

APPENDIX C

Meta-Analysis of Cancer Results from Epidemiological Studies

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1 analyses, an RR estimate for the highest exposure group was selected from studies that presented
2 results for different exposure groups. Exposure groups based on some measure of cumulative
3 exposure were preferred, if available; however, often duration was the sole exposure metric used.
4 Specific selection choices are described in the following subsections detailing the actual
5 analyses.

6 The meta-analysis calculations are based on (natural) logarithm-transformed values.
7 Thus, each RR estimate was transformed to its natural logarithm (referred to here as “log RR,”
8 the conventional terminology in epidemiology), and either an estimate of the standard error (SE)
9 of the log RR was obtained, from which to estimate the variance for the weights, or an estimate
10 of the variance of the log RR was calculated directly. If the reported 95% confidence interval
11 limits were proportionally symmetric about the observed RR estimate (i.e., upper confidence
12 limit/RR \approx RR/lower confidence limit), then an estimate of the SE of the log RR estimate was
13 obtained using the formula
14

$$SE = \frac{[\log(UCL) - \log(LCL)]}{3.92}, \quad (\text{Eq. C-1})$$

16
17 where UCL is the upper confidence limit and LCL is the lower confidence limit (for 90%
18 confidence intervals [CIs], the divisor is 3.29) (Rothman and Greenland, 1998). In all the TCE
19 cohort studies reporting SMRs or SIRs as the overall RR estimates, reported CIs were calculated
20 assuming the number of deaths (or cases) is approximately Poisson distributed. In such cases,
21 the CIs are not proportionally symmetric about the RR estimate (unless the number of deaths is
22 fairly large), and the SE of the log RR estimate was estimated as the inverse of the square root of
23 the observed number of deaths (or cases) (Breslow and Day, 1987). In some case-control
24 studies, no overall odds ratio (OR) was reported, so a crude OR estimate was calculated as
25 $OR = (a/b)/(c/d)$, where a, b, c, and d are the cell frequencies in a 2×2 table of cancer cases vs.
26 TCE exposure, and the variance of the log OR was estimated using the formula
27

$$Var[\log(OR)] = \frac{1}{a} + \frac{1}{b} + \frac{1}{c} + \frac{1}{d}, \quad (\text{Eq. C-2})$$

29
30 in accordance with the method proposed by Woolf (1955), as described by Breslow and Day
31 (1980).
32

1 The analyses that were performed for this assessment include

- 2
- 3 • meta-analyses to obtain overall summary estimates of RR
- 4 • heterogeneity analyses
- 5 • analyses of the influence of single studies on the summary estimates
- 6 • analyses of the sensitivity of the summary estimate to alternate study inclusion selections
- 7 or to alternate selections of RR estimates from a study
- 8 • publication bias analyses
- 9 • meta-analyses to obtain summary estimates for the highest exposure groups in studies
- 10 that provide data by exposure group, and
- 11 • consideration of some potential sources of heterogeneity across studies.
- 12

13 The analyses were conducted using Excel spreadsheets and the software package Comprehensive
14 Meta-Analysis, Version 2 (© 2006, Biostat, Inc.). Figures were generated using the
15 Comprehensive Meta-Analysis software. Note that for these figures, this software recalculates
16 CIs for the studies based on the SE inputs, and the resulting CIs are not always identical to those
17 reported in the original studies, in particular those based on Poisson distributions. However, the
18 recalculated CIs are merely outputs and are not the basis for any calculations in the software; SEs
19 were obtained as described above, and these SEs and the log RRs constitute the inputs for the
20 meta-analysis calculations.

21 The heterogeneity (or homogeneity) analysis tests the hypothesis that the study results are
22 homogeneous, i.e., that all the RR estimates are estimating the same population RR and the total
23 variance is no more than would be expected from within-study variance. Heterogeneity was
24 assessed using the statistic Q described by DerSimonian and Laird (1986). The Q -statistic
25 represents the sum of the weighted squared differences between the summary RR estimate
26 (obtained under the null hypothesis, i.e., using a fixed-effect model) and the RR estimate from
27 each study, and, under the null hypothesis, Q approximately follows a χ^2 distribution with
28 degrees of freedom equal to the number of studies minus one. However, this test can be under-
29 powered when the number of studies is small, and it is only a significance test, i.e., it is not very
30 informative about the *extent* of any heterogeneity. Therefore, the I^2 value (Higgins et al., 2003)
31 was also considered. $I^2 = 100\% \times (Q - df)/Q$, where Q is the Q -statistic and df is the degrees of
32 freedom, as described above. This value estimates the percentage of variation that is due to
33 study heterogeneity. Typically, I^2 values of 25%, 50%, and 75% are considered low, moderate,
34 and high amounts of heterogeneity, respectively.

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1 Subgroup analyses were sometimes conducted to examine whether or not the combined
2 RR estimate varied significantly between different types of studies (e.g., case-control vs. cohort
3 studies). In such subgroup analyses of categorical variables (e.g., study design), analysis of
4 variance was used to determine if there was significant heterogeneity between the subgroups.
5 Applying analysis of variance to meta-analyses with two subgroups ($df = 1$), $Q_{\text{between subgroups}} =$
6 $Q_{\text{overall}} - (Q_{\text{subgroup1}} + Q_{\text{subgroup2}}) = z\text{-value}^2$, where Q_{overall} is the Q -statistic calculated across all the
7 studies and $Q_{\text{subgroup1}}$ and $Q_{\text{subgroup2}}$ are the Q -statistics calculated within each subgroup.

8 Publication bias is a systematic error that occurs if statistically significant studies are
9 more likely to be submitted and published than nonsignificant studies. Studies are more likely to
10 be statistically significant if they have large effect sizes (in this case, RR estimates); thus, an
11 upward bias would result in a meta-analysis if the available published studies have higher effect
12 sizes than the full set of studies that was actually conducted. One feature of publication bias is
13 that smaller studies tend to have larger effect sizes than larger studies, since smaller studies need
14 larger effect sizes in order to be statistically significant. Thus, many of the techniques used to
15 analyze publication bias examine whether or not effect size is associated with study size.
16 Methods used to investigate potential publication bias for this assessment included funnel plots,
17 which plot effect size vs. study size (actually, SE vs. log RR here); the “trim and fill” procedure
18 of Duvall and Tweedie (2000), which imputes the “missing” studies in a funnel plot (i.e., the
19 studies needed to counterbalance an asymmetry in the funnel plot resulting from an ostensible
20 publication bias) and recalculates a summary effect size with these studies present; forest plots
21 (arrays of RRs and CIs by study) sorted by precision (i.e., SE) to see if effect size shifts with
22 study size; Begg and Mazumdar rank correlation test (Begg and Mazumdar, 1994), which
23 examines the correlation between effect size estimates and their variances after standardizing the
24 effect sizes to stabilize the variances; Egger’s linear regression test (Egger et al., 1997), which
25 tests the significance of the bias reflected in the intercept of a regression of effect size/SE on
26 $1/SE$; and cumulative meta-analyses after sorting by precision to assess the impact on the
27 summary effect size estimate of progressively adding the smaller studies.

28 29 **C.2. META-ANALYSIS FOR LYMPHOMA**

30 **C.2.1. Overall Effect of TCE Exposure**

31 **C.2.1.1. Selection of RR Estimates**

32 The selected RR estimates for lymphoma associated with TCE exposure from the
33 selected epidemiological studies are presented in Table C-1 for cohort studies and in Table C-2
34 for case-control studies. A few of the more recent case-control studies classified lymphomas
35 along the lines of the recent WHO/REAL classification system (World Health

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1 Organization/Revised European-American Classification of Lymphoid Neoplasms) (Harris et al.,
2 2000); however, most of the available TCE studies reported lymphoma results according to the
3 International Classification of Diseases (ICD), Revisions 7, 8, and 9, and focused on
4 non-Hodgkin lymphoma (NHL; ICD 200 + 202). For consistency of endpoint in the lymphoma
5 meta-analyses, RR estimates for ICD 200 + 202 were selected, wherever possible; otherwise,
6 estimates for the classification(s) best approximating NHL were selected. In addition, many of
7 the studies provided RR estimates only for males and females combined, and we are not aware of
8 any basis for a sex difference in the effects of TCE on lymphoma risk; thus, wherever possible,
9 RR estimates for males and females combined were used. The only study of much size (in terms
10 of number of lymphoma cancer cases) that provided results separately by sex was
11 Raaschou-Nielsen (2003). This study reports an insignificantly higher SIR for females (1.4,
12 95% CI: 0.73, 2.34) than for males (1.2, 95% CI: 0.98, 1.52).

13 Beyond selecting adjusted RR estimates for lymphoma classification and both sexes,
14 when multiple estimates were available, the preference was to select the RR estimate that
15 represented the largest population in a study, while trying to minimize the likelihood of TCE
16 exposure misclassification. Sensitivity analyses were generally done to investigate the impact of
17 these alternate selection choices, as well as to estimate the impacts of study findings that were
18 not reported.

19 Thus, for example, for Axelson et al. (1994), in which a small subcohort of females was
20 studied but only results for the larger male subcohort were reported, the reported male-only
21 results were used in the primary analysis; however, an attempt was made to estimate the female
22 contribution to an overall RR estimate for both sexes and its impact on the meta-analysis.
23 Axelson et al. (1994) reported that there were no cases of lymphoma observed in females, but the
24 expected number was not presented. To estimate the expected number, the expected number for
25 males was multiplied by the ratio of female-to-male person-years in the study and by the ratio of
26 female-to-male age-adjusted incidence rates for NHL.¹ The male results and the estimated
27 female contribution were then combined into an RR estimate for both sexes assuming a Poisson
28 distribution, and this alternate RR estimate for the Axelson et al. (1994) study was used in a
29 sensitivity analysis.

¹Person-years for men and women ≤ 79 years were obtained from Axelson et al. (1994): 23516.5 and 3691.5, respectively. Lifetime age-adjusted incidence rates for NHL for men and women were obtained from the National Cancer Institute's 2000-2004 SEER-17 (Surveillance Epidemiology and End Results from 17 geographical areas) database (<http://seer.cancer.gov/statfacts/html/nhl.html>): 23.2/100,000 and 16.3/100,000, respectively. The calculation for estimating the expected number of cases in females in the cohort assumes that the males and females have similar TCE exposures and that the relative distributions of age-related incidence risk for the males and females in the cohort are adequately represented by the ratios of person-years and lifetime incidence rates used in the calculation.

