij} \quad (4) \]

This single-equation specification contains both level-1 and level-2 effects. In the hierarchical form, we specify a separate equation for each level.

\[ y_{ij} = \gamma_{0j} + \beta_1 x_{ij} + \epsilon_{ij} \]

\[ \gamma_{0j} = \beta_{00} + u_{0j} \quad (5) \]

The equation for the intercept \( \gamma_{0j} \) consists of the overall mean intercept \( \beta_{00} \) and a cluster-specific random intercept \( u_{0j} \). To fit this model by using, for example, `mixed`, we must translate the multiple-equation notation into a single-equation form. We substitute the second equation into the first one and rearrange terms.

\[ y_{ij} = \beta_{00} + u_{0j} + \beta_1 x_{ij} + \epsilon_{ij} \]

\[ = \beta_{00} + \beta_1 x_{ij} + u_{0j} + \epsilon_{ij} \quad (6) \]

Note that model (6) is the same as model (4) with \( \beta_{00} \equiv \beta_0 \) and \( u_{0j} \equiv u_j \). Thus the syntax for our generic random-intercept model is

```
. mixed y x || id:
```

where `id` is the variable designating the clusters.
We can extend model (5) to include a random slope. We do so by specifying an additional equation for the slope on $x_{ij}$.

$$y_{ij} = \gamma_0 + \gamma_1 x_{ij} + \epsilon_{ij}$$

$$\gamma_0 = \beta_0 + u_0$$

$$\gamma_1 = \beta_1 + u_1$$  \hspace{1cm} (7)

The additional equation for the slope $\gamma_{1j}$ consists of the overall mean slope $\beta_{10}$ and a cluster-specific random slope $u_{1j}$. We substitute the last two equations into the first one to obtain a reduced-form model.

$$y_{ij} = (\beta_0 + u_0) + (\beta_{10} + u_{1j})x_{ij} + \epsilon_{ij}$$

$$= \beta_0 + \beta_{10}x_{ij} + u_0 + u_{1j}x_{ij} + \epsilon_{ij}$$

The syntax for this model becomes

```
.mixed y x || id: x, covariance(unstructured)
```

where we specified an unstructured covariance structure for the level-2 $u$ terms.

Here we further extend the random-slope random-intercept model (7) by adding a level-2 covariate $z_j$ into the level-2 equations.

$$y_{ij} = \gamma_0 + \gamma_1 x_{ij} + \epsilon_{ij}$$

$$\gamma_0 = \beta_0 + \beta_{01}z_j + u_0$$

$$\gamma_1 = \beta_1 + \beta_{11}z_j + u_1$$

We substitute as before to obtain a single-equation form:

$$y_{ij} = (\beta_0 + \beta_{01}z_j + u_0) + (\beta_{10} + \beta_{11}z_j + u_{1j})x_{ij} + \epsilon_{ij}$$

$$= \beta_0 + \beta_{01}z_j + \beta_{10}x_{ij} + \beta_{11}z_jx_{ij} + u_0 + u_{1j}x_{ij} + \epsilon_{ij}$$

Now the fixed-effects portion of the equation contains a constant and variables $x$, $z$, and their interaction. Assuming both $x$ and $z$ are continuous variables, we can use the following Stata syntax to fit this model:

```
.mixed y x z c.x#c.z || id: x, covariance(unstructured)
```

Although the `menl` command is not as suitable for fitting LME models as `mixed`, it can accommodate a multistage formulation. For example, (5) can be fit in `menl` as

```
.menl y = {gamma0:}+{b1}*x, define(gamma0: {b00}+{U0[id]})
```

and (7) as

```
.menl y = {gamma0:}+{gamma1:}*x, define(gamma0: {b00}+{U0[id]}) ///
    define(gamma1: {b10}+{U1[id]})
```

In the above `menl`’s specifications, `gamma0` and `gamma1` can be specified more efficiently by using linear combinations; see [ME] menl for details.

We refer you to Raudenbush and Bryk (2002) and Rabe-Hesketh and Skrondal (2012) for a more thorough discussion and further examples of multistage mixed-model formulations, including three-level models.
The key to fitting mixed models lies in estimating the variance components, and for that there exist many methods. Most of the early literature in LME models dealt with estimating variance components in ANOVA models. For simple models with balanced data, estimating variance components amounts to solving a system of equations obtained by setting expected mean-squares expressions equal to their observed counterparts. Much of the work in extending the ANOVA method to unbalanced data for general ANOVA designs is attributed to Henderson (1953).

The ANOVA method, however, has its shortcomings. Among these is a lack of uniqueness in that alternative, unbiased estimates of variance components could be derived using other quadratic forms of the data in place of observed mean squares (Searle, Casella, and McCulloch 1992, 38–39). As a result, ANOVA methods gave way to more modern methods, such as minimum norm quadratic unbiased estimation (MINQUE) and minimum variance quadratic unbiased estimation (MIVQUE); see Rao (1973) for MINQUE and LaMotte (1973) for MIVQUE. Both methods involve finding optimal quadratic forms of the data that are unbiased for the variance components.

Stata uses maximum likelihood (ML) to fit LME and GLME models. The ML estimates are based on the usual application of likelihood theory, given the distributional assumptions of the model. In addition, for linear mixed-effects models, mixed offers the method of restricted maximum likelihood (REML). The basic idea behind REML (Thompson 1962) is that you can form a set of linear contrasts of the response that do not depend on the fixed effects $\beta$ but instead depend only on the variance components to be estimated. You then apply ML methods by using the distribution of the linear contrasts to form the likelihood; see the Methods and formulas section of [ME] mixed for a detailed discussion of ML and REML methods in the context of linear mixed-effects models.

Log-likelihood calculations for fitting any mixed-effects model require integrating out the random effects. For LME models, this integral has a closed-form solution; for GLME and NLME models, it does not. In dealing with this difficulty, early estimation methods avoided the integration altogether. Two such popular methods are the closely related penalized quasilikelihood (PQL) and marginal quasilikelihood (MQL) (Breslow and Clayton 1993). Both PQL and MQL use a combination of iterative reweighted least squares (see [R] glm) and standard estimation techniques for fitting LME models. Efficient computational methods for fitting LME models have existed for some time (Bates and Pinheiro 1998; Littell et al. 2006), and PQL and MQL inherit this computational efficiency. However, both of these methods suffer from two key disadvantages. First, they have been shown to be biased, and this bias can be severe when clusters are small or intracluster correlation is high (Rodríguez and Goldman 1995; Lin and Breslow 1996). Second, because they are “quasilikelihood” methods and not true likelihood methods, their use prohibits comparing nested models via likelihood-ratio (LR) tests, blocking the main avenue of inference involving variance components.

The advent of modern computers has brought with it the development of more computationally intensive methods, such as bias-corrected PQL (Lin and Breslow 1996), Bayesian Markov-Chain Monte Carlo, and simulated maximum likelihood, just to name a few; see Ng et al. (2006) for a discussion of these alternate strategies (and more) for mixed-effects models for binary outcomes.

One widely used modern method is to directly estimate the integral required to calculate the log likelihood by Gauss–Hermite quadrature or some variation thereof. Because the log likelihood itself is estimated, this method has the advantage of permitting LR tests for comparing nested models. Also, if done correctly, quadrature approximations can be quite accurate, thus minimizing bias. Stata commands for fitting GLME models such as meglm support three types of Gauss–Hermite quadratures: mean–variance adaptive Gauss–Hermite quadrature (MVAGH), mode-curvature adaptive Gauss–Hermite quadrature (MCAGH), and nonadaptive Gauss–Hermite quadrature (GHQ); see Methods and formulas of [ME] meglm for a detailed discussion of these quadrature methods. A fourth method, the Laplacian approximation, that does not involve numerical integration is also offered; see Computation time
and the Laplacian approximation below and Methods and formulas of [ME] \textit{meglm} for a detailed discussion of the Laplacian approximation method.

In the context of NLME models, the use of an adaptive quadrature to fit these models can be often computationally infeasible. A popular alternative method used to fit NLME models is the linearization method of Lindstrom and Bates (1990), also known as the conditional first-order linearization method. It is based on a first-order Taylor-series approximation of the mean function and essentially linearizes the mean function with respect to fixed and random effects. The linearization method is computationally efficient because it avoids the intractable integration, but the approximation cannot be made arbitrarily accurate. Despite its potential limiting accuracy, the linearization method has proven the most popular in practice (Fitzmaurice et al. 2009, sec 5.4.8). The linearization method of Lindstrom and Bates (1990), with extensions from Pinheiro and Bates (1995), is the method of estimation in \textit{menl}.

\section*{Computation time and the Laplacian approximation}

Like many programs that fit generalized linear mixed models, \texttt{me} commands can be computationally intensive. This is particularly true for large datasets with many lowest-level clusters, models with many random coefficients, models with many estimable parameters (both fixed effects and variance components), or any combination thereof.

Computation time will also depend on hardware and other external factors but in general is (roughly) a function of $p^2\{M + M(N_Q)^{q_t}\}$, where $p$ is the number of estimable parameters, $M$ is the number of lowest-level (smallest) clusters, $N_Q$ is the number of quadrature points, and $q_t$ is the total dimension of the random effects, that is, the total number of random intercepts and coefficients at all levels.

For a given model and a given dataset, the only prevailing factor influencing computation time is $(N_Q)^{q_t}$. However, because this is a power function, this factor can get prohibitively large. For example, using five quadrature points for a model with one random intercept and three random coefficients, we get $(N_Q)^{q_t} = 5^4 = 625$. Even a modest increase to seven quadrature points would increase this factor by almost fourfold $(7^4 = 2,401)$, which, depending on $M$ and $p$, could drastically slow down estimation. When fitting mixed-effects models, you should always assess whether the approximation is adequate by refitting the model with a larger number of quadrature points. If the results are essentially the same, the lower number of quadrature points can be used.

However, we do not deny a tradeoff between speed and accuracy, and in that spirit we give you the option to choose a (possibly) less accurate solution in the interest of getting quicker results. Toward this end is the limiting case of $N_Q = 1$, otherwise known as the Laplacian approximation; see Methods and formulas of [ME] \textit{meglm}. The computational benefit is evident—1 raised to any power equals 1—and the Laplacian approximation has been shown to perform well in certain situations (Liu and Pierce 1994; Tierney and Kadane 1986). When using Laplacian approximation, keep the following in mind:

1. Fixed-effects parameters and their standard errors are well approximated by the Laplacian method. Therefore, if your interest lies primarily here, then the Laplacian approximation may be a viable alternative.
2. Estimates of variance components exhibit bias, particularly the variances.
3. The model log likelihood and comparison LR test are in fair agreement with statistics obtained via quadrature methods.

Although this is by no means the rule, we find the above observations to be fairly typical based on our own experience. Pinheiro and Chao (2006) also make observations similar to points 1 and 2 on the basis of their simulation studies: bias due to Laplace (when present) tends to exhibit itself.
more in the estimated variance components than in the estimates of the fixed effects as well as at the lower levels in higher-level models.

**Item 3** is of particular interest, because it demonstrates that the Laplacian approximation can produce a decent estimate of the model log likelihood. Consequently, you can use the Laplacian approximation during the model building phase of your analysis, during which you are comparing competing models by using LR tests. Once you settle on a parsimonious model that fits well, you can then increase the number of quadrature points and obtain more accurate parameter estimates for further study.

Of course, sometimes the Laplacian approximation will perform either better or worse than observed here. This behavior depends primarily on cluster size and intracluster correlation, but the relative influence of these factors is unclear. The idea behind the Laplacian approximation is to approximate the posterior density of the random effects given the response with a normal distribution; see Methods and formulas of [ME] `meglm`. Asymptotic theory dictates that this approximation improves with larger clusters. Of course, the key question, as always, is “How large is large enough?” Also, there are data situations where the Laplacian approximation performs well even with small clusters. Therefore, it is difficult to make a definitive call as to when you can expect the Laplacian approximation to yield accurate results across all aspects of the model.

In conclusion, consider our above advice as a rule of thumb based on empirical evidence.

### Diagnosing convergence problems

Given the flexibility of mixed-effects models, you will find that some models fail to converge when used with your data. The default gradient-based method used by mixed-effects commands, except `menl`, is the Newton–Raphson algorithm, requiring the calculation of a gradient vector and Hessian (second-derivative) matrix; see [R] `ml`.

A failure to converge can take any one of three forms:

1. repeated nonconcave or backed-up iterations without convergence;
2. a Hessian (second-derivative) calculation that has become asymmetric, unstable, or has missing values; or
3. the message “standard-error calculation has failed” when computing standard errors.

All three situations essentially amount to the same thing: the Hessian calculation has become unstable, most likely because of a ridge in the likelihood function, a subsurface of the likelihood in which all points give the same value of the likelihood and for which there is no unique solution.

Such behavior is usually the result of one of the following two situations:

A. A model that is not identified given the data, for example, fitting the three-level nested random intercept model

\[ y_{jk} = x_{jk} \beta + u^{(3)}_k + u^{(2)}_{jk} + \epsilon_{jk} \]

without any replicated measurements at the \((j, k)\) level, that is, with only one \(i\) per \((j, k)\) combination. This model is unidentified for such data because the random intercepts \(u^{(2)}_{jk}\) are confounded with the overall errors \(\epsilon_{jk}\).

B. A model that contains a variance component whose estimate is really close to 0. When this occurs, a ridge is formed by an interval of values near 0, which produce the same likelihood and look equally good to the optimizer.
For LME models, one useful way to diagnose problems of nonconvergence is to rely on the expectation-maximization (EM) algorithm (Dempster, Laird, and Rubin 1977), normally used by mixed only as a means of refining starting values; see Diagnosing convergence problems of [ME] mixed for details.

If your data and model are nearly unidentified, as opposed to fully unidentified, you may be able to obtain convergence with standard errors by changing some of the settings of the gradient-based optimization. Adding the difficult option can be particularly helpful if you are seeing many “nonconcave” messages; you may also consider changing the technique() or using the nonrtolerance option; see [R] Maximize.

Regardless of how the convergence problem revealed itself, you may try to obtain better starting values; see Obtaining better starting values in [ME] meglm for details.

Achieving convergence and diagnosing convergence problems can be even more challenging with NLME models. As with other mixed-effects models, complicated variance–covariance structures for random effects and errors can often lead to overparameterized models that fail to converge. In addition, highly nonlinear mean specifications can lead to multiple solutions and thus to potential convergence to a local maximum. menl uses the linearization estimation method that alternates between the penalized least-squares estimation of the fixed-effects parameters and the Newton–Raphson estimation of the random-effects parameters of the approximating LME model, which was the result of the linearization of the original NLME model. This alternating method does not provide a joint Hessian matrix for all parameters, so there is no check for the tolerance of the scaled gradient, and thus the convergence cannot be established in its strict sense. The convergence is declared based on the stopping rules described in Methods and formulas of [ME] menl. Exploring different initial values to investigate convergence is particularly important with NLME models; see Obtaining initial values in [ME] menl.

**Distribution theory for likelihood-ratio test**

When determining the asymptotic distribution of an LR test comparing two nested mixed-effects models, issues concerning boundary problems imposed by estimating strictly positive quantities (that is, variances) can complicate the situation. For example, when performing LR tests involving linear mixed-effects models (whether comparing with linear regression within mixed or comparing two separate linear mixed-effects models with lrtest), you may thus sometimes see a test labeled as chibar rather than the usual chi2, or you may see a chi2 test with a note attached stating that the test is conservative or possibly conservative depending on the hypothesis being tested.

At the heart of the issue is the number of variances being restricted to 0 in the reduced model. If there are none, the usual asymptotic theory holds, and the distribution of the test statistic is \(\chi^2\) with degrees of freedom equal to the difference in the number of estimated parameters between both models.

When there is only one variance being set to 0 in the reduced model, the asymptotic distribution of the LR test statistic is a 50:50 mixture of a \(\chi_p^2\) and a \(\chi_{p+1}^2\) distribution, where \(p\) is the number of other restricted parameters in the reduced model that are unaffected by boundary conditions. Stata labels such test statistics as chibar and adjusts the significance levels accordingly. See Self and Liang (1987) for the appropriate theory or Gutierrez, Carter, and Drukker (2001) for a Stata-specific discussion.

When more than one variance parameter is being set to 0 in the reduced model, however, the situation becomes more complicated. For example, consider a comparison test versus linear regression for a mixed model with two random coefficients and unstructured covariance matrix
Because the random component of the mixed model comprises three parameters \((\sigma^2_0, \sigma_{01}, \sigma^2_1)\), on the surface it would seem that the LR comparison test would be distributed as \(\chi^2_3\). However, two complications need to be considered. First, the variances \(\sigma^2_0\) and \(\sigma^2_1\) are restricted to be positive, and second, constraints such as \(\sigma^2_1 = 0\) implicitly restrict the covariance \(\sigma_{01}\) to be 0 as well. From a technical standpoint, it is unclear how many parameters must be restricted to reduce the model to linear regression.

Because of these complications, appropriate and sufficiently general distribution theory for the more-than-one-variance case has yet to be developed. Theory (for example, Stram and Lee [1994]) and empirical studies (for example, McLachlan and Basford [1988]) have demonstrated that, whatever the distribution of the LR test statistic, its tail probabilities are bounded above by those of the \(\chi^2\) distribution with degrees of freedom equal to the full number of restricted parameters (three in the above example).

The mixed and me commands use this reference distribution, the \(\chi^2\) with full degrees of freedom, to produce a conservative test and place a note in the output labeling the test as such. Because the displayed significance level is an upper bound, rejection of the null hypothesis based on the reported level would imply rejection on the basis of the actual level.