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Table C-1. Selected RR estimates for lymphoma associated with TCE exposure (overall effect) from cohort studies

Study	RR	95% LCL	95% UCL	RR type	log RR	SE(log RR)	Alternate RR estimates	Comments
Anttila et al., 1995	1.81	0.78	3.56	SIR	0.593	0.354	None	ICD-7 200 + 202.
Axelson et al., 1994	1.52	0.49	3.53	SIR	0.419	0.447	1.36 (0.44, 3.18) with estimated female contribution to SIR added (see text)	ICD-7 200 and 202. Results reported separately; combined assuming Poisson distribution. Results reported for males only, but there was a small female component to the cohort.
Boice et al., 1999	1.19	0.65	1.99	SMR	0.174	0.267	1.19 (0.83, 1.65) for any potential exposure	ICD-9 200 + 202. For potential routine exposure.
Greenland et al., 1994	0.76	0.24	2.42	OR	-0.274	0.590	None	ICD-8 200-202. Nested case-control study.
Hansen et al., 2001	3.1	1.3	6.1	SIR	1.13	0.354	None	ICD-7 200 + 202. Male and female results reported separately; combined assuming Poisson distribution.
Morgan et al., 1998	1.01	0.46	1.92	SMR	0.00995	0.333	1.36 (0.35, 5.21) unpublished RR for ICD 200 (see text)	ICD 200 + 202. Results reported by Mandel et al. (2006). ICD Revision 7, 8, or 9, depending on year of death.
Raaschou-Nielsen et al., 2003	1.24	1.01	1.52	SIR	0.215	0.104	None	ICD-7 200 + 202.
Radican et al., 2008	1.36	0.77	2.39	Mortality HR	0.307	0.289	None	ICD-8,-9 200 + 202; ICD-10 C82-C85. Time variable = age; covariates = sex and race. Referent group is workers with no chemical exposures.

Table C-1. Selected RR estimates for lymphoma associated with TCE exposure (overall effect) from cohort studies (continued)

Study	RR	95% LCL	95% UCL	RR type	log RR	SE(log RR)	Alternate RR estimates	Comments
Zhao et al., 2005	1.44	0.90	2.30	Mortality RR	0.363	0.239	Incidence RR: 0.77 (0.42, 1.39) Boice 2006 SMR for ICD-9 200 + 202: 0.21 (0.01, 1.18)	All lymphohematopoietic cancer (ICD-9 200-208), not just 200 + 202. Males only; adjusted for age, socioeconomic status (SES), time since first employment. Mortality results reflect more exposed cases (33) than do incidence results (17). Overall RR estimated by combining across exposure groups (see text). Boice 2006 cohort overlaps Zhao cohort; just 1 exposed death for ICD 200 + 202; 9 for 200-208 (vs. 33 in Zhao).

Table C-2. Selected RR estimates for lymphoma associated with TCE exposure from case-control studies^a

Study	RR	95% LCL	95% UCL	log RR	SE(log RR)	Lymphoma type	Comments
Hardell et al., 1994	7.2	1.3	42	1.97	0.887	NHL	Rappaport classification system. Males only; controls matched for age, place of residence, vital status.
Miligi et al., 2006	0.93	-- ^b	-- ^b	-0.0726	0.168	NHL + CLL	NCI working formulation. Crude OR; overall adjusted OR not presented.
Nordstrom et al., 1998	1.5	0.7	3.3	0.405	0.396	HCL	HCL specifically. Males only; controls matched for age and county; analysis controlled for age.
Persson and Frederikson, 1999	1.2	0.5	2.4	0.182	0.400	NHL	Classification system not specified. Controls selected from same geographic areas; ORs stratified on age and sex.
Seidler et al., 2007	1.0	0.74	1.4	-0.223	0.177	B-cell and T-cell NHL	WHO classification. Overall results for B-cell and T-cell NHL from personal communication (see text). Adjusted for smoking and alcohol consumption. Case-control pairs matched on sex, region, and age.
Siemiatycki, 1991	1.1	0.5	2.5	0.0953	0.424	NHL	ICD-9 200 + 202. SE and 95% CI calculated from reported 90% CIs; males only; adjusted for age, income, and cigarette smoking index.
Wang et al., 2009	1.2	0.9	1.8	0.182	0.177	"NHL"; various lymphoma subtypes + mast cell tumors	ICD-O M-9590-9642, 9690-9701, 9740-9750. Females only; adjusted for age, family history of lymphohematopoietic cancers, alcohol consumption, and race.

^aThe RR estimates are all ORs for incident cases.

^bNot calculated.

NHL: non-Hodgkin lymphoma; CLL: chronic lymphocytic leukemia; HCL: hairy cell leukemia (a subgroup of NHL).

1 Most of the selections in Tables C-1 and C-2 should be self-evident, but some are
2 discussed in more detail here, in the order the studies are presented in the tables. For Boice et al.
3 (1999), results for “potential routine exposure” were selected for the primary analysis, because
4 this exposure category was considered to have less exposure misclassification, and results for
5 “any potential exposure” were used in a sensitivity analysis. The Greenland et al. (1994) study is
6 a case-control study nested within a worker cohort, and we treat it here as a cohort study (see
7 Appendix B, Section II-9.1). For Morgan et al. (1998), the reported results did not allow for the
8 combination of ICD 200 and 202, so the SMR estimate for the combined 200 + 202 grouping
9 was taken from the meta-analysis paper of Mandel et al. (2006), who included one of the
10 investigators from the Morgan et al. (1998) study. RR estimates for overall TCE exposure from
11 internal analyses of the Morgan et al. (1998) cohort data were available from an unpublished
12 report (Environmental Health Strategies, 1997; the published paper only presented the internal
13 analyses results for exposure subgroups), but only for ICD 200; from these, the RR estimate
14 from the Cox model which included age and sex was selected, because those are the variables
15 deemed to be important in the published paper (Morgan et al., 1998). Although the results from
16 internal analyses are generally preferred, in this case the SMR estimate was used in the primary
17 analysis and the internal analysis RR estimate was used in a sensitivity analysis because the latter
18 estimate represented an appreciably smaller number of deaths (3, based on ICD 200 only) than
19 the SMR estimate (9, based on ICD 200 + 202). For Radican et al. (2008), the Cox model hazard
20 ratio (HR) from the 2000 follow-up was used. In the Radican et al. (2008) Cox regressions, age
21 was the time variable, and sex and race were covariates. It should also be noted that the referent
22 group is composed of workers with no chemical exposures, not just no exposure to TCE.

23 For Zhao et al. (2005), RR estimates were only reported for ICD-9 200–208 (all
24 lymphohematopoietic cancers), and not for 200 + 202 alone. Given that other studies have not
25 reported associations between leukemias and TCE exposure, combining all lymphohematopoietic
26 cancers would dilute any lymphoma effect, and the Zhao results are expected to be an
27 underestimate of any TCE effect on lymphoma alone. Another complication with the Zhao et al.
28 (2005) study is that no results for an overall TCE effect are reported. We were unable to obtain
29 any overall estimates from the study authors, so, as a best estimate, the results across the
30 “medium” and “high” exposure groups were combined, under assumptions of group
31 independence, even though the exposure groups are not independent (the “low” exposure group
32 was the referent group in both cases). Zhao et al. (2005) present RR estimates for both incidence
33 and mortality; however, the time frame for the incidence accrual is smaller than the time frame
34 for mortality accrual and fewer exposed incident cases (17) were obtained than deaths (33).
35 Thus, because better case ascertainment occurred for mortality than for incidence, the mortality

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1 results were used for the primary analysis, and the incidence results were used in a sensitivity
2 analysis. A sensitivity analysis was also done using results from Boice et al. (2006) in place of
3 the Zhao et al. (2005) RR estimate. The cohorts for these studies overlap, so they are not
4 independent studies and should not be included in the meta-analysis concurrently. Boice et al.
5 (2006) report an RR estimate for an overall TCE effect for lymphoma alone; however, it is based
6 on far fewer cases (1 death in ICD-9 200 + 202; 9 deaths for 200–208) and is an SMR rather
7 than an internal analysis RR estimate, so the Zhao et al. (2005) estimates are preferred for the
8 primary analysis.

9 For the case-control studies, the main issue was the lymphoma classifications.
10 Miligi et al. (2006) include chronic lymphocytic leukemias (CLLs) in their NHL results,
11 consistent with the current WHO/REAL classification. Also, Miligi et al. (2006) do not report an
12 overall adjusted RR estimate, so a crude estimate of the OR was calculated for the two TCE
13 exposure categories together vs. no TCE exposure. The Nordstrom et al. (1998) study was a
14 case-control study of hairy cell leukemias (HCLs), which are a subgroup of NHLs, so only
15 results for HCL were reported. For Seidler et al. (2007), an overall adjusted OR for B-cell and
16 T-cell NHL combined was kindly provided by Dr. Seidler (personal communication from
17 Andreas Seidler, Bundesanstalt für Arbeitsschutz u. Arbeitsmedizin, to Cheryl Scott, U.S. EPA,
18 13 November 2007). Wang et al. (2009) refer to their cases as “NHL” cases; however, according
19 to the ICD-O classification system that they used, their cases are more specifically various
20 particular subtypes of malignant lymphoma (9590-9642, 9690-9701) and mast cell tumors (9740-
21 9750) (Morton et al., 2003). No alternate RR estimates were considered for any of the case-
22 control studies of lymphoma.

23 24 **C.2.1.2. Results of Meta-Analyses**

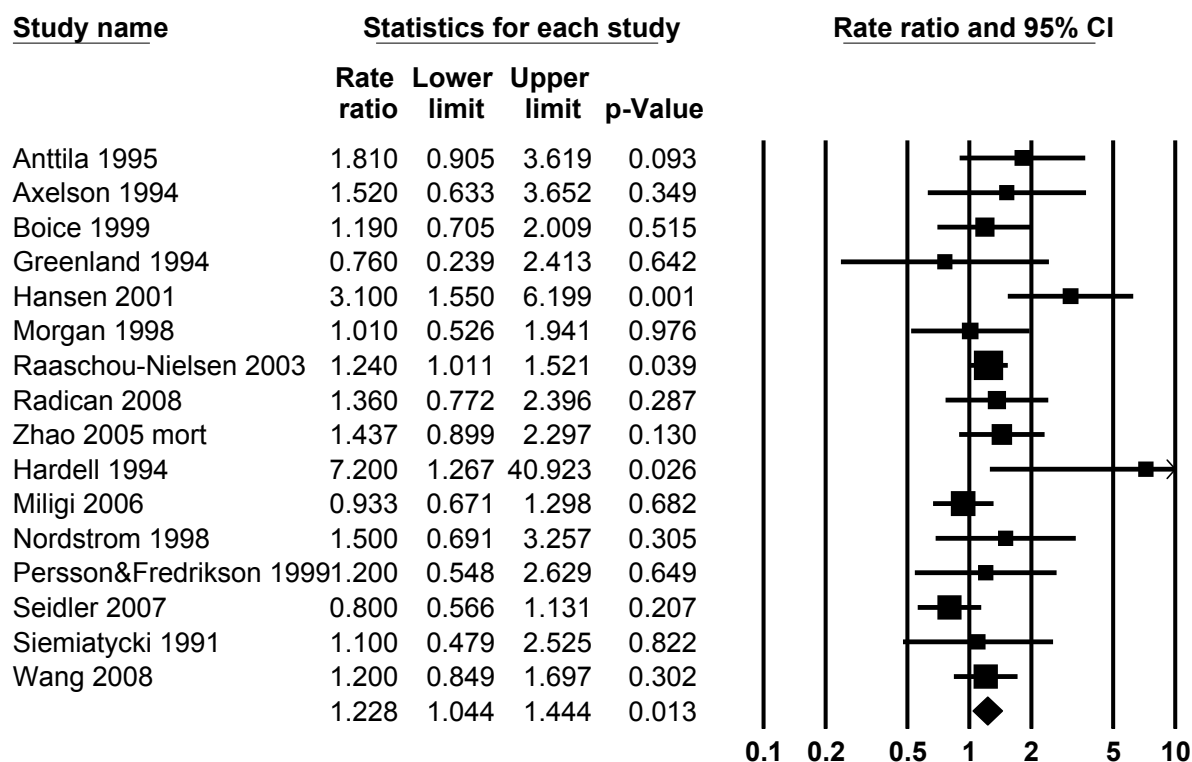
25 Results from some of the meta-analyses that were conducted on the epidemiological
26 studies of TCE and lymphoma are summarized in Table C-3. The summary estimate from the
27 primary random effects meta-analysis of the 16 studies was 1.23 (95% CI: 1.04, 1.44) (see
28 Figure C-1). No single study was overly influential; removal of individual studies resulted in
29 summary, or “pooled,” RR (RRp) estimates that ranged from 1.16 (with the removal of Hansen)
30 to 1.28 (with the removal of Seidler) and were all statistically significant. Removal of Hardell,
31 whose RR estimate is a relative outlier (see Figure C-1), only decreased the RRp estimate to 1.20
32 (1.04, 1.39), since this study does not contribute a lot of weight to the meta-analysis. Removal of
33 studies other than Hansen or Hardell resulted in RRp estimates that were all greater than 1.20.

Table C-3. Summary of some meta-analysis results for TCE (overall) and lymphoma

Analysis	# of studies	Model	Summary RR estimate (RRp)	95% LCL	95% UCL	Heterogeneity	Comments
All studies	16	Random	1.23	1.04	1.44	Not significant ($p = 0.10$)	Statistical significance of RRp not dependent on individual studies.
		Fixed	1.19	1.06	1.34		
Cohort	9	Random	1.35	1.13	1.61	Not significant ($p = 0.35$)	Not significant difference between CC and cohort studies ($p = 0.13$).
		Fixed	1.33	1.14	1.54		Significant difference between CC and cohort studies ($p = 0.03$).
Case-control	7	Random	1.07	0.84	1.37	Not significant ($p = 0.17$)	
		Fixed	1.03	0.86	1.23		
Alternate RR selections ^a	16	Random	1.19	1.00	1.41	Not significant ($p = 0.07$)	With estimated Zhao overall RR for incidence rather than mortality.
	16	Random	1.21	1.01	1.45	Not significant ($p = 0.053$)	With Boice (2006) study rather than Zhao.
	16	Random	1.22	1.04	1.44	Not significant ($p = 0.10$)	With estimated female contribution to Axelson.
	16	Random	1.22	1.05	1.43	Not significant ($p = 0.10$)	With Boice (1999) any potential exposure SMR.
	16	Random	1.24	1.05	1.46	Not significant ($p = 0.10$)	With Morgan et al. (1998) unpublished RR.
Highest exposure groups	12	Random	1.57	1.27	1.94	None observable (fixed = random)	Statistical significance not dependent on single study. See Table C-5 for results with alternate RR selections.
		Fixed	1.57	1.27	1.94		

^aChanging the primary analysis by one alternate RR each time; more details on alternate RR estimates in text.

TCE and Lymphoma



random effects model

Figure C-1. Meta-analysis of lymphoma and overall TCE exposure. The pooled estimate is in the bottom row. Symbol sizes reflect relative weights of the studies. The horizontal midpoint of the bottom diamond represents the summary RR estimate, and the horizontal extremes depict the 95% CI limits.

Similarly, the RR_p estimate was not highly sensitive to alternate RR estimate selections. Use of the five alternate selections, individually, resulted in RR_p estimates that ranged from 1.19 to 1.24 (see Table C-3) and were all statistically significant except when the Zhao incidence estimate ($p = 0.050$) was used instead of the Zhao mortality estimate. As discussed above, the Zhao mortality estimate is preferred over the incidence estimate in this instance because it is based on nearly twice as many cases (33 vs. 17).

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There was some heterogeneity apparent across the 16 studies, although it was not statistically significant ($p = 0.10$). The I^2 value (see Section C.1) was 33%, suggesting low-to-moderate heterogeneity. Subgroup analyses were done examining the cohort and case-control studies separately. With the random effects model (and tau-squared not pooled across subgroups), the resulting RRp estimates were 1.35 (95% CI: 1.13, 1.61) for the cohort studies and 1.07 (0.84, 1.37) for the case-control studies. There was residual heterogeneity in each of the subgroups, but in neither case was it statistically. I^2 values were 10% for the cohort studies, suggesting low heterogeneity, and 33% for the case-control studies, suggesting low-to-moderate heterogeneity. The difference between the RRp estimates for the cohort and case-control subgroups was not statistically significant under the random effects model, although it was under the fixed effect model (see Table C-3). Some thought was given to further analyses to investigate the source(s) of the heterogeneity, such as qualitative tiering or subgroups based on likelihood for correct exposure classification or on likelihood for higher vs. lower exposures across the studies. Ultimately, these approaches were rejected because in many of the studies it was difficult to judge (and weight) the extent of exposure misclassification or the degree of TCE exposure with any precision. In other words, there was inadequate information to reliably assess either the extent to which each study accurately classified exposure status or the relative TCE exposure levels and prevalences of exposure to different levels across studies. See Section C.2.3 below for a qualitative discussion of some potential sources of heterogeneity.

As discussed in Section C.1, publication bias was examined in several different ways. The funnel plot in Figure C-2 suggests some relationship between RR estimate and study size (if there were no relationship, the studies would be symmetrically distributed around the pooled RR estimate rather than veering towards higher RR estimates with increasing SEs), although the observed asymmetry is highly influenced by the Hardell study, which is a relative outlier and which contributes little weight to the overall meta-analysis, as discussed above. The Begg and Mazumdar rank correlation test and Egger's linear regression test were not statistically significant; it should be noted, however, that both of these tests have low power. Duval and Tweedie's trim-and-fill procedure yielded a pooled RR estimate (under the random effects model) of 1.13 (95% CI: 0.94, 1.35) when the 4 studies deemed missing from the funnel plot were filled into the meta-analysis (these studies are filled in so as to counter-balance the apparent asymmetry of the more extreme values in the funnel plot). Eliminating the Hardell study made little difference to the results of the publication bias analyses. The results of a cumulative meta-analysis, incorporating studies with increasing SE one at a time, are depicted in Figure C-3. This procedure is a transparent way of examining the effects of including studies with increasing SE. The figure shows that the pooled RR estimate is 1.05 after inclusion of the 4 largest (i.e.,

most precise) studies, which constitute about 50% of the weight. The pooled RR estimate increases to 1.12 with inclusion of the 8 most precise studies, which represent ½ of the total number of studies and about 75% of the weight. The pooled RR estimate becomes fairly stable after addition of the next 2 most precise study (RRp = 1.21), which adds another 9% of the weight. Adding in the 6 least precise studies (16% of the weight) barely increases the pooled RR estimate further. In summary, there is some evidence of potential publication bias in this data set. It is uncertain, however, that this reflects actual publication bias rather than an association between effect size and SE resulting for some other reason, e.g., a difference in study populations or protocols in the smaller studies. Furthermore, if there is publication bias in this data set, it does not appear to account completely for the findings of an increased lymphoma risk.

Funnel Plot of Standard Error by Log rate ratio

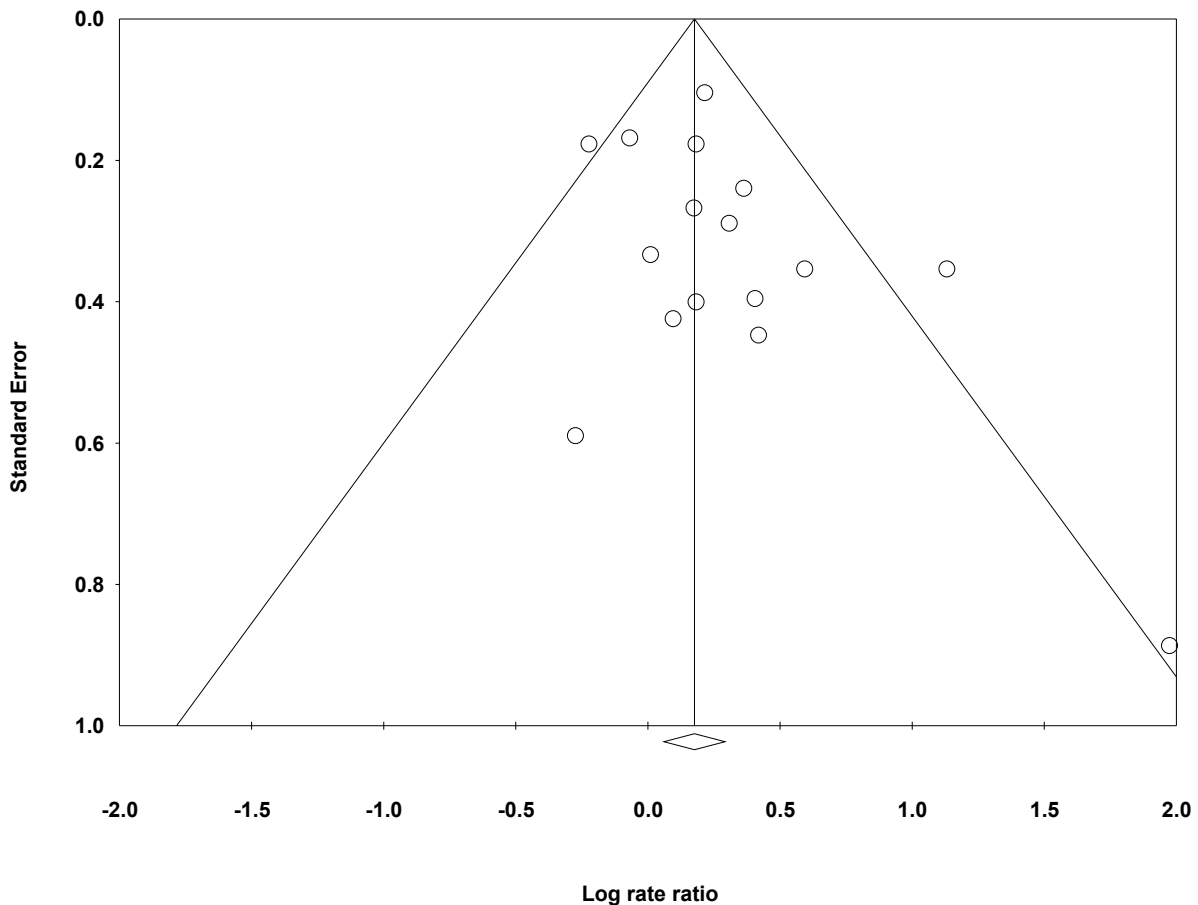
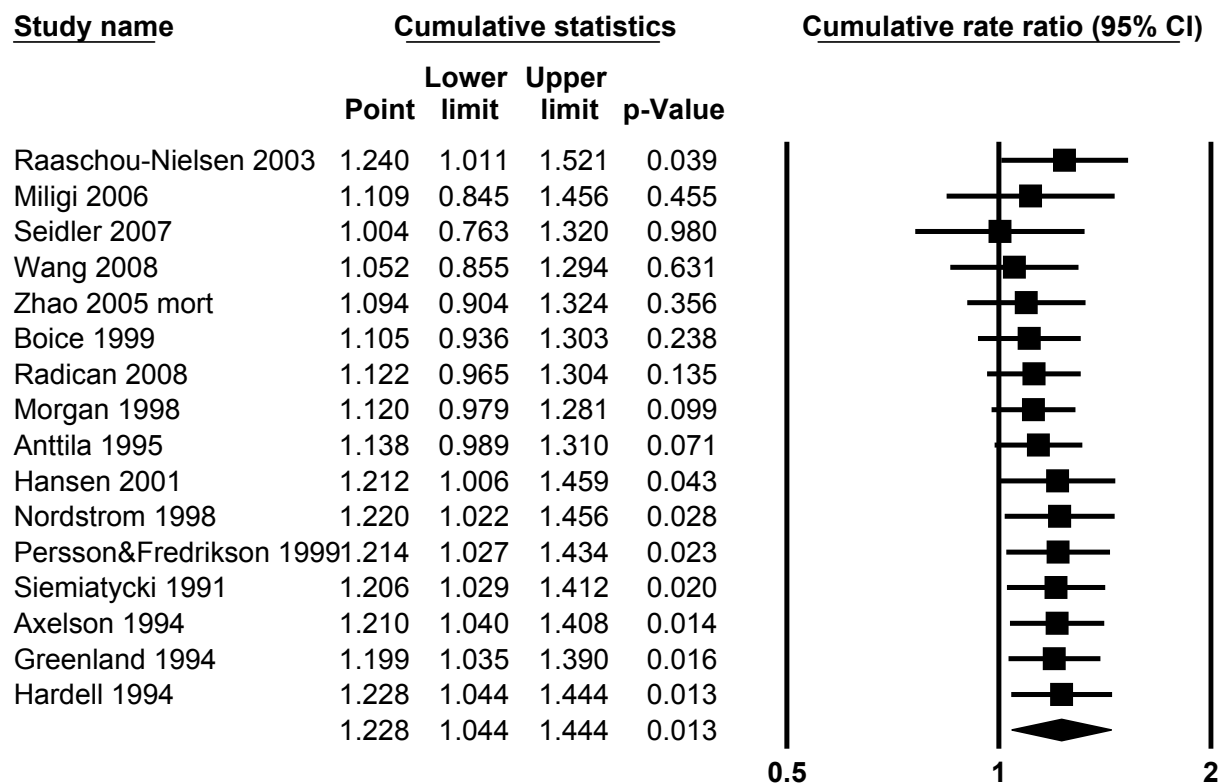