**Examples**

**Two-level models**

Example 1: Growth-curve model

Consider a longitudinal dataset, used by both Ruppert, Wand, and Carroll (2003) and Diggle et al. (2002), consisting of weight measurements of 48 pigs on 9 successive weeks. Pigs are identified by the variable id. Each pig experiences a linear trend in growth, but overall weight measurements vary from pig to pig. Because we are not really interested in these particular 48 pigs per se, we instead treat them as a random sample from a larger population and model the between-pig variability as a random effect, or in the terminology of (2), as a random-intercept term at the pig level. We thus wish to fit the model

\[
\text{weight}_{ij} = \beta_0 + \beta_1 \text{week}_{ij} + u_j + \epsilon_{ij}
\]

for \(i = 1, \ldots, 9\) weeks and \(j = 1, \ldots, 48\) pigs. The fixed portion of the model, \(\beta_0 + \beta_1 \text{week}_{ij}\), simply states that we want one overall regression line representing the population average. The random effect \(u_j\) serves to shift this regression line up or down according to each pig. Because the random effects occur at the pig level (id), we fit the model by typing.
. use https://www.stata-press.com/data/r16/pig
(Longitudinal analysis of pig weights)
. mixed weight week || id:
Performing EM optimization:
Performing gradient-based optimization:
Iteration 0:  log likelihood = -1014.9268
Iteration 1:  log likelihood = -1014.9268
Computing standard errors:
Mixed-effects ML regression
Group variable: id
Number of obs = 432
Number of groups = 48
Obs per group:
  min = 9
  avg = 9.0
  max = 9
Wald chi2(1) = 25337.49
Log likelihood = -1014.9268
Prob > chi2 = 0.0000

|      | Coef.  | Std. Err. | z    | P>|z|  | [95% Conf. Interval] |
|------|--------|-----------|------|------|---------------------|
| week | 6.209896 | 0.0390124 | 159.18 | 0.000 | 6.133433  6.286359 |
| _cons| 19.35561 | 0.5974059 | 32.40 | 0.000 | 18.18472  20.52651 |

Random-effects Parameters
<table>
<thead>
<tr>
<th></th>
<th>Estimate</th>
<th>Std. Err.</th>
<th>[95% Conf. Interval]</th>
</tr>
</thead>
<tbody>
<tr>
<td>id: Identity</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>var(_cons)</td>
<td>14.81751</td>
<td>3.124226</td>
<td>9.801716  22.40002</td>
</tr>
<tr>
<td>var(Residual)</td>
<td>4.383264</td>
<td>0.3163348</td>
<td>3.805112  5.04926</td>
</tr>
</tbody>
</table>

LR test vs. linear model: chibar2(01) = 472.65  Prob >= chibar2 = 0.0000

We explain the output in detail in example 1 of [ME] mixed. Here we only highlight the most important points.

1. The first estimation table reports the fixed effects. We estimate $\beta_0 = 19.36$ and $\beta_1 = 6.21$.
2. The second estimation table shows the estimated variance components. The first section of the table is labeled id: Identity, meaning that these are random effects at the id (pig) level and that their variance–covariance matrix is a multiple of the identity matrix; that is, $\Sigma = \sigma_u^2 I$. The estimate of $\hat{\sigma}_u^2$ is 14.82 with standard error 3.12.
3. The row labeled var(Residual) displays the estimated standard deviation of the overall error term; that is, $\hat{\sigma}_e^2 = 4.38$. This is the variance of the level-one errors or the variance of the residuals.
4. An LR test comparing the model with one-level ordinary linear regression is provided and is highly significant for these data.
We can predict the random intercept $u_j$ and list the predicted random intercept for the first 10 pigs by typing

```
. predict r_int, reffects
. egen byte tag = tag(id)
. list id r_int if id<=10 & tag
```

<table>
<thead>
<tr>
<th>id</th>
<th>r_int</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>-1.683105</td>
</tr>
<tr>
<td>10</td>
<td>0.8987018</td>
</tr>
<tr>
<td>19</td>
<td>-1.952043</td>
</tr>
<tr>
<td>28</td>
<td>-1.79068</td>
</tr>
<tr>
<td>37</td>
<td>-3.189159</td>
</tr>
<tr>
<td>46</td>
<td>-3.780823</td>
</tr>
<tr>
<td>55</td>
<td>-2.382344</td>
</tr>
<tr>
<td>64</td>
<td>-1.952043</td>
</tr>
<tr>
<td>73</td>
<td>-6.739143</td>
</tr>
<tr>
<td>82</td>
<td>1.16764</td>
</tr>
</tbody>
</table>

In example 3 of [ME] mixed, we show how to fit a random-slope model for these data, and in example 1 of [ME] mixed postestimation, we show how to plot the estimated regression lines for each of the pigs.

---

#### Example 2: Split-plot design

Here we replicate the example of a split-plot design from Kuehl (2000, 477). The researchers investigate the effects of nitrogen in four different chemical forms and the effects of thatch accumulation on the quality of golf turf. The experimental plots were arranged in a randomized complete block design with two replications. After two years of nitrogen treatment, the second treatment factor, years of thatch accumulation, was added to the experiment. Each of the eight experimental plots was split into three subplots. Within each plot, the subplots were randomly assigned to accumulate thatch for a period of 2, 5, and 8 years.
Nitrogen treatment is stored in the variable `nitrogen`, and the chemicals used are urea, ammonium sulphate, isobutylidene diurea (IBDU), and sulphur-coated urea (urea SC). The length of thatch accumulation is stored in the variable `thatch`. The response is the chlorophyll content of grass clippings, recorded in mg/g and stored in the variable `chlorophyll`. The `block` variable identifies the replication group.

There are two sources of variation in this example corresponding to the whole-plot errors and the subplot errors. The subplot errors are the residual errors. The whole-plot errors represents variation in the chlorophyll content across nitrogen treatments and replications. We create the variable `wpunit` to represent the whole-plot units that correspond to the levels of the nitrogen treatment and block interaction.

```stata
. egen wpunit = group(nitrogen block)
. mixed chlorophyll ibn.nitrogen##ibn.thatch ibn.block, noomitted noconstant || wpunit:, reml
```

Performing EM optimization:
Performing gradient-based optimization:

Iteration 0:  log restricted-likelihood = -13.212401
Iteration 1:  log restricted-likelihood = -13.203149
Iteration 2:  log restricted-likelihood = -13.203125
Iteration 3:  log restricted-likelihood = -13.203125
Computing standard errors:

Mixed-effects REML regression

Number of obs = 24
Group variable: wpunit
Number of groups = 8

Obs per group:
min = 3
avg = 3.0
max = 3

Wald chi2(13) = 2438.36
Log restricted-likelihood = -13.203125
Prob > chi2 = 0.0000

| chlorophyll | Coef.  | Std. Err. | z     | P>|z|    | [95% Conf. Interval] |
|-------------|--------|-----------|-------|--------|---------------------|
| nitrogen    |        |           |       |        |                     |
| urea        | 5.24583 | .3986014  | 13.16 | 0.000  | 4.464589 6.027078   |
| ammonium s. | 5.94583 | .3986014  | 14.92 | 0.000  | 5.164589 6.727078   |
| IBDU        | 7.94583 | .3986014  | 19.93 | 0.000  | 7.164589 8.727078   |
| Urea (SC)   | 8.59583 | .3986014  | 21.56 | 0.000  | 7.814589 9.377078   |
| thatch      |        |           |       |        |                     |
| 2           | -1.1   | .4632314  | -2.37 | 0.018  | -2.007917-.1920828 |
| 5           | .150006 | .4632314  | 0.32  | 0.746  | -.7579163 1.057917  |
| nitrogen#   |        |           |       |        |                     |
| thatch      |        |           |       |        |                     |
| urea#2     | -.1500005 | .6551081 | -0.23 | 0.819  | -1.433989 1.133988  |
| urea#5     | .0999994 | .6551081  | 0.15  | 0.879  | -1.183989 1.383988  |
| ammonium s. | .8999996 | .6551081  | 1.37  | 0.169  | -.383987 2.183988   |
| 2           | -.1000006 | .6551081 | -0.15 | 0.879  | -1.383989 1.183988  |
| 5           | -.2000005 | .6551081 | -.31  | 0.760  | -1.483989 1.083988  |
| IBDU#2     | -1.950001 | .6551081 | -2.98 | 0.003  | -3.233989 -.6660124 |
| IBDU#5     |        |           |       |        |                     |
| block       |        |           |       |        |                     |
| 1           | -.2916666 | .2643563 | -1.10 | 0.270  | -.8097955 .2264622  |

Random-effects Parameters

<table>
<thead>
<tr>
<th>wpunit: Identity</th>
<th>Estimate</th>
<th>Std. Err.</th>
<th>[95% Conf. Interval]</th>
</tr>
</thead>
<tbody>
<tr>
<td>var(_cons)</td>
<td>.0682407</td>
<td>.1195933</td>
<td>.0021994 2.117344</td>
</tr>
<tr>
<td>var(Residual)</td>
<td>.2145833</td>
<td>.1072917</td>
<td>.080537 .5717376</td>
</tr>
</tbody>
</table>

LR test vs. linear model: chibar2(01) = 0.53
Prob >= chibar2 = 0.2324
We can calculate the cell means for source of nitrogen and years of thatch accumulation by using `margins`.

```
. margins thatch#nitrogen
Predictive margins
Expression : Linear prediction, fixed portion, predict()
Number of obs = 24

| thatch#nitrogen | Margin | Std. Err. | z     | P>|z| | [95% Conf. Interval] |
|-----------------|--------|-----------|-------|------|----------------------|
| 2#urea          | 3.85   | .3760479  | 10.24 | 0.000 | 3.11296 | 4.58704 |
| 2#ammonium s.. | 5.6    | .3760479  | 14.89 | 0.000 | 4.86296 | 6.33704 |
| 2#IBDU          | 6.5    | .3760479  | 17.29 | 0.000 | 5.76296 | 7.23704 |
| 2#Urea (SC)     | 7.35   | .3760479  | 19.55 | 0.000 | 6.61296 | 8.08704 |
| 5#urea          | 5.35   | .3760479  | 14.23 | 0.000 | 4.61296 | 6.08704 |
| 5#ammonium s.. | 5.85   | .3760479  | 15.56 | 0.000 | 5.11296 | 6.58704 |
| 5#IBDU          | 6      | .3760479  | 15.96 | 0.000 | 5.26296 | 6.73704 |
| 5#Urea (SC)     | 8.6    | .3760479  | 22.87 | 0.000 | 7.86296 | 9.33704 |
| 8#urea          | 5.1    | .3760479  | 13.56 | 0.000 | 4.36296 | 5.83704 |
| 8#ammonium s.. | 5.8    | .3760479  | 15.42 | 0.000 | 5.06296 | 6.53704 |
| 8#IBDU          | 7.8    | .3760479  | 20.74 | 0.000 | 7.06296 | 8.53704 |
| 8#Urea (SC)     | 8.45   | .3760479  | 22.47 | 0.000 | 7.712959| 9.18704 |
```

It is easier to see the effect of the treatments if we plot the impact of the four nitrogen and the three thatch treatments. We can use `marginsplot` to plot the means of chlorophyll content versus years of thatch accumulation by nitrogen source.
We can see an increase in the mean chlorophyll content over the years of thatch accumulation for all but one nitrogen source.

The marginal means can be obtained by using `margins` on one variable at a time.

```
. margins thatch
Predictive margins       Number of obs  =       24
Expression : Linear prediction, fixed portion, predict()

<table>
<thead>
<tr>
<th></th>
<th>Delta-method</th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Margins</td>
<td>Std. Err.</td>
<td>z</td>
<td>P&gt;</td>
<td>z</td>
<td></td>
</tr>
<tr>
<td>thatch</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>2</td>
<td>5.825</td>
<td>.188024</td>
<td>30.98</td>
<td>0.000</td>
<td>5.45648</td>
<td>6.19352</td>
</tr>
<tr>
<td>5</td>
<td>6.45</td>
<td>.188024</td>
<td>34.30</td>
<td>0.000</td>
<td>6.08148</td>
<td>6.81852</td>
</tr>
<tr>
<td>8</td>
<td>6.7875</td>
<td>.188024</td>
<td>36.10</td>
<td>0.000</td>
<td>6.41898</td>
<td>7.15602</td>
</tr>
</tbody>
</table>
```
. margins nitrogen
Predictive margins Number of obs = 24
Expression : Linear prediction, fixed portion, predict()

|                | Delta-method        | Margin | Std. Err. | z    | P>|z| | [95% Conf. Interval] |
|----------------|---------------------|--------|-----------|------|------|----------------------|
| nitrogen       |                     | 4.766667 | .2643563 | 18.03 | 0.000 | 4.248538 - 5.284796 |
| urea           |                     | 5.75    | .2643563 | 21.75 | 0.000 | 5.231871 - 6.268129 |
| ammonium s.    |                     | 6.766667 | .2643563 | 25.60 | 0.000 | 6.248538 - 7.284796 |
| IBDU           |                     | 8.133333 | .2643563 | 30.77 | 0.000 | 7.615205 - 8.651462 |

Marchenko (2006) shows more examples of fitting other experimental designs using linear mixed-effects models.

Example 3: Binomial counts

We use the data taken from Agresti (2013, 219) on graduate school applications to the 23 departments within the College of Liberal Arts and Sciences at the University of Florida during the 1997–1998 academic year. The dataset contains the department ID (department), the number of applications (napplied), and the number of students admitted (nadmitted) cross-classified by gender (female).

. use https://www.stata-press.com/data/r16/admissions, clear
(Graduate school admissions data)
. describe
Contains data from https://www.stata-press.com/data/r16/admissions.dta
obs: 46 Graduate school admissions data
vars: 4 25 Feb 2018 09:28
(_dta has notes)

<table>
<thead>
<tr>
<th>variable name</th>
<th>storage</th>
<th>display</th>
<th>value</th>
<th>label</th>
</tr>
</thead>
<tbody>
<tr>
<td>department</td>
<td>long</td>
<td>%8.0g</td>
<td>dept</td>
<td>department id</td>
</tr>
<tr>
<td>nadmitted</td>
<td>byte</td>
<td>%8.0g</td>
<td>number of admissions</td>
<td></td>
</tr>
<tr>
<td>napplied</td>
<td>float</td>
<td>%9.0g</td>
<td>number of applications</td>
<td></td>
</tr>
<tr>
<td>female</td>
<td>byte</td>
<td>%8.0g</td>
<td>=1 if female, =0 if male</td>
<td></td>
</tr>
</tbody>
</table>

We wish to investigate whether admission decisions are independent of gender. Given department and gender, the probability of admission follows a binomial model, that is, \( \Pr(Y_{ij} = y_{ij}) = \text{Binomial}(n_{ij}, \pi_{ij}) \), where \( i = \{0, 1\} \) and \( j = 1, \ldots, 23 \). We fit a mixed-effects binomial logistic model with a random intercept at the department level.
The odds of being admitted are higher for females than males but without statistical significance. The estimate of $\hat{\sigma}_u^2$ is 1.35 with the standard error of 0.46. An LR test comparing the model with the one-level binomial regression model favors the random-intercept model, indicating that there is a significant variation in the number of admissions between departments.

We can further assess the model fit by performing a residual analysis. For example, here we predict and plot Anscombe residuals.
Anscombe residuals are constructed to be approximately normally distributed, thus residuals that are above two in absolute value are usually considered outliers. In the graph above, the residual for female admissions in department 2 is a clear outlier, suggesting a poor fit for that particular observation; see [ME] `meglm postestimation` for more information about Anscombe residuals and other model diagnostics tools.

Covariance structures

Example 4: Growth-curve model with correlated random effects

Here we extend the model from example 1 of [ME] `me` to allow for a random slope on `week` and an unstructured covariance structure between the random intercept and the random slope on `week`.
use https://www.stata-press.com/data/r16/pig, clear
(Longitudinal analysis of pig weights)
mixed weight week || id: week, covariance(unstructured)
Performing EM optimization:
Performing gradient-based optimization:
Iteration 0:  log likelihood = -868.96185
Iteration 1:  log likelihood = -868.96185
Computing standard errors:
Mixed-effects ML regression  Number of obs = 432
Group variable: id  Number of groups = 48
Obs per group:
  min =  9
  avg =  9.0
  max =  9
Wald chi2(1) = 4649.17
Log likelihood = -868.96185  Prob > chi2 = 0.0000

| weight  | Coef. | Std. Err. | z    | P>|z| | [95% Conf. Interval] |
|---------|-------|-----------|------|------|----------------------|
| week    | 6.209896 | .0910745  | 68.18 | 0.000 | 6.031393  6.388399 |
| _cons   | 19.35561 | .3996387  | 48.43 | 0.000 | 18.57234  20.13889 |

Random-effects Parameters

<table>
<thead>
<tr>
<th>id: Unstructured</th>
<th>Estimate</th>
<th>Std. Err.</th>
<th>[95% Conf. Interval]</th>
</tr>
</thead>
<tbody>
<tr>
<td>var(week)</td>
<td>.3715251</td>
<td>.0812958</td>
<td>.2419532  .570486</td>
</tr>
<tr>
<td>var(_cons)</td>
<td>6.823363</td>
<td>1.566194</td>
<td>4.351297  10.69986</td>
</tr>
<tr>
<td>cov(week,_cons)</td>
<td>-.0984378</td>
<td>.2545767</td>
<td>-.5973991 .4005234</td>
</tr>
<tr>
<td>var(Residual)</td>
<td>1.596829</td>
<td>.123198</td>
<td>1.372735  1.857505</td>
</tr>
</tbody>
</table>

LR test vs. linear model: chi2(3) = 764.58  Prob > chi2 = 0.0000
Note: LR test is conservative and provided only for reference.

The unstructured covariance structure allows for correlation between the random effects. Other covariance structures supported by mixed, besides the default independent, include identity and exchangeable; see [ME] mixed for details. You can also specify multiple random-effects equations at the same level, in which case the covariance types can be combined to form more complex blocked-diagonal covariance structures; see example 5 below.

We can predict the fitted values and plot the estimated regression line for each of the pigs. The fitted values are based on both the fixed and the random effects.
Example 5: Blocked-diagonal covariance structures

In this example, we fit a logistic mixed-effects model with a blocked-diagonal covariance structure of random effects.

We use the data from the 1989 Bangladesh fertility survey (Huq and Cleland 1990), which polled 1,934 Bangladeshi women on their use of contraception. The women sampled were from 60 districts, identified by the variable district. Each district contained either urban or rural areas (variable urban) or both. The variable c_use is the binary response, with a value of 1 indicating contraceptive use. Other covariates include mean-centered age and three indicator variables recording number of children. Below we fit a standard logistic regression model amended to have random coefficients on each indicator variable for children and an overall district random intercept.