Figure C-2. Funnel plot of SE by log RR estimate for TCE and lymphoma studies.

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TCE and Lymphoma



random effects model; cumulative analysis, sorted by SE

Figure C-3. Cumulative meta-analysis of TCE and lymphoma studies, progressively including studies with increasing SEs.

C.2.2. Lymphoma Effect in the Highest Exposure Groups

C.2.2.1. Selection of RR Estimates

The selected RR estimates for lymphoma in the highest TCE exposure categories, for studies that provided such estimates, are presented in Table C-4. All 8 cohort studies (but not the nested case-control study of Greenland et al. [1994]) and 4 of the 7 case-control studies did report lymphoma risk estimates categorized by exposure level. As in Section C.2.1.1 for the overall risk estimates, estimates to best correspond to NHL as represented by ICD-7, -8, and -

9 200 and 202 were selected, and, wherever possible, RR estimates for males and females combined were used.

As above for the overall TCE effect, for Axelson et al. (1994), in which a small subcohort of females was studied but only results for the larger male subcohort were reported, the reported male-only high-exposure group results were used in the primary analysis; however, an attempt was made to estimate the female contribution to a high-exposure group RR estimate for both sexes and its impact on the meta-analysis. To estimate the expected number in the highest exposure group for females, the expected number in the highest exposure group for males was multiplied by the ratio of total female-to-male person-years in the study and by the ratio of female-to-male age-adjusted incidence rates for NHL. The RR estimate for both sexes was used as an alternate RR estimate for the Axelson et al. (1994) study in a sensitivity analysis.

For Boice et al. (1999), only results for workers with “any potential exposure” (rather than “potential routine exposure”) were presented by exposure category, and the referent group is workers not exposed to any solvent. For Hansen et al. (2001), exposure group data were presented only for males. To estimate the female contribution to a highest-exposure group RR estimate for both sexes, it was assumed that the expected number of cases in females had the same overall-to-highest-exposure group ratio as in males. The RR estimate for both sexes was then calculated assuming a Poisson distribution, and this estimate was used in the primary analysis. Hansen et al. (2001) present results for three exposure metrics; the cumulative exposure metric was preferred for the primary analysis, and results for the other two metrics were used in sensitivity analyses. For Morgan et al. (1998), results did not allow for the combination of ICD 200 and 202, so the highest-exposure group RR estimate for ICD 200 only was used. The primary analysis used results for the cumulative exposure metric, and a sensitivity analysis was done with the results for the peak exposure metric.

For Radican et al. (2008), it should be noted that the referent group is composed of workers with no chemical exposures, not just no exposure to TCE. In addition, exposure group results were reported separately for males and females and were combined for this assessment using inverse-variance weighting, as in a fixed effect meta-analysis. Radican et al. (2008) present only mortality HR estimates by exposure group; however, in an earlier follow-up of this same cohort, Blair et al. (1998) present both incidence and mortality RR estimates by exposure group. The mortality RR estimate based on the more recent follow-up of Radican et al. (2008) (17 deaths in the highest exposure group) was used in the primary analysis, while the incidence RR estimate based on similarly combined results from Blair et al. (1998) (9 cases) was used as an alternate estimate in a sensitivity analysis.

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Table C-4. Selected RR estimates for lymphoma risk in highest TCE exposure groups

Study	RR	95% LCL	95% UCL	Exposure category	log RR	SE(log RR)	Alternate RR estimates	Comments
Anttila et al., 1995	1.4	0.17	5.04	100+ $\mu\text{mol/L}$ U-TCA ^a	0.336	0.707	none	SIR. ICD 200 + 202.
Axelson et al., 1994	6.25	0.16	34.83	≥ 2 -yr exposure and 100+ mg/L U-TCA	1.83	1.00	5.62 (0.14, 31.3) with estimated female contribution added (see text)	SIR. ICD 200 + 202. Results reported for males only, but there was a small female component to the cohort.
Boice et al., 1999	1.62	0.82	3.22	≥ 5 -yr exposure	0.482	0.349	None	Mortality RR. ICD 200 + 202. For potential routine or intermittent exposure. Adjusted for date of birth, dates 1 st and last employed, race, and sex. Referent group is workers not exposed to any solvent.
Hansen et al., 2001	2.7	0.56	8.0	≥ 1080 mos \times mg/m ³	0.993	0.577	3.7 (1.0, 9.5) for ≥ 75 mos exposure duration 2.9 (0.79, 7.5) for ≥ 19 mg/m ³ mean exposure	SIR. ICD 200 + 202. Exposure-group results presented only for males. Female results estimated and combined with male results assuming Poisson distribution (see text).
Morgan et al., 1998	0.81	0.1	6.49	High cumulative exp. score	-0.211	1.06	1.31 (0.28, 6.08) for med/high peak vs. low/no	Mortality RR. ICD 200 only. Adjusted for age and sex.
Raaschou-Nielsen et al., 2003	1.6	1.1	2.2	≥ 5 yrs in subcohort with expected higher exp. levels	0.470	0.183	None	SIR. ICD 200 + 202.

Table C-4. Selected RR estimates for lymphoma risk in highest TCE exposure groups (continued)

Study	RR	95% LCL	95% UCL	Exposure category	log RR	SE(log RR)	Alternate RR estimates	Comments
Radican et al., 2008	1.41	0.71	2.81	>25 unit-yr	0.337	0.350	Blair et al. (1998) 0.97 (0.42, 2.2) incidence RR	Mortality HR. ICD 200 + 202. Male and female results presented separately and combined (see text). Cox regression time variable = age; covariate = race. Referent group is workers with no chemical exposures.
Zhao et al., 2005	1.30	0.52	3.23	High exposure score	0.262	0.466	Incidence RR: 0.20 (0.03, 1.46)	Mortality RR. Results for all lymphohematopoietic cancer (ICD-9 200–208), not just 200 + 202. Males only; adjusted for age, SES, time since first employment. Mortality results reflect more exposed cases (6 in high-exposure group) than do incidence results (1 in high-exposure group).
Miligi et al., 2006	1.2	0.7	2.0	Med/high exposure intensity	0.182	0.268	1.0 (0.5, 2.6) for med/high intensity and >15-yr exp.	Incidence OR. NHL + CLL (see Section C.2.1.1).
Seidler et al., 2007	2.3	1.0	5.2	>35 ppm-yr	0.833	0.421	None	Incidence OR. Results for B-cell and T-cell NHL from personal communication (see Section C.2.1.1). Adjusted for smoking and alcohol consumption. Case-control pairs matched on sex, region, and age.
Siemiatycki 1991	0.8	0.2	3.3	Substantial	-0.223	0.719	None	Incidence OR. NHL. SE and 95% CI calculated from reported 90% CIs. Males only; adjusted for age, income, and cigarette smoking index.
Wang et al., 2009	2.2	0.9	5.4	Medium-high intensity	0.788	0.457	None	Incidence OR. "NHL" (various malignant lymphoma subtypes and mast cell tumors). Females only; adjusted for age, family history of lymphohematopoietic cancers, alcohol consumption, and race.

^aMean personal trichloroacetic acid in urine. 1 µmol/L = 0.1634 mg/L.

1 For Zhao et al. (2005), RR estimates were only reported for ICD-9 200–208 (all
2 lymphohematopoietic cancers), and not for 200 + 202 alone. Given that other studies have not
3 reported associations between leukemias and TCE exposure, combining all lymphohematopoietic
4 cancers would dilute any lymphoma effect, and the Zhao results are expected to be an
5 underestimate of any TCE effect on lymphoma alone. Zhao et al. (2005) present RR estimates
6 for both incidence and mortality in the highest exposure group; however, the time frame for the
7 incidence accrual is smaller than the time frame for mortality accrual and fewer incident cases
8 (1) were obtained than deaths (6), so the mortality results were used for the primary analysis to
9 reflect the better case ascertainment in the mortality data, and the incidence results were used in
10 a sensitivity analysis.

11 Miligi et al. (2006) include CLLs in their NHL results, consistent with the current
12 WHO/REAL classification. Miligi et al. (2006) report RR estimates for medium and high
13 exposure intensity overall and by duration of exposure; however, there was incomplete
14 information for the duration breakdowns (e.g., a case missing), so the RR estimate for med/high
15 exposure intensity overall was used in the primary analysis, and the RR estimate for med/high
16 exposure for >15 years was used in a sensitivity analysis. For Seidler et al. (2007), an adjusted
17 OR for B-cell and T-cell NHL combined for the >35 ppm-years exposure category was kindly
18 provided by Dr. Seidler (personal communication from Andreas Seidler, Bundesanstalt fur
19 Arbeitsschutz u. Arbeitsmedizin, to Cheryl Scott, U.S. EPA, 13 November 2007). Wang et al.
20 (2009) refer to their cases as "NHL" cases; however, according to the ICD-O classification
21 system that they used, their cases are more specifically various particular subtypes of malignant
22 lymphoma (9590-9642, 9690-9701) and mast cell tumors (9740-9750) (Morton et al., 2003).