. use https://www.stata-press.com/data/r16/bangladesh, clear
(Bangladesh Fertility Survey, 1989)
. melogit c_use urban age child\* || district: child\*, cov(exchangeable)  
> || district:, or

Fitting fixed-effects model:
Iteration 0:  log likelihood = -1229.5485
Iteration 1:  log likelihood = -1228.5268
Iteration 2:  log likelihood = -1228.5263
Iteration 3:  log likelihood = -1228.5263
Refining starting values:
Grid node 0:  log likelihood = -1234.3979
Fitting full model:
Iteration 0:  log likelihood = -1234.3979 (not concave)
Iteration 1:  log likelihood = -1208.0052
Iteration 2:  log likelihood = -1206.4497
Iteration 3:  log likelihood = -1206.2417
Iteration 4:  log likelihood = -1206.2397
Iteration 5:  log likelihood = -1206.2397

Mixed-effects logistic regression  
Number of obs = 1,934
Group variable: district  
Number of groups = 60

Obs per group:
min = 2
avg = 32.2
max = 118

Integration method: mvaghermite  
Integration pts. = 7

Log likelihood = -1206.2397  
Prob > chi2 = 0.0000

( 1) \[/\]var(child1[district]) - \[/\]var(child3[district]) = 0
( 2) \[/\]cov(child1[district],child2[district]) - \[/\]cov(child2[district],child3[district]) = 0
( 3) \[/\]cov(child1[district],child3[district]) - \[/\]cov(child2[district],child3[district]) = 0
( 4) \[/\]var(child2[district]) - \[/\]var(child3[district]) = 0

| c_use     | Odds Ratio | Std. Err. | z    | P>|z|  | [95% Conf. Interval] |
|-----------|------------|-----------|------|------|---------------------|
| urban     | 2.105163   | .254604   | 6.15 | 0.000| 1.660796 2.668426   |
| age       | .9735765   | .007746   | -3.37| 0.001| .9585122 .9888775  |
| child1    | 2.992596   | .502149   | 6.53 | 0.000| 2.153867 4.157931  |
| child2    | 3.879345   | .709412   | 7.41 | 0.000| 2.710815 5.51584   |
| child3    | 3.774627   | .705582   | 7.11 | 0.000| 2.616744 4.444863  |
| _cons     | .1859471   | .027481   | -11.38| 0.000| .1391841 2.484214  |

| district  |          |           |      |      |                     |
| var(child1)|  .0841518| .0880698  | .0108201| .654479 |
| var(child2)|  .0841518| .0880698  | .0108201| .654479 |
| var(child3)|  .0841518| .0880698  | .0108201| .654479 |
| var(_cons)|  .1870273| .0787274  | .0819596| .426786 |

district  
cov(child1, child2) |  .0616875| .0844681  | .073  | 0.465 | -.1038669 .2272419 |
cov(child1, child3) |  .0616875| .0844681  | .073  | 0.465 | -.1038669 .2272419 |
cov(child2, child3) |  .0616875| .0844681  | .073  | 0.465 | -.1038669 .2272419 |

Note: Estimates are transformed only in the first equation.
Note: _cons estimates baseline odds (conditional on zero random effects).
LR test vs. logistic model: \chi2(3) = 44.57  
Prob > \chi2 = 0.0000
Note: LR test is conservative and provided only for reference.
The fixed effects can be interpreted just as you would the output from `logit`. Urban women have roughly double the odds of using contraception as compared with their rural counterparts. Having any number of children will increase the odds from three- to fourfold when compared with the base category of no children. Contraceptive use also decreases with age.

Because we specified `cov(exchangeable)`, the estimated variances on each indicator variable for children are constrained to be the same, and the estimated covariances on each indicator variable for children are constrained to be the same. More complex covariance structures with constraints can be specified using `covariance(pattern())` and `covariance(fixed())`; see example 6 below.

### Example 6: Meta analysis

In this example, we present a mixed-effects model for meta analysis of clinical trials. The term “meta-analysis” refers to a statistical analysis that involves summary data from similar but independent studies.

Turner et al. (2000) performed a study of nine clinical trials examining the effect of taking diuretics during pregnancy on the risk of pre-eclampsia. The summary data consist of the log odds-ratio (variable `lnor`) estimated from each study, and the corresponding estimated variance (variable `var`). The square root of the variance is stored in the variable `std` and the trial identifier is stored in the variable `trial`.

```
use https://www.stata-press.com/data/r16/diuretics
(Meta analysis of clinical trials studying diuretics and pre-eclampsia)
list
```

<table>
<thead>
<tr>
<th>trial</th>
<th>lnor</th>
<th>var</th>
<th>std</th>
</tr>
</thead>
<tbody>
<tr>
<td>1.</td>
<td>1</td>
<td>.04</td>
<td>.16</td>
</tr>
<tr>
<td>2.</td>
<td>2</td>
<td>-.92</td>
<td>.12</td>
</tr>
<tr>
<td>3.</td>
<td>3</td>
<td>-.12</td>
<td>.18</td>
</tr>
<tr>
<td>4.</td>
<td>4</td>
<td>-1.47</td>
<td>.3</td>
</tr>
<tr>
<td>5.</td>
<td>5</td>
<td>-1.39</td>
<td>.11</td>
</tr>
<tr>
<td>6.</td>
<td>6</td>
<td>-.3</td>
<td>.01</td>
</tr>
<tr>
<td>7.</td>
<td>7</td>
<td>-.26</td>
<td>.12</td>
</tr>
<tr>
<td>8.</td>
<td>8</td>
<td>1.09</td>
<td>.69</td>
</tr>
<tr>
<td>9.</td>
<td>9</td>
<td>.14</td>
<td>.07</td>
</tr>
</tbody>
</table>

In a random-effects modeling of summary data, the observed log odds-ratios are treated as a continuous outcome and assumed to be normally distributed, and the true treatment effect varies randomly among the trials. The random-effects model can be written as

\[
y_i \sim N(\theta + \nu_i, \sigma_i^2)
\]

\[
\nu_i \sim N(0, \tau^2)
\]

where \(y_i\) is the observed treatment effect corresponding to the \(i\)th study, \(\theta + \nu_i\) is the true treatment effect, \(\sigma_i^2\) is the variance of the observed treatment effect, and \(\tau\) is the between-trial variance component. Our aim is to estimate \(\theta\), the global mean.

Notice that the responses \(y_i\) do not provide enough information to estimate this model, because we cannot estimate the group-level variance component from a dataset that contains one observation per group. However, we already have estimates for the \(\sigma_i^2\)'s, therefore we can constrain each \(\sigma_i\) to
be equal to its estimated value, which will allow us to estimate $\theta$ and $\tau$. We use `meglm` to estimate this model because the `mixed` command does not support constraints.

In `meglm`, one way to constrain a group of individual variances to specific values is by using the fixed covariance structure (an alternative way is to define each constraint individually with the `constraint` command and specify them in the `constraints()` option). The `covariance(fixed())` option requires a Stata matrix defining the constraints, thus we first create matrix $f$ with the values of $\sigma_i$, stored in variable `var`, on the main diagonal. We will use this matrix to constrain the variances.

```
. mkmat var, mat(f)
. matrix f = diag(f)
```

In the random-effects equation part, we need to specify nine random slopes, one for each trial. Because random-effects equations do not support factor variables (see [U] 11.4.3 Factor variables), we cannot use the `i.trial` notation. Instead, we `tabulate` the variable `trial` and use the `generate()` option to create nine dummy variables named `tr1`, `tr2`, ..., `tr9`. We can then fit the model. Because the model is computationaly demanding, we use Laplacian approximation instead of the default mean-variance adaptive quadrature; see `Computation time and the Laplacian approximation` above for details.
. qui tabulate trial, gen(tr)
. meglm lnor || _all: tr1-tr9, nocons cov(fixed(f)) intm(laplace) nocnsreport
Fitting fixed-effects model:
Iteration 0: log likelihood = -10.643432
Iteration 1: log likelihood = -10.643432
Refining starting values:
Grid node 0: log likelihood = -10.205455
Fitting full model:
Iteration 0: log likelihood = -10.205455
Iteration 1: log likelihood = -9.4851561 (backed up)
Iteration 2: log likelihood = -9.4587068
Iteration 3: log likelihood = -9.4552982
Iteration 4: log likelihood = -9.4552759
Iteration 5: log likelihood = -9.4552759
Mixed-effects GLM Number of obs = 9
Family: Gaussian
Link: identity
Group variable: _all Number of groups = 1
Obs per group:
  min = 9
  avg = 9
  max = 9
Integration method: laplace
Log likelihood = -9.4552759 Wald chi2(0) = .

|         | Coef.   | Std. Err. |    z    |   P>|z|  | [95% Conf. Interval] |
|---------|---------|-----------|---------|-------|----------------------|
| lnor    | -.5166151 | .2059448  | -2.51   | 0.012 | -.9202594 -.1129707  |
| _cons   |         |           |         |       |                      |
| _all    |         |           |         |       |                      |
| var(tr1)| .16 (constrained) |       |         |       |                      |
| var(tr2)| .12 (constrained) |       |         |       |                      |
| var(tr3)| .18 (constrained) |       |         |       |                      |
| var(tr4)| .3 (constrained)  |       |         |       |                      |
| var(tr5)| .11 (constrained) |       |         |       |                      |
| var(tr6)| .01 (constrained) |       |         |       |                      |
| var(tr7)| .12 (constrained) |       |         |       |                      |
| var(tr8)| .69 (constrained) |       |         |       |                      |
| var(tr9)| .07 (constrained) |       |         |       |                      |
| var(e.lnor) | .2377469 | .1950926  | .0476023 | 1.187413 |                      |

We estimate $\hat{\theta} = -0.52$, which agrees with the estimate reported by Turner et al. (2000).