23 24 **C.2.2.2. Results of Meta-Analyses**

25 Results from the meta-analyses that were conducted for lymphoma in the highest exposure
26 groups are summarized at the bottom of Table C-3 and reported in more detail in Table C-5. The
27 pooled RR estimate from the primary random effects meta-analysis of the 12 studies with results
28 presented for exposure groups was 1.57 (95% CI: 1.27, 1.94) (see Figure C-4). No single study
29 was overly influential; removal of individual studies resulted in RRp estimates that were all
30 statistically significant (all with $p \leq 0.001$) and that ranged from 1.53 (with the removal of
31 Seidler) to 1.65 (with the removal of Miligi). Similarly, the RRp estimate was not highly
32 sensitive to alternate RR estimate selections. Use of the 7 alternate selections, individually,
33 resulted in RRp estimates that were all statistically significant (all with $p < 0.001$) and all in the
34 narrow range from 1.54 to 1.60 (see Table C-5). There was no observable heterogeneity across
35 the 12 studies in either the primary analysis or any of the alternate RR analyses.

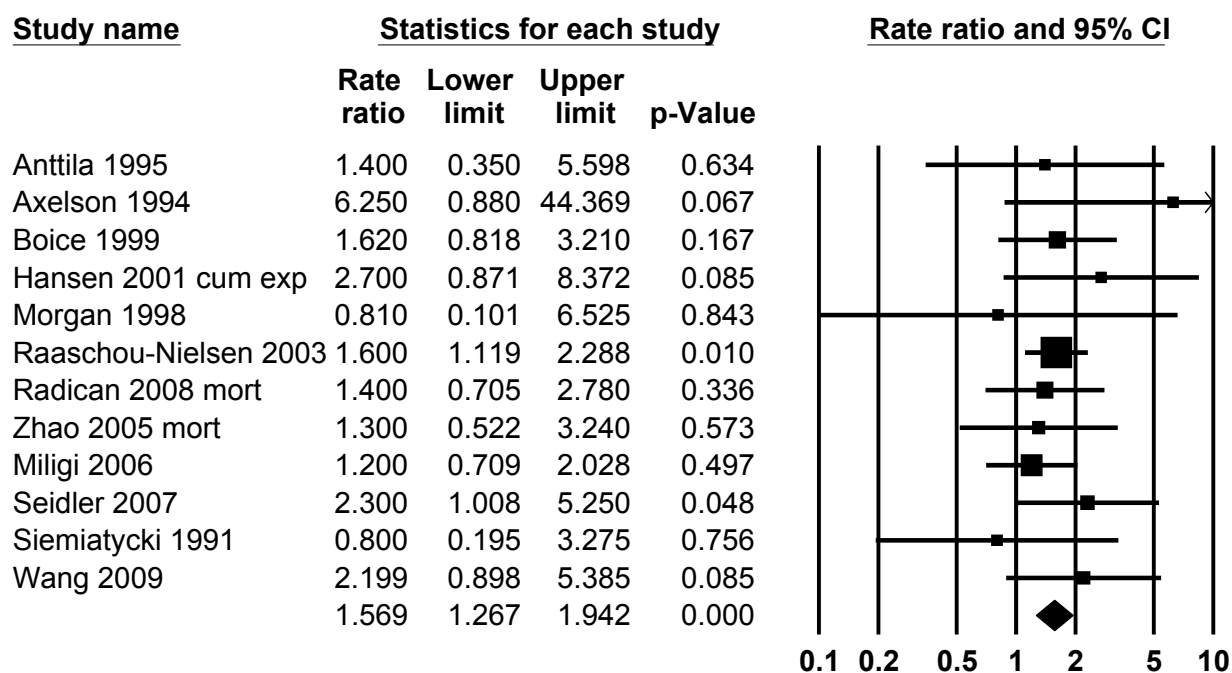
Table C-5. Summary of some meta-analysis results for TCE (highest exposure groups) and lymphoma

Analysis	Model	Combined RR estimate	95% LCL	95% UCL	Heterogeneity	Comments
Primary analysis	Random	1.57	1.27	1.94	None obs (fixed = random)	Statistical significance not dependent on single study.
Alternate RR selections ^a	Random	1.54	1.24	1.91	None obs	With Blair et al. (1998) incidence RR instead of Radican mortality HR.
	Random	1.55	1.24	1.92	None obs	With Zhao incidence.
	Random	1.57	1.27	1.94	None obs	With estimated female contribution for Axelson.
	Random	1.57	1.27	1.95	None obs	With Morgan peak.
	Random	1.58	1.28	1.96	None obs	With Hansen mean exposure.
	Random	1.60	1.28	2.00	None obs	With Miligi with >15 yrs.
	Random	1.60	1.30	1.98	None obs	With Hansen duration.

^aChanging the primary analysis by one alternate RR estimate each time.

obs = observable.

TCE and Lymphoma - highest exposure groups



random effects model; same for fixed

1
2 **Figure C-4. Meta-analysis of lymphoma and TCE exposure—highest exposure**
3 **groups.** (The pooled estimate is in the bottom row. Symbol sizes reflect relative
4 weights of the studies. The horizontal midpoint of the bottom diamond represents
5 the pooled RR estimate, and the horizontal extremes depict the 95% CI limits.)
6
7

8 C.2.3. Discussion of Lymphoma Meta-Analysis Results

9 For the most part, the meta-analyses of the overall effect of TCE exposure on lymphoma
10 suggest a small, statistically significant increase in risk. The pooled estimate from the primary
11 random effects meta-analysis of the 16 studies was 1.23 (95% CI: 1.04, 1.44). This result was
12 not overly influenced by any single study, nor was it overly sensitive to individual RR estimate
13 selections. In terms of the statistical significance of the RRp estimate, the only alternate analysis
14 (involving either a study removal or an alternate RR estimate) that did not yield a statistically
15 significant RRp was the analysis in which the Zhao mortality RR estimate was substituted with
16 the incidence estimate, resulting in an RRp estimate of 1.19 (1.00, 1.41); although, as noted

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1 above, this substitution is considered clearly inferior to the Zhao mortality estimate that was used
2 in the primary analysis. Thus, the finding of an increased risk of lymphoma associated with TCE
3 exposure, though the increased risk is not large in magnitude, is fairly robust.

4 There is some evidence of potential publication bias in this data set; however, it is
5 uncertain that this is actually publication bias rather than an association between SE and effect
6 size resulting for some other reason, e.g., a difference in study populations or protocols in the
7 smaller studies. Furthermore, if there is publication bias in this data set, it does not appear to
8 account completely for the finding of an increased lymphoma risk.

9 Although there was some heterogeneity across the 16 studies, it was not statistically
10 significant ($p = 0.10$). The I^2 value was 33%, suggesting low-to-moderate heterogeneity.
11 Similarly, when subgroup analyses were done of cohort and case-control studies separately, there
12 was some observable heterogeneity in each of the subgroups, but it was not statistically
13 significant in either case. I^2 values were 10% for the cohort studies, suggesting low
14 heterogeneity, and 33% for the case-control studies, suggesting low-to-moderate heterogeneity.
15 In the subgroup analyses, the increased risk of lymphoma was strengthened in the cohort study
16 analysis and virtually eliminated in the case-control study analysis, although the subgroup RRs
17 estimates were not statistically significantly different under the random effects model. Study
18 design itself is unlikely to be an underlying cause of heterogeneity and, to the extent that it may
19 explain some of the differences across studies, is more probably a surrogate for some other
20 difference(s) across studies that may be associated with study design. Furthermore, other
21 potential sources of heterogeneity may be masked by the broad study design subgroupings. The
22 true source(s) of heterogeneity across these studies is an uncertainty. As discussed above, further
23 quantitative investigations of heterogeneity were ruled out because of database limitations. A
24 qualitative discussion of some potential sources of heterogeneity follows.

25 Study differences in exposure assessment approach, exposure prevalence, average
26 exposure intensity, and lymphoma classification are possible sources of heterogeneity. Many
27 studies included TCE assignment from information on job and task exposures, e.g., a
28 job-exposure matrix (JEM) (Siemiatycki, 1991; Morgan et al., 1998; Boice et al., 1999, 2006;
29 Zhao et al., 2005; Miligi et al., 2006; Seidler et al., 2007; Radican et al., 2008; Wang et al.,
30 2009), or from an exposure biomarker in either breath or urine (Axelson et al., 1994; Anttila et
31 al., 1995; Hansen et al., 2001). Three case-control studies relied on self-reported exposure to
32 TCE (Hardell et al., 1994; Nordstrom et al., 1998; Persson and Fredrikson, 1999).
33 Misclassification is possible with all exposure assessment approaches. No information is
34 available to judge the degree of possible misclassification bias associated with a particular
35 exposure assessment approach; it is quite possible that in some cohort studies, in which past

1 exposure is inferred from various data sources, exposure misclassification may be as great as in
2 population-based or hospital-based case-control studies. Approaches based upon JEMs can
3 provide order-of-magnitude estimates that are useful for distinguishing groups of workers with
4 large differences in exposure; however, smaller differences usually cannot be reliably
5 distinguished (NRC, 2006). Biomonitoring can provide information on potential TCE exposure
6 in an individual, but the biomarkers used aren't necessarily specific for TCE and they reflect only
7 recent exposures. The lack of heterogeneity in the analysis of the highest exposure groups
8 provides some evidence of exposure misclassification as a source of heterogeneity in the overall
9 analysis.

10 General population studies have special problems in evaluating exposure, because the
11 subjects could have worked in any job or setting that is present within the population (Copeland
12 et al., 1977; Nelson et al., 1994; McGuire et al., 1998; 't Mannetje et al., 2002; NRC, 2006).
13 Low exposure prevalence in the four population case-control studies (Siemiatycki, 1991;
14 Miligi et al., 2006; Seidler et al., 2007; Wang et al., 2009) may be another source of
15 heterogeneity. Prevalence of TCE exposure among cases in the case-control studies was low,
16 ranging from 3% in Siemiatycki (1991) to 13% in Seidler et al. (2007) and Wang et al. (2009).
17 However, prevalence of high TCE exposure in these case-control studies was even rarer—3% of
18 all cases in Miligi et al. (2006) and Seidler et al. (2007), 2% in Wang et al. (2009), and less than
19 1% in Siemiatycki (1991). Low exposure prevalence, especially in the relatively large Miligi et
20 al. (2006) and Seidler et al. (2007) case-control studies (see Figure C-1), may be one of the
21 underlying characteristics differentiating the case-control and cohort studies and explaining some
22 of the heterogeneity across the studies.

23 Study differences in lymphoma groupings and in lymphoma classification schemes are
24 another potential source of heterogeneity in the meta-analysis. All studies included a broad but
25 sometimes slightly different group of lymphosarcoma, reticulum-cell sarcoma, and other
26 lymphoid tissue neoplasms, with the exception of the Nordstrom et al. (1998) case-control study,
27 which examined hairy cell leukemia, now considered a lymphoma, and the Zhao et al. (2005)
28 cohort study, which reported only results for *all* lymphohematopoietic cancers, including
29 nonlymphoid types. Persson and Fredrikson (1999) do not identify the classification system for
30 defining NHL, and Hardell et al. (1994) define NHL using the Rappaport classification system.
31 Miligi et al. (2006) used an NCI classification system and considered chronic lymphocytic
32 leukemias and NHLs together as lymphomas, while Seidler et al. (2007) used the REAL
33 classification system, which reclassifies lymphocytic leukemias and NHLs as lymphomas of
34 B-cell or T-cell origin. The cohort studies (except for Zhao et al.) and the case-control study of
35 Siemiatycki (1991) have some consistency in coding NHL, with NHL defined as lymphosarcoma

1 and reticulum-cell sarcoma (ICD code 200) and other lymphoid tissue neoplasms (ICD 202)
2 using the ICD Revisions 7, 8, or 9. Revisions 7 and 8 are essentially the same with respect to
3 NHL; under Revision 9, the definition of NHL was broadened to include some neoplasms
4 previously classified as Hodgkin's lymphomas (Banks, 1992). Wang et al. (2009) refer to their
5 cases as "NHL" cases; however, according to the ICD-O classification system that they used,
6 their cases are more specifically various particular subtypes of malignant lymphoma (9590-9642,
7 9690-9701) and mast cell tumors (9740-9750) (Morton et al., 2003).