We can fit the above model in a more efficient way. We can consider the trials as nine independent random variables, each with variance unity, and each being multiplied by a different standard error. To accomplish this, we treat trial as a random-effects level, use the standard deviations of the log odds-ratios as a random covariate at the trial level, and constrain the variance component of trial to unity.
constraint 1 _b[var(std[trial])] = 1
meglm lnor || trial: std, nocons constraints(1)

Fitting fixed-effects model:
Iteration 0: log likelihood = -10.643432
Iteration 1: log likelihood = -10.643432

Refining starting values:
Grid node 0: log likelihood = -10.205455

Fitting full model:
Iteration 0: log likelihood = -10.205455
Iteration 1: log likelihood = -9.4851164 (backed up)
Iteration 2: log likelihood = -9.45869
Iteration 3: log likelihood = -9.4552794
Iteration 4: log likelihood = -9.4552759
Iteration 5: log likelihood = -9.4552759

Mixed-effects GLM Number of obs = 9
Family: Gaussian
Link: identity
Group variable: trial Number of groups = 9
Obs per group:
  min = 1
  avg = 1.0
  max = 1

Integration method: mvaghermite Integration pts. = 7
Wald chi2(0) = .
( 1) [var(std[trial])] = 1

|      | Coef.       | Std. Err. | z    | P> |z|  [95% Conf. Interval]       |
|------|-------------|-----------|------|----|-----------------------------|
|     | _cons       |           |      |    |                             |
|     | -.5166151   | .2059448  | -2.51| .012| -.9202594 -.1129708        |
| trial|             |           |      |    |                             |
|     | var(std)    | 1         | (constrained) |
|     | .2377469   | .1950926  | .0476023 | 1.187413 |

The results are the same, but this model took a fraction of the time compared with the less efficient specification.

Three-level models

The methods we have discussed so far extend from two-level models to models with three or more levels with nested random effects. By “nested”, we mean that the random effects shared within lower-level subgroups are unique to the upper-level groups. For example, assuming that classroom effects would be nested within schools would be natural, because classrooms are unique to schools. Below we illustrate a three-level mixed-effects ordered probit model.

Example 7: Three-level ordinal response model

In this example, we fit a three-level ordered probit model. The data are from the Television, School, and Family Smoking Prevention and Cessation Project (Flay et al. 1988; Rabe-Hesketh and Skrondal 2012, chap. 11), where schools were randomly assigned into one of four groups defined
by two treatment variables. Students within each school are nested in classes, and classes are nested in schools. The dependent variable is the tobacco and health knowledge (THK) scale score collapsed into four ordered categories. We regress the outcome on the treatment variables and their interaction and control for the pretreatment score.

```
. use https://www.stata-press.com/data/r16/tvsfpors, clear
. meoprobit thk prethk cc##tv || school: || class:
Fitting fixed-effects model:
Iteration 0:  log likelihood =  -2212.775
Iteration 1:  log likelihood =  -2127.811
Iteration 2:  log likelihood =  -2127.761
Iteration 3:  log likelihood =  -2127.761
Refining starting values:
Grid node 0:  log likelihood =  -2195.642
Fitting full model:
Iteration 0:  log likelihood =  -2195.642 (not concave)
Iteration 1:  log likelihood =  -2167.958 (not concave)
Iteration 2:  log likelihood =  -2140.264 (not concave)
Iteration 3:  log likelihood =  -2128.695 (not concave)
Iteration 4:  log likelihood =  -2119.923
Iteration 5:  log likelihood =  -2117.095
Iteration 6:  log likelihood =  -2116.700
Iteration 7:  log likelihood =  -2116.698
Iteration 8:  log likelihood =  -2116.698
Mixed-effects oprobit regression  Number of obs =  1,600

<table>
<thead>
<tr>
<th>Group Variable</th>
<th>No. of Groups</th>
<th>Observations per Group</th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Minimum</td>
<td>Average</td>
<td>Maximum</td>
<td></td>
<td></td>
</tr>
<tr>
<td>school</td>
<td>18</td>
<td>57.1</td>
<td>137</td>
<td></td>
<td></td>
</tr>
<tr>
<td>class</td>
<td>11</td>
<td>11.9</td>
<td>28</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Integration method: mvaghermite  Integration pts. =  7
Wald chi2(4) =  124.20
Log likelihood =  -2116.698
Prob > chi2 =  0.0000

| thk   | Coef.  | Std. Err. | z     | P>|z| | [95% Conf. Interval] |
|-------|--------|-----------|-------|------|----------------------|
| prethk| .238841| .0231446  | 10.32 | 0.000| .1934784 -.2842036   |
| 1.cc  | .5254813| .1285816  | 4.09  | 0.000| .2734659 .7774967    |
| 1.tv  | .1455573| .1255827  | 1.16  | 0.246| -.1005803 .3916949   |
| cc#tv | - .2426203| .1811999 | -1.34| 0.181| -.5977656 .1125251   |
| /cut1 | -.074617| .1029791  | -.74  | 0.462| -.2764523 .1272184   |
| /cut2 | .6863046| .1034813  | 6.58  | 0.000| .4834849 .8891242    |
| /cut3 | 1.413686| .1064889  | 13.23 | 0.000| 1.204972 1.622401    |

```

school var(_cons) = .0186456 .0160226 .0034604 .1004695
school>class var(_cons) = .0519974 .0224014 .0223496 .1209745

LR test vs. oprobit model: chi2(2) = 22.13  Prob > chi2 = 0.0000

Note: LR test is conservative and provided only for reference.
Notes:
1. Our model now has two random-effects equations, separated by ||. The first is a random intercept (constant only) at the school level (level three), and the second is a random intercept at the class level (level two). The order in which these are specified (from left to right) is significant—meoprobit assumes that class is nested within school.
2. The information on groups is now displayed as a table, with one row for each grouping. You can suppress this table with the nogroup or the noheader option, which will also suppress the rest of the header.
3. The variance-component estimates are now organized and labeled according to level. The variance component for class is labeled school>class to emphasize that classes are nested within schools.

The above extends to models with more than two levels of nesting in the obvious manner, by adding more random-effects equations, each separated by ||. The order of nesting goes from left to right as the groups go from biggest (highest level) to smallest (lowest level).

Crossed-effects models
Not all mixed-effects models contain nested levels of random effects.

Example 8: Crossed random effects

Returning to our longitudinal analysis of pig weights, suppose that we wish to fit

$$weight_{ij} = \beta_0 + \beta_1 week_{ij} + u_i + v_j + \epsilon_{ij}$$

for the $i = 1, \ldots, 9$ weeks and $j = 1, \ldots, 48$ pigs and

$$u_i \sim N(0, \sigma_u^2); \quad v_j \sim N(0, \sigma_v^2); \quad \epsilon_{ij} \sim N(0, \sigma_\epsilon^2)$$

all independently. That is, we assume an overall population-average growth curve $\beta_0 + \beta_1 week$ and a random pig-specific shift. In other words, the effect due to $week$, $u_i$, is systematic to that week and common to all pigs. The rationale behind (8) could be that, assuming that the pigs were measured contemporaneously, we might be concerned that week-specific random factors such as weather and feeding patterns had significant systematic effects on all pigs.

Model (8) is an example of a two-way crossed-effects model, with the pig effects $v_j$ being crossed with the week effects $u_i$. One way to fit such models is to consider all the data as one big cluster, and treat $u_i$ and $v_j$ as a series of $9 + 48 = 57$ random coefficients on indicator variables for week and pig. The random effects $u$ and the variance components $G$ are now represented as
\[ \mathbf{u} = \begin{bmatrix} u_1 \\ \vdots \\ u_9 \\ v_1 \\ \vdots \\ v_{48} \end{bmatrix} \sim N(0, \mathbf{G}); \quad \mathbf{G} = \begin{bmatrix} \sigma_u^2 \mathbf{I}_9 & 0 \\ 0 & \sigma_v^2 \mathbf{I}_{48} \end{bmatrix} \]

Because \( \mathbf{G} \) is block diagonal, it can be represented as repeated-level equations. All we need is an ID variable to identify all the observations as one big group and a way to tell mixed-effects commands to treat \texttt{week} and \texttt{pig} as factor variables (or equivalently, as two sets of overparameterized indicator variables identifying weeks and pigs, respectively). The mixed-effects commands support the special group designation \texttt{all} for the former and the \texttt{R.varname} notation for the latter.

```
. use https://www.stata-press.com/data/r16/pig
(Longitudinal analysis of pig weights)
. mixed weight week || _all: R.id || _all: R.week
Performing EM optimization:
Performing gradient-based optimization:
Iteration 0:  log likelihood =  -1013.824
Iteration 1:  log likelihood =  -1013.824
```

```
Computing standard errors:
 Mixed-effects ML regression Number of obs =  432
 Group variable:  _all Number of groups =  1
Obs per group:
    min = 432
    avg = 432.0
    max = 432
Wald chi2(1)  = 13258.28
Log likelihood =  -1013.824  Prob > chi2 =  0.0000

weight  Coef.  Std. Err.     z  P>|z|     [95% Conf. Interval]
    week  6.209896  .0539313 115.14  0.000     6.104192    6.315599
   _cons 19.35561  .6333982  30.56  0.000    18.11418    20.59705

Random-effects Parameters  |   Estimate  Std. Err.     [95% Conf. Interval]
  _all: Identity      var(R.id)  =  14.83623  3.126142     9.816733    22.42231
  _all: Identity      var(R.week) =  .0849874  .0868856   .0114588    .6303302
  var(Residual)       =  4.297328  .3134404     3.724888    4.957741
LR test vs. linear model: chi2(2)  = 474.85  Prob > chi2 =  0.0000
Note: LR test is conservative and provided only for reference.
```

We estimate \( \hat{\sigma}_u^2 = 0.08 \) and \( \hat{\sigma}_v^2 = 14.84 \).

The \texttt{R.varname} notation is equivalent to giving a list of overparameterized (none dropped) indicator variables for use in a random-effects specification. When you use \texttt{R.varname}, mixed-effects commands handle the calculations internally rather than creating the indicators in the data. Because the set of indicators is overparameterized, \texttt{R.varname} implies \texttt{noconstant}.
Note that the column dimension of our random-effects design is 57. Computation time and memory requirements grow (roughly) quadratically with the dimension of the random effects. As a result, fitting such crossed-effects models is feasible only when the total column dimension is small to moderate. For this reason, mixed-effects commands use the Laplacian approximation as the default estimation method for crossed-effects models; see *Computation time and the Laplacian approximation* above for more details.

It is often possible to rewrite a mixed-effects model in a way that is more computationally efficient. For example, we can treat pigs as nested within the `_all` group, yielding the equivalent and more efficient (total column dimension 10) way to fit (8):

```
mixed weight week || _all: R.week || id:
```

The results of both estimations are identical, but the latter specification, organized at the cluster (pig) level with random-effects dimension 1 (a random intercept) is much more computationally efficient. Whereas with the first form we are limited in how many pigs we can analyze, there is no such limitation with the second form.

All the mixed-effects commands—except `mixed`—automatically attempt to recast the less efficient model specification into a more efficient one. However, this automatic conversion may not be sufficient for some complicated mixed-effects specifications, especially if both crossed and nested effects are involved. Therefore, we strongly encourage you to always specify the more efficient syntax; see Rabe-Hesketh and Skrondal (2012) and Marchenko (2006) for additional techniques to make calculations more efficient in more complex mixed-effects models.

### Nonlinear models

NLME models are popular in population pharmacokinetics, bioassays, studies of biological and agricultural growth processes, and other applications, where the mean function is a nonlinear function of fixed and random effects. *Remarks and examples* of [ME] `mnen` provide many examples of fitting different NLME models by using `mnen`, including a pharmacokinetics model in example 15. Here we consider simple data from Draper and Smith (1998) that contain trunk circumference (in mm) of five different orange trees measured over seven different time points.
Let’s plot our data first.

```
. use https://www.stata-press.com/data/r16/orange
   (Growth of orange trees (Draper and Smith, 1998))
. twoway scatter circumf age, connect(L) ylabel(#6 175)
```

Consider the following nonlinear growth model for these data,

\[
\text{circumf}_{ij} = \frac{\beta_1}{1 + \exp\left\{-\left(\text{age}_{ij} - \beta_2\right)/\beta_3\right\}} + \epsilon_{ij}
\]

where \(\epsilon_{ij}\)’s are i.i.d. N(0, \(\sigma^2_\epsilon\)). In this model, \(\beta_1\) can be interpreted as the average asymptotic trunk circumference of trees as \(\text{age}_{ij} \to \infty\). We can crudely estimate it as the average of the trunk circumference values at the last observed time point, which for these data is roughly 175 mm. \(\beta_2\) is the age at which a tree attains half of the average asymptotic trunk circumference \(\beta_1\); that is, if we set \(\text{age}_{ij} = \beta_2\), then \(E(\text{circumf}_{ij}) = 0.5\beta_1\). \(\beta_3\) is a scale parameter that represents the number of days it takes for a tree to grow from 50% to about 73% of the average asymptotic trunk circumference. That is, if we set \(\text{age} = t_{0.73} = \beta_2 + \beta_3\), then \(E(\text{circumf}_{ij}) = \beta_1/[1 + \exp(-1)] = 0.73\beta_1\) and then \(\beta_3 = t_{0.73} - \beta_2\).

The above model can be easily fit by using, for example, `nl`; see [R] nl. However, if we study the graph more carefully, we will notice that there is an increasing variability in the trunk circumferences of trees as they approach their limiting age. So it may be more reasonable to allow \(\beta_1\) to vary between trees,

\[
\text{circumf}_{ij} = \frac{\beta_1 + u_{1ij}}{1 + \exp\left\{-\left(\text{age}_{ij} - \beta_2\right)/\beta_3\right\}} + \epsilon_{ij}
\]

where \(u_{1ij}\)’s are i.i.d. N(0, \(\sigma^2_{u_1}\)). We use `menl` to fit this model.

The specification of NLME models in `menl` is fairly straightforward. Following the dependent variable and the equality sign (=), we specify the expression for the mean function as a usual Stata expression but with parameters and random effects enclosed in curly braces ({}).
. menl circumf = ({b1}+{U1[tree]})/(1+exp(-(age-{b2})/{b3}))

Obtaining starting values by EM:

Alternating PNLS/LME algorithm:

Iteration 1:  linearization log likelihood = -131.584579

Computing standard errors:

Mixed-effects ML nonlinear regression

| Coef. | Std. Err. | z   | P>|z|  | [95% Conf. Interval] |
|-------|-----------|-----|------|------------------------|
| /b1   | 191.049   | 16.15403 | 11.83 | 0.000 | 159.3877 222.7103 |
| /b2   | 722.556   | 35.15082 | 20.56 | 0.000 | 653.6616 791.4503 |
| /b3   | 344.1624  | 27.14739 | 12.68 | 0.000 | 290.9545 397.3703 |

Random-effects Parameters

<table>
<thead>
<tr>
<th>Coef.</th>
<th>Std. Err.</th>
<th>[95% Conf. Interval]</th>
</tr>
</thead>
<tbody>
<tr>
<td>var(U1)</td>
<td>991.1514</td>
<td>639.4636 3510.038</td>
</tr>
<tr>
<td>var(Residual)</td>
<td>61.56371</td>
<td>15.89568 102.1184</td>
</tr>
</tbody>
</table>

In the above specification, we used {U1[tree]} to include random intercepts at the tree level in our model. U1 is the name or label associated with these random intercepts.

The output of menl is similar to that of mixed—the header information is displayed first, fixed-effects parameter estimates are displayed in the first or the fixed-effects parameter table, and the estimates of variance components are displayed in the second or the random-effects parameter table. The header information is similar to that of mixed, but unlike mixed, menl in general does not report a model $\chi^2$ statistic in the header because a test of the joint significance of all fixed-effects parameters (except the constant term) may not be relevant in a nonlinear model. menl also reports the so-called linearization log likelihood. menl uses the linearization method of Lindstrom and Bates (1990), with extensions from Pinheiro and Bates (1995), for estimation. This method is based on the approximation of the NLME model by an LME model, in which a first-order Taylor-series approximation is used to linearize the nonlinear mean function with respect to fixed and random effects; see *Introduction* and *Methods and formulas* in [ME] menl for details. The linearization log likelihood is the log likelihood of this approximating LME model. We can use this log likelihood for model comparison of different NLME models and to form likelihood-ratio tests, but note that this is not the log likelihood of the corresponding NLME model. Depending on the accuracy of the approximation, the linearization log likelihood may be close to the true NLME log likelihood.

As part of Stata’s standard estimation output, menl reports $z$ tests against zeros for the estimated fixed-effects parameters. Testing a parameter against zero may or may not be of interest, or may not even be appropriate, in a nonlinear model. In our example, {b3} is the denominator of a fraction, so the test of {b3} against zero may not be feasible in this model. Instead, we may be interested in testing {b3} against, for example, 300, which would correspond to testing whether the average trunk circumference of orange trees increases from 50% to 73% of its asymptotic value in 300 days. We can perform this test by using, for instance, the test command; see [R] test. As a side note, setting $\beta_3 = 0$ in (9) results in a simple random-intercept model, in a limiting sense.
From the random-effects table, the variability in limiting growth $\beta_1$ between trees, labeled as $\text{var}(U_1)$, is statistically significant in this model with an estimate of 991 (mm$^2$) and a 95% CI of [280, 3510].

We can rewrite (9) as a two-stage model,

$$\text{circumf}_{ij} = \frac{\phi_{1j}}{1 + \exp\left\{-\left(\text{age}_{ij} - \phi_{2j}\right)/\phi_{3j}\right\}} + \epsilon_{ij}$$

(10)

where the stage 2 specification is

$$\phi_j = \begin{bmatrix} \phi_{1j} \\ \phi_{2j} \\ \phi_{3j} \end{bmatrix} = \begin{bmatrix} \beta_1 + u_{1j} \\ \beta_2 \\ \beta_3 \end{bmatrix}$$

(11)

The model defined by (10) and (11) is the same as that defined by (9) but with a different parameterization.

In menl, we can accommodate this two-stage formulation with the define() option. For example, we can fit the two-stage model defined by (10) and (11) as follows:

```
. menl circumf = {phi1:}/(1+exp(-(age-{b2})/{b3})), define(phi1: {b1}+{U1[tree]})
```

Obtaining starting values by EM:
Alternating PNLS/LME algorithm:
Iteration 1: linearization log likelihood = -131.584579
Computing standard errors:
Mixed-effects ML nonlinear regression
Number of obs = 35
Group variable: tree
Number of groups = 5
Obs per group:
  min = 7
  avg = 7.0
  max = 7

Linearization log likelihood = -131.58458

```
phi1: {b1}+{U1[tree]})
```

| circumf | Coef. | Std. Err. | z     | P>|z| | [95% Conf. Interval] |
|---------|-------|-----------|-------|------|---------------------|
| /b1     | 191.049 | 16.15403  | 11.83 | 0.000 | 159.3877  222.7103 |
| /b2     | 722.556 | 35.15082  | 20.56 | 0.000 | 653.6616  791.4503 |
| /b3     | 344.1624 | 27.14739 | 12.68 | 0.000 | 290.9545  397.3703 |

The results are identical to the previous model. Here we defined a substitutable expression phi1 in the define() option as a function of {b1} and {U1[tree]} and included it in our main expression as {phi1:}. Including a colon (?) in {phi1:} is important to notify menl that it is a substitutable expression rather than a simple scalar parameter {phi1}.