8 Twelve of the 16 studies categorized results by exposure level. Different exposure
9 metrics were used, and the purpose of combining results across the different highest exposure
10 groups was not to estimate an RRp associated with some level of exposure, but rather to see the
11 impacts of combining RR estimates that should be less affected by exposure misclassification.
12 In other words, the highest exposure category is more likely to represent a greater differential
13 TCE exposure compared to people in the referent group than the exposure differential for the
14 overall (typically any vs. none) exposure comparison. Thus, if TCE exposure increases the risk
15 of lymphoma, the effects should be more apparent in the highest exposure groups. Indeed, the
16 RRp estimate from the primary meta-analysis of the highest exposure group results was 1.57
17 (95% CI: 1.27, 1.94), which is greater than the RRp estimate of 1.23 (95% CI: 1.04, 1.44) from
18 the overall exposure analysis. This result for the highest exposure groups was not overly
19 influenced by any single study, nor was it overly sensitive to individual RR estimate selections.
20 Heterogeneity was not observed in any of the relevant analyses. The robustness of this finding
21 lends substantial support to a conclusion that TCE exposure increases the risk of lymphoma.
22

23 **C.3. META-ANALYSIS FOR KIDNEY CANCER**

24 **C.3.1. Overall Effect of TCE Exposure**

25 **C.3.1.1. Selection of RR Estimates**

26 The selected RR estimates for kidney cancer associated with TCE exposure from the
27 epidemiological studies are presented in Table C-6 for cohort studies and in Table C-7 for
28 case-control studies. The majority of the cohort studies reported results for all kidney cancers,
29 including cancers of the renal pelvis and ureter (i.e., ICD-7 180; ICD-8 and -9 189.0–189.2;
30 ICD-10 C64–C66); whereas the majority of the case-control studies focused on renal cell
31 carcinoma (RCC), which comprises roughly 85% of kidney cancers. Where both all kidney
32 cancer and RCC were reported, the primary analysis used the results for RCC, because RCC and
33 the other forms of kidney cancer are very different cancer types and it seemed preferable not to
34 combine them; the results for all kidney cancers were then used in a sensitivity analysis. The
35 preference for the RCC results alone is supported by the results in rodent cancer bioassays,

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1 where TCE-associated rat kidney tumors are observed in the renal tubular cells (Section 4.3.5),
2 and in metabolism studies, where the focus of studies for the GSH conjugation pathway
3 (considered the primary metabolic pathway for kidney toxicity) is in renal cortical and tubular
4 cells (Sections 3.3.3.2 and 4.3.6).

5 As for lymphoma, many of the studies provided RR estimates only for males and females
6 combined, and we are not aware of any basis for a sex difference in the effects of TCE on kidney
7 cancer risk; thus, wherever possible, RR estimates for males and females combined were used.
8 Of the three larger (in terms of number of cases) studies that did provide results separately by
9 sex, Dosemeci et al. (1999) suggest that there may be a sex difference for TCE exposure and
10 RCC (OR = 1.04 [95% CI: 0.6, 1.7] in males and 1.96 [1.0, 4.0] in females), while
11 Raaschou-Nielsen et al. (2003) report the same SIR (1.2) for both sexes and crude ORs
12 calculated from data from the Pesch et al. (2000) study (provided in a personal communication
13 from Baeta Pesch, Forschungsinstitut für Arbeitsmedizin (BGFA), to Cheryl Scott, U.S. EPA,
14 21 February 2008) are 1.28 for males and 1.23 for females. Radican et al. (2008) and Hansen et
15 al. (2001) also present some results by sex, but both of these studies have too few cases to be
16 informative about a sex difference for kidney cancer.

17 Most of the selections in Tables C-6 and C-7 should be self-evident, but some are
18 discussed in more detail here, in the order the studies are presented in the tables. For Axelson et
19 al. (1994), in which a small subcohort of females was studied but only results for the larger male
20 subcohort were reported, the reported male-only results were used in the primary analysis;
21 however, as for lymphoma, an attempt was made to estimate the female contribution to an
22 overall RR estimate for both sexes and its impact on the meta-analysis. Axelson et al. (1994)
23 reported neither the observed nor the expected number of kidney cancer cases for females. It
24 was assumed that none were observed. To estimate the expected number, the expected number
25 for males was multiplied by the ratio of female-to-male person-years in the study and by the ratio
26 of female-to-male age-adjusted incidence rates for kidney cancer.² The male results and the
27 estimated female contribution were then combined into an RR estimate for both sexes assuming
28 a Poisson distribution, and this alternate RR estimate for the Axelson et al. (1994) study was
29 used in a sensitivity analysis.

²Person-years for men and women ≤ 79 years were obtained from Axelson et al. (1994): 23516.5 and 3691.5, respectively. Lifetime age-adjusted incidence rates for cancer of the kidney and renal pelvis for men and women were obtained from the National Cancer Institute's 2000–2004 SEER-17 (Surveillance Epidemiology and End Results from 17 geographical locations) database (<http://seer.cancer.gov/statfacts/html/kidrp.html>): 17.8/100,000 and 8.8/100,000, respectively. The calculation for estimating the expected number of cases in females in the cohort assumes that the males and females have similar TCE exposures and that the relative distributions of age-related incidence risk for the males and females in the cohort are adequately represented by the ratios of person-years and lifetime incidence rates used in the calculation.

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Table C-6. Selected RR estimates for kidney cancer associated with TCE exposure (overall effect) from cohort studies

Study	RR	95% LCL	95% UCL	RR type	log RR	SE(log RR)	Alternate RR estimates	Comments
Anttila et al., 1995	0.87	0.32	1.89	SIR	-0.139	0.408	none	ICD-7 180.
Axelson et al., 1994	1.16	0.42	2.52	SIR	0.148	0.408	1.07 (0.39, 2.33) with estimated female contribution to SIR added (see text)	ICD-7 180. Results reported for males only, but there was a small female component to the cohort.
Boice et al., 1999	0.99	0.4	2.04	SMR	-0.010	0.378	None	ICD-9 189.0–189.2. For potential routine exposure. Results for any potential exposure not reported.
Greenland et al., 1994	0.99	0.30	3.32	OR	-0.010	0.613	None	Nested case-control study. ICD-8 codes not specified, presumably all of 189.
Hansen et al., 2001	1.1	0.3	2.8	SIR	0.095	0.500	None	ICD-7 180. Male and female results reported separately; combined assuming Poisson distribution.
Morgan et al., 1998	1.14	0.51	2.58	Mortality RR	0.134	0.415	Published SMR 1.32 (0.57, 2.6)	ICD-9 189.0–189.2. Unpublished RR, adjusted for age and sex (see text).
Raaschou-Nielsen et al., 2003	1.20	0.94	1.50	SIR	0.182	0.199	1.20 (0.98, 1.46) for ICD-7 180	RCC.
Radican et al., 2008	1.18	0.47	2.94	Mortality HR	0.166	0.468	None	ICD-8, -9 189.0, ICD-10 C64. Time variable = age; covariates = sex and race. Referent group is workers with no chemical exposures.
Zhao et al., 2005	1.7	0.38	7.9	Mortality RR	0.542	0.775	Incidence RR: 2.0 (0.47, 8.2) Mortality RR no lag: 0.89 (0.22, 3.6) Incidence RR no lag : 2.1 (0.56, 8.1) Boice (2006) SMR: 2.22 (0.89, 4.57)	ICD-9 189. Males only. Adjusted for age, SES, time since first employment, exposure to other carcinogens. 20-yr lag. Mortality results reflect same number exposed cases (10 with no lag) as do incidence results, so no reason to prefer mortality results, but they are used in primary analysis to avoid appearance of “cherry-picking.” Overall RR estimated by combining across exposure groups (see text). Boice (2006) cohort overlaps Zhao cohort; just 7 exposed deaths.

Table C-7. Selected RR estimates for renal cell carcinoma associated with TCE exposure from case-control studies^a

Study	RR estimate	95% LCL	95% UCL	log RR	SE(log RR)	Alternate RR estimates	Comments
Brüning et al., 2003	2.47	1.36	4.49	0.904	0.305	1.80 (1.01, 3.20) for longest job held in industry with TCE exposure	Self-assessed exposure. Adjusted for age, sex, and smoking.
Charbotel et al., 2006	1.88	0.89	3.98	0.631	0.382	1.64 (0.95, 2.84) for full study	Subgroup with good level of confidence about exp assessment. Matched on sex, age. Adjusted for smoking, body mass index.
Dosemeci et al., 1999	1.30	0.9	1.9	0.262	0.191		Adjusted for age, sex, smoking, hypertension and/or use of diuretics and/or anti-hypertension drugs, body mass index.
Pesch et al., 2000	1.24	-- ^b	-- ^b	0.215	0.094	1.13 with German JEM	With JTEM (job task exposure matrix). Crude OR calculated from data provided in personal communication (see text).
Siemiatycki 1991	0.8	0.3	2.2	-0.223	0.524		"Kidney cancer." SE and 95% CI calculated from reported 90% CIs. Males only; adjusted for age, income, and cigarette smoking index.

^aThe RR estimates are all ORs for incident cases.

^bNot calculated.

1 For Boice et al. (1999), only results for “potential routine exposure” were reported for
2 kidney cancer. This is our preferred TCE exposure definition for the Boice study, because it was
3 considered to have less exposure misclassification than “any potential exposure;” however, since
4 the results for the latter definition were not presented, they could not be used in a sensitivity
5 analysis, as was done for lymphoma. Boice et al. (1999) report in general that the SMRs for
6 workers with any potential exposure “were similar to those for workers with daily potential
7 exposure.” In their published paper, Morgan et al. (1998) present only SMRs for overall TCE
8 exposure, although the results from internal analyses are presented for exposure subgroups. RR
9 estimates for overall TCE exposure from the internal analyses of the Morgan et al. (1998) cohort
10 data were available from an unpublished report (Environmental Health Strategies, 1997); from
11 these, the RR estimate from the Cox model which included age and sex was selected, because
12 those are the variables deemed to be important in the published paper. The internal analysis RR
13 estimate was preferred for the primary analysis, and the published SMR result was used in a
14 sensitivity analysis. Raaschou-Nielsen et al. (2003) reported results for RCC and renal
15 pelvis/ureter separately. As discussed above, RCC estimates were used in the primary analysis,
16 and the results for both kidney cancer categories were combined (across sexes as well), assuming
17 a Poisson distribution, and used in a sensitivity analysis. For Radican et al. (2008), the Cox
18 model hazard ratio (HR) from the 2000 follow-up was used. In the Radican et al. (2008) Cox
19 regressions, age was the time variable, and sex and race were covariates. It should also be noted
20 that the referent group is composed of workers with no chemical exposures, not just no exposure
21 to TCE.

22 For Zhao et al. (2005), no results for an overall TCE effect are reported. We were unable
23 to obtain any overall estimates from the study authors, so, as a best estimate, as was done for
24 lymphoma, the results across the “medium” and “high” exposure groups were combined, under
25 assumptions of group independence, even though the exposure groups are not independent (the
26 “low” exposure group was the referent group in both cases). Unlike for lymphoma, adjustment
27 for exposure to other carcinogens made a considerable difference, so Zhao et al. (2005) also
28 present kidney results with this additional adjustment, with and without a 20-year lag. Estimates
29 of RR with this additional adjustment were selected over those without. In addition, a 20-year
30 lag seemed reasonable for kidney cancer, so the lagged estimates were preferred to the unlagged;
31 unlagged estimates were used in sensitivity analyses. Zhao et al. (2005) present RR estimates for
32 both incidence and mortality. Unlike for lymphoma, the number of exposed incident cases (10
33 with no lag) was identical to the number of deaths, so there was no reason to prefer the mortality
34 results over the incidence results. (In fact, there were more exposed incident cases [10 vs. 7]
35 after lagging.) However, the mortality results, which yield a lower RR estimate, were selected

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1 for the primary analysis to avoid any appearance of “cherry-picking,” and incidence RR
2 estimates were used in sensitivity analyses. A sensitivity analysis was also done using results
3 from Boice et al. (2006) in place of the Zhao et al. (2005) RR estimate. The cohorts for these
4 studies overlap, so they are not independent studies and should not be included in the
5 meta-analysis concurrently. Boice et al. (2006) report results for an overall TCE effect for
6 kidney cancer; however, the results are SMR estimates rather than internal comparisons and are
7 based on fewer exposed deaths (7), so either Zhao et al. (2005) estimate is preferred over the
8 Boice et al. (2006) estimate.

9 Regarding the case-control studies, for Brüning et al. (2003), the results based on
10 self-assessed exposure were preferred because, although TCE exposure was probably under
11 ascertained with this measure, there were greater concerns about the result based on the alternate
12 measure reported—longest-held job in an industry with TCE exposure. Even though this study
13 was conducted in the Arnsberg region of Germany, an area with high prevalence of exposure to
14 TCE, the exposure prevalence in both cases (87%) and controls (79%) seemed inordinately high,
15 and this for not just any job in an industry with TCE exposure, but for the longest-held job.
16 Furthermore, Table V of Brüning et al., which presents this result, states that the result is for
17 longest-held job in industries with TCE *or tetrachloroethylene* exposure. Additionally, some of
18 the industries with exposure to TCE presented in Table V have many jobs that would not entail
19 TCE exposure (e.g., white-collar workers), so the assessment based on industry alone likely has
20 substantial misclassification. Both of these—inclusion of tetrachloroethylene and exposure
21 assessment by industry—could result in overstating TCE exposure prevalence. Results based on
22 the longest-held-job measure were used in a sensitivity analysis.

23 For Charbotel et al. (2006), results from the analysis that considered “only job periods
24 with a good level of confidence for TCE exposure assessment” (Table 7 of Charbotel et al.,
25 2006) were preferred, as these estimates would presumably be less influenced by exposure
26 misclassification. Estimates from the full study analysis were used in a sensitivity analysis. For
27 Pesch et al. (2000), TCE results were presented for 2 different exposure assessments. Estimates
28 using the job-task-exposure-matrix (JTEM) approach were preferred because they seemed to
29 represent a more comprehensive exposure assessment (see Appendix B, Section II-4); estimates
30 based on the JEM approach were used in a sensitivity analysis. Furthermore, results were
31 presented only by exposure category, with no overall RR estimate reported. Case and control
32 numbers for the different exposure categories were kindly provided by Dr. Pesch (personal
33 communication from Baete Pesch, BGFA, to Cheryl Scott, U.S. EPA, 21 February 2008), and we
34 calculated crude overall ORs for males and females combined for each exposure assessment
35 approach.

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1 **C.3.1.2. Results of Meta-Analyses**

2 Results from some of the meta-analyses that were conducted on the epidemiological
3 studies of TCE and kidney cancer are summarized in Table C-8. The pooled estimate from the
4 primary random effects meta-analysis of the 14 studies was 1.25 (95% CI: 1.11, 1.41) (see
5 Figure C-5). As shown in Figure C-5, the analysis was dominated by 2 (contributing almost 70%
6 of the weight) or 3 (almost 80% of the weight) large studies. No single study was overly
7 influential; removal of individual studies resulted in RRp estimates that were all statistically
8 significant (all with $p < 0.005$) and that ranged from 1.22 (with the removal of Brüning) to 1.27
9 (with the removal of Raaschou-Nielsen).

10 Similarly, the RRp estimate was not highly sensitive to alternate RR estimate selections.
11 Use of the 10 alternate selections, individually, resulted in RRp estimates that were all
12 statistically significant (all with $p < 0.002$) and that ranged from 1.19 to 1.27 (see Table C-8). In
13 fact, as can be seen in Table C-8, all but one of the alternates had negligible impact. The Zhao,
14 Axelson, Brüning, and Charbotel original values and alternate selections were associated with
15 very little weight and, thus, have little influence in the RRp. The Raaschou-Nielsen value carried
16 more weight, but the alternate RR estimate was identical to the original, although with a
17 narrower CI, and so did not alter the RRp. Only the Pesch alternate (with the JEM exposure
18 assessment approach instead of the JTEM approach) had much impact, resulting in an RRp
19 estimate of 1.19 (95% CI: 1.07, 1.32). As noted above, the JTEM approach is preferred. The
20 JEM approach takes jobs into account but not tasks; thus, it is expected to have greater potential
21 for exposure misclassification. Indeed, a comparison of exposure prevalences for the
22 two approaches suggests that the JEM approach is less discriminating about exposure; 42% of
23 cases were defined as TCE-exposed under the JEM approach, but only 18% of cases were
24 exposed under the JTEM approach.

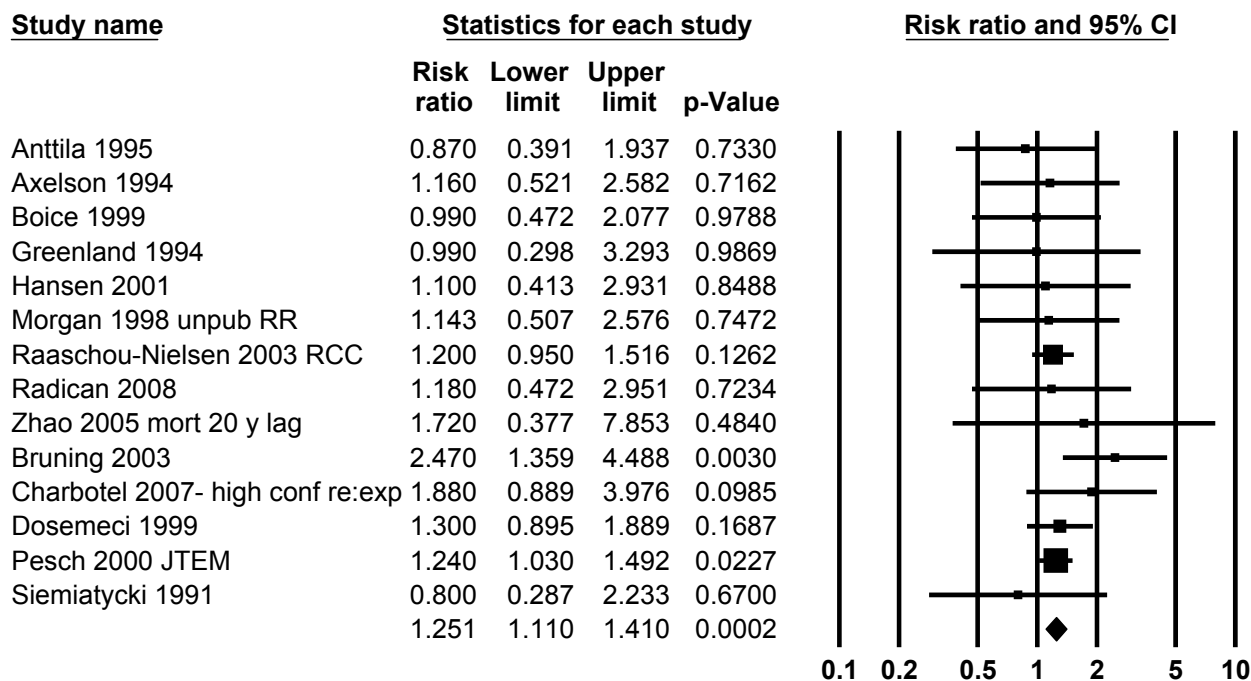
Table C-8. Summary of some meta-analysis results for TCE (overall) and kidney cancer

Analysis	# of studies	Model	Combined RR estimate	95% LCL	95% UCL	Heterogeneity	Comments
All studies	14	Random	1.25	1.11	1.41	None obs	Statistical significance not dependent on single study. No apparent publication bias.
		Fixed	1.25	1.11	1.41		
Cohort	9	Random	1.16	0.96	1.40	None obs	Not significant difference between CC and cohort studies ($p = 0.23$).
		Fixed	1.16	0.96	1.40		Not significant difference between CC and cohort studies ($p = 0.29$).
Case-control	5	Random	1.41	1.08	1.83	Not significant ($p = 0.17$)	
		Fixed	1.32	1.13	1.54		
Alternate RR selections ^a	14	Random	1.25	1.11	1.40–1.41	None obs	With 3 different alternates from Zhao (see Table C-6).
	14	Random	1.27	1.13	1.43	None obs	With Boice (2006) study rather than Zhao
	14	Random	1.25	1.11	1.41	None obs	With estimated female contribution to Axelson.
	14	Random	1.26	1.11	1.41	None obs	With Morgan published SMR.
	14	Random	1.25	1.11	1.40	None obs	With Raaschou-Nielsen all kidney cancer.
	14	Random	1.24	1.10	1.39	None obs	With Brüning longest job held in industry with TCE.
	14	Random	1.25	1.11	1.41	None obs	With Charbotel full study
	14	Random	1.19	1.07	1.32	None obs	With Pesch JEM.
Highest exposure groups	9	Random	1.59	1.26	2.01	None obs	
	12	Random	1.53	1.23	1.91	None obs	Using RR = 1 for Anttila, Axelson, and Hansen (see text). See Table C-10 for alternate RR selection results.

^aChanging the primary analysis by one alternate RR each time.

obs = observable.

TCE and Kidney Cancer



random effects model; same for fixed

Figure C-5. Meta-analysis of kidney cancer and overall TCE exposure.

The pooled estimate is in the bottom row. Symbol sizes reflect relative weights of the studies. The horizontal midpoint of the bottom diamond represents the pooled RR estimate and the horizontal extremes depict the 95% CI limits.