In general, we can accommodate multistage formulations by using the define() option repeatedly.
More conveniently, we can use a linear-combination specification (see *Linear combinations in [ME] menl*) within the `define()` option to define the linear combination \{b_1\}+{U_1[tree]}.

```
. menl circumf = {phi1:}/(1+exp(-\(\text{age}-{b_2})/{b_3}\)), define(phi1: U1[tree], xb)
```

Obtaining starting values by EM:

Alternating PNLS/LME algorithm:

Iteration 1:  linearization log likelihood = -131.584579

Computing standard errors:

Mixed-effects ML nonlinear regression

<table>
<thead>
<tr>
<th>Group variable: tree</th>
</tr>
</thead>
</table>

| Number of obs = 35 |
| Number of groups = 5 |
| Obs per group: |
| min = 7 |
| avg = 7.0 |
| max = 7 |

Linearization log likelihood = -131.58458

```
phi1:  U1[tree], xb
```

```

| circumf | Coef. | Std. Err. | z  | P>|z| | [95% Conf. Interval] |
|---------|-------|-----------|----|-----|----------------------|
| phi1_cons | 191.049 | 16.15403 | 11.83 | 0.000 | 159.3877 | 222.7103 |
| /b2 | 722.566 | 35.15082 | 20.56 | 0.000 | 653.6616 | 791.4503 |
| /b3 | 344.1624 | 27.14739 | 12.68 | 0.000 | 290.9545 | 397.3703 |

<table>
<thead>
<tr>
<th>Random-effects Parameters</th>
<th>Estimate</th>
<th>Std. Err.</th>
<th>[95% Conf. Interval]</th>
</tr>
</thead>
<tbody>
<tr>
<td>tree: Identity</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>var(U1)</td>
<td>991.1514</td>
<td>639.4636</td>
<td>279.8776</td>
</tr>
<tr>
<td>var(Residual)</td>
<td>61.56371</td>
<td>15.89568</td>
<td>37.11466</td>
</tr>
</tbody>
</table>
```

The \{phi1: U1[tree], xb\} specification used in the `define()` option, but without curly braces, creates a linear combination named `phi1` that contains a constant \{phi1: _cons\} and random intercepts \{U1\} at the `tree` level. In the linear-combination specification, the constant is included automatically unless you specify the `noconstant` option such as `phi1: U1[tree], xb noconstant`. Also, you do not specify curly braces around random effects within the linear-combination specification. If we had covariates, say, x1 and x2, that we also wanted to include in the linear combination, we would have used `phi1: x1 x2 U1[tree]`. Notice that we did not specify the `xb` option in the previous linear combination. When a linear combination contains more than one term, this option is implied. When a linear combination contains only one term, such as in \{phi1: U1[tree], xb\}, the `xb` option must be specified to request that `menl` treat the specification as a linear combination instead of a scalar parameter; see *Random-effects substitutable expressions in [ME] menl* for details.

Instead of using `define()`, we could have similarly specified the linear combination directly in the main expression:

```
. menl circumf = {phi1: U1[tree], xb}/(1+exp(-\(\text{age}-{b_2})/{b_3}\))
```

(output omitted)

However, by using the `define()` option, we simplified the look of the main equation.
We can extend the stage 2 specification (11) to allow, for example, $\beta_2$ to vary across trees by including random intercepts at the tree level for $\phi_{2j}$,

$$
\phi_j = \begin{bmatrix}
\phi_{1j} \\
\phi_{2j} \\
\phi_{3j}
\end{bmatrix} = \begin{bmatrix}
\beta_1 + u_{1j} \\
\beta_2 + u_{2j} \\
\beta_3
\end{bmatrix}
$$

We can then fit the corresponding model by using `menl` as follows:

```
.menl circumf = {phi1:}/(1+exp(-(age-{phi2:})/{b3})),
> define(phi1: U1[tree], xb) define(phi2: U2[tree], xb)
```

Obtaining starting values by EM:

Alternating PNLS/LME algorithm:

Iteration 1:  linearization log likelihood = -131.605387
Iteration 2:  linearization log likelihood = -131.5827
Iteration 3:  linearization log likelihood = -131.580503
Iteration 4:  linearization log likelihood = -131.580265
Iteration 5:  linearization log likelihood = -131.580257

Computing standard errors:

Mixed-effects ML nonlinear regression  Number of obs = 35
Group variable: tree  Number of groups = 5
Obs per group:
  min = 7
  avg = 7.0
  max = 7

Linearization log likelihood = -131.58026

|     | Coef.   | Std. Err. |     z  | P>|z|  |  [95% Conf. Interval] |
|-----|---------|-----------|-------|-----|----------------------|
| phi1 | _cons   | 190.5939  | 16.211| 11.76| 0.000  | 158.8209  | 222.3669 |
|     | _cons   | 719.6027  | 35.77597| 20.11| 0.000  | 649.4831  | 789.7223 |
|     | /b3     | 342.0794  | 26.42036| 12.95| 0.000  | 290.2965  | 393.8624 |

Random-effects Parameters  Estimate  Std. Err.  [95% Conf. Interval]

<table>
<thead>
<tr>
<th>tree: Independent</th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>var(U1)</td>
<td>1012.15</td>
<td>666.2807</td>
<td>278.557</td>
</tr>
<tr>
<td>var(U2)</td>
<td>503.2308</td>
<td>2401.322</td>
<td>.043651</td>
</tr>
<tr>
<td>var(Residual)</td>
<td>59.27073</td>
<td>18.21297</td>
<td>32.45483</td>
</tr>
</tbody>
</table>

The large standard error for the estimate of the variance component var(U2) suggests that our model is overparameterized—a common problem when fitting NLME models. We could verify this, for instance, by computing information criteria ([R] estimates stats) or by performing a likelihood-ratio test ([R] lrtest).

By default, `menl` assumes an independent covariance structure for the random effects such as U1 and U2 in our example. We can specify, for example, an unstructured model by using the `covariance()` option. We demonstrate this only for illustration, given that our simpler model that assumed independence between U1 and U2 was already overparameterized.
Introduction to multilevel mixed-effects models

. menl circumf = {phi1:}/(1+exp(-(age-{phi2:})/{b3})),
> define(phi1: U1[tree], xb) define(phi2: U2[tree], xb)
> covariance(U1 U2, unstructured)

Obtaining starting values by EM:
Alternating PNLS/LME algorithm:
Iteration 1:  linearization log likelihood = -130.90452
Iteration 2:  linearization log likelihood = -130.902051
Iteration 3:  linearization log likelihood = -130.901772
Iteration 4:  linearization log likelihood = -130.901766

Computing standard errors:
Mixed-effects ML nonlinear regression
Number of obs = 35
Group variable: tree  Number of groups = 5
Obs per group:
  min = 7
  avg = 7.0
  max = 7

Linearization log likelihood = -130.90177
phi1:  U1[tree], xb
phi2:  U2[tree], xb

|         | Coef.  | Std. Err. | z     | P>|z|  | [95% Conf. Interval] |
|---------|--------|-----------|-------|------|---------------------|
| phi1    |        |           |       |      |                     |
| _cons   | 189.8349 | 17.20035  | 11.04 | 0.000 | 156.1228 223.5469   |
| phi2    |        |           |       |      |                     |
| _cons   | 709.5333 | 37.24229  | 19.05 | 0.000 | 636.5397 782.5268   |
| /b3     | 340.4731 | 25.52176  | 13.34 | 0.000 | 290.4514 390.4948   |

Random-effects Parameters

<table>
<thead>
<tr>
<th></th>
<th>Estimate</th>
<th>Std. Err.</th>
<th>[95% Conf. Interval]</th>
</tr>
</thead>
<tbody>
<tr>
<td>tree: Unstructured</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>var(U1)</td>
<td>1180.097</td>
<td>775.0818</td>
<td>325.7264 4275.458</td>
</tr>
<tr>
<td>var(U2)</td>
<td>1469.879</td>
<td>2777.131</td>
<td>36.22888 59638.94</td>
</tr>
<tr>
<td>cov(U1,U2)</td>
<td>1015.504</td>
<td>1124.568</td>
<td>-1188.608 3219.617</td>
</tr>
<tr>
<td>var(Residual)</td>
<td>56.07332</td>
<td>16.20294</td>
<td>31.82682 98.79144</td>
</tr>
</tbody>
</table>

In `menl`, we need to list the names of the random effects in the `covariance()` option for which we want to specify a covariance structure other than the independent one used by default.

In our example, parameters \( \phi_{1j} \) and \( \phi_{2j} \) were modeled as linear functions of random effects and parameters \( \beta_1 \) and \( \beta_2 \). The relationship does not have to be linear; see example 15 in [ME] menl.

This example has a small number of trees or clusters, so REML estimation would have been more appropriate. We could have obtained REML estimates in our examples by specifying the `reml` option with `menl`.

See [ME] menl for more examples of and details about the menl command.

Acknowledgments

We are indebted to Sophia Rabe-Hesketh of the University of California, Berkeley, and coauthor of the Stata Press book Multilevel and Longitudinal Modeling Using Stata; Anders Skrondal of the
University of Oslo and the Norwegian Institute of Public Health, and coauthor of the Stata Press book *Multilevel and Longitudinal Modeling Using Stata*; and Andrew Pickles of the University of Manchester for their extensive body of work in Stata, both previous and ongoing, in this area.

References


**Also see**

[ME] *Glossary*