There was no apparent heterogeneity across the 14 studies, i.e., the random effects model and the fixed effect model gave the same results. Nonetheless, subgroup analyses were done examining the cohort and case-control studies separately. With the random effects model (and tau-squared not pooled across subgroups), the resulting RRp estimates were 1.16 (95% CI: 0.96, 1.40) for the cohort studies and 1.41 (1.08, 1.83) for the case-control studies. There was heterogeneity in the case-control subgroup, but it was not statistically significant and the I^2 value

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1 of 38% suggests that the extent of the heterogeneity in this subgroup was low-to-moderate. Nor
2 was the difference between the RRp estimates for the cohort and case-control subgroups
3 statistically significant under either the random effects model or the fixed effect model. Further
4 quantitative investigations of heterogeneity were not pursued because of database limitations
5 and, in any event, there is no evidence for heterogeneity of study results in this database. A
6 qualitative discussion of some potential sources of heterogeneity across studies is nonetheless
7 included in Section C.3.3.

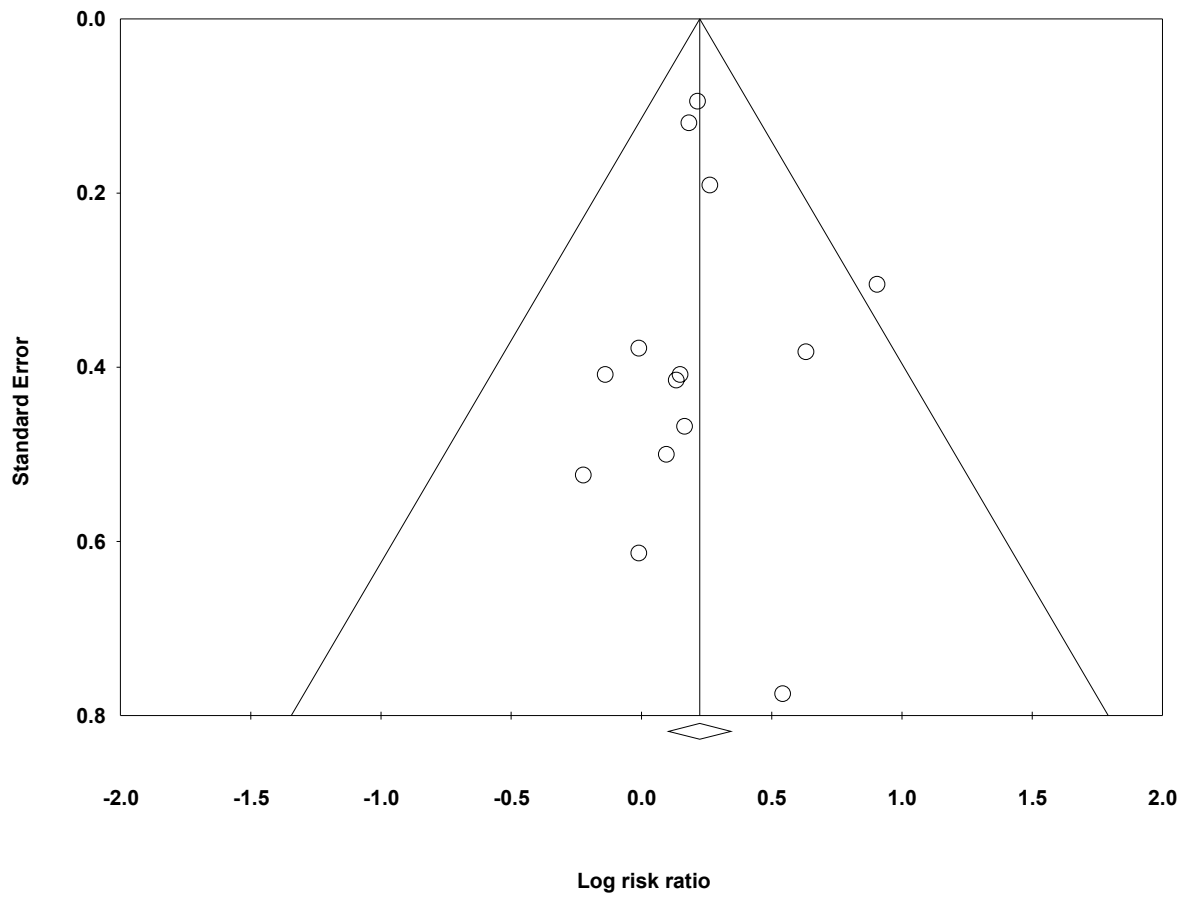
8 As discussed in Section C.1, publication bias was examined in several different ways.
9 The funnel plot in Figure C-6 shows little relationship between RR estimate and study size, and,
10 indeed, none of the other tests performed found any evidence of publication bias. Duval and
11 Tweedie's trim-and-fill procedure, for example, determined that no studies were missing from
12 the funnel plot, i.e., there was no asymmetry to counterbalance. Similarly, the results of a
13 cumulative meta-analysis, incorporating studies with increasing SE one at a time, shows no
14 evidence of a trend of increasing effect size with addition of the less precise studies. Including
15 the 3 most precise studies, reflecting 78% of the weight, the RRp goes from 1.24 to 1.22 to 1.23.
16 The addition of the Brüning study brings the RRp to 1.32 and the weight to 82%. After the
17 addition of the next 5 studies, the RRp stabilizes at about 1.26, and further addition of the 5 least
18 precise studies has little impact.

19 20 **C.3.2. Kidney Cancer Effect in the Highest Exposure Groups**

21 ***C.3.2.1. Selection of RR Estimates***

22 The selected RR estimates for kidney cancer in the highest TCE exposure categories, for
23 studies that provided such estimates, are presented in Table C-9. Five of the 9 cohort studies and
24 4 of the 5 case-control studies reported kidney cancer risk estimates categorized by exposure
25 level. As in Section C.3.1.1 for the overall risk estimates, estimates for RCC were preferentially
26 selected when presented, and, wherever possible, RR estimates for males and females combined
27 were used.

Funnel Plot of Standard Error by Log risk ratio



1
2
3

Figure C-6. Funnel plot of SE by log RR estimate for TCE and kidney cancer studies

Table C-9. Selected RR estimates for kidney cancer risk in highest TCE exposure groups

Study	RR	95% LCL	95% UCL	Exposure category	log RR	SE(log RR)	Alternate RR estimates	Comments
Anttila et al., 1995				100+ $\mu\text{mol/L}$ U-TCA ^a			1.0 assumed	Reported high exposure group results for some cancer sites but not kidney.
Axelson et al., 1994				≥ 2 yr exposure and 100+ mg/L U-TCA			1.0 assumed	Reported high exposure group results for some cancer sites but not kidney.
Boice et al., 1999	0.69	0.22	2.12	≥ 5 yr exp	-0.371	0.578	None	Mortality RR. ICD-9 189.0–189.2. For potential routine or intermittent exposure. adjusted for date of birth, dates 1 st and last employed, race, and sex. Referent group is workers not exposed to any solvent.
Hansen et al., 2001				≥ 1080 mos \times mg/m ³			1.0 assumed	Reported high exposure group results for some cancer sites but not kidney.
Morgan et al., 1998	1.59	0.68	3.71	High cumulative exposure score	0.464	0.433	1.89 (0.85, 4.23) for med/high peak vs. low/no	Mortality RR. ICD-9 189.0–189.2. Adjusted for age and sex.
Raaschou-Nielsen et al., 2003	1.7	1.1	2.4	≥ 5 yrs in subcohort with expected higher exposure levels	0.531	0.183	1.4 (0.99, 1.9) ICD-7 180 ≥ 5 yrs in total cohort	SIR. RCC.
Radican et al., 2008	1.11	0.35	3.49	>25 unit-yr	0.104	0.582	Blair et al. (1998) incidence RR 0.9 (0.3, 3.2)	Mortality HR. ICD-8, -9 189.0, ICD-10 C64. Male and female results presented separately and combined (see text). Referent group is workers with no chemical exposures.

Table C-9. Selected RR estimates for kidney cancer risk in highest TCE exposure groups (continued)

Study	RR	95% LCL	95% UCL	Exposure category	log RR	SE(log RR)	Alternate RR estimates	Comments
Zhao et al., 2005	7.40	0.47	116	High exposure score	2.00	1.41	Mortality RR: 1.82 (0.09, 38.6) Incidence RR no lag: 7.71 (0.65, 91.4) Mortality RR no lag: 0.96 (0.09, 9.91) Boice 2006 mortality RR: 2.12 (0.63, 7.11) for ≥ 5 yrs as test stand mechanic; 3.13 (0.74, 13.2) for ≥ 4 test-yr engine flush	Incidence RR. ICD-9 189. Males only. Adjusted for age, SES, time since first employment, exposure to other carcinogens. 20-yr lag. Incidence results reflect more exposed cases (4 with no lag) than do mortality results (3), so they are used in primary analysis.
Brüning et al., 2003	2.69	0.84	8.66	≥ 20 yrs self-assessed exposure	0.990	0.595	None	Incidence OR. RCC. Adjusted for age, sex, and smoking.
Charbotel et al., 2006	3.34	1.27	8.74	High cumulative dose	1.21	0.492	3.80 (1.27, 11.40) for high cum + peaks 1.96 (0.71, 5.37) for high cum + peaks in full study 2.63 (0.79, 8.83) for high cum in full study	Incidence OR. RCC. In subgroup with good level of confidence for TCE exposure. Adjusted for smoking and body mass index. Matched on sex and age. Alternate full study estimates were additionally adjusted for exposure to cutting fluids and other petroleum oils.

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Table C-9. Selected RR estimates for kidney cancer risk in highest TCE exposure groups (continued)

Study	RR	95% LCL	95% UCL	Exposure category	log RR	SE(log RR)	Alternate RR estimates	Comments
Pesch et al., 2000	1.4	0.9	2.1	Substantial	0.336	0.219	1.2 (0.9, 1.7) for JEM	Incidence OR. RCC. JTEM approach. Adjusted for age, study center, and smoking. Sexes combined.
Siemiatycki 1991	0.8	0.2	3.4	Substantial	-0.233	0.736	none	Incidence OR. Kidney cancer. SE and 95% CI calculated from reported 90% CIs. Males only; adjusted for age, income, and cigarette smoking index.

^aMean personal trichloroacetic acid in urine. 1 $\mu\text{mol/L}$ = 0.1634 mg/L.

1 Three of the 9 cohort studies (Anttila et al., 1995; Axelson et al., 1994; Hansen et al.,
2 2001) did not report kidney cancer risk estimates categorized by exposure level even though
3 these same studies reported such estimates for selected other cancer sites. To address this
4 reporting bias, attempts were made to obtain the results from the primary investigators, and,
5 failing that, an alternate analysis was performed in which null estimates (RR = 1.0) were
6 included for all 3 studies. This alternate analysis was then used as the main analysis, e.g., the
7 basis of comparison for the sensitivity analyses. For the SE (of the logRR) estimates for these
8 null estimates, SE estimates from other sites for which highest-exposure-group results were
9 available were used. For Anttila et al. (1995), the SE estimate for liver cancer in the highest
10 exposure group was used, because liver cancer and kidney cancer had similar numbers of cases
11 in the overall study (5 and 6, respectively). For Axelson et al. (1994), the SE estimate for NHL
12 in the highest exposure group was used, because NHL and kidney cancer had similar numbers of
13 cases in the overall study (5 and 6, respectively). For Hansen et al. (2001), the SE estimate for
14 NHL in the highest exposure group was used, because NHL was the only cancer site of interest
15 in this assessment for which highest-exposure-group results were available.

16 For Boice et al. (1999), only results for workers with “any potential exposure” (rather
17 than “potential routine exposure”) were presented by exposure category, and the referent group is
18 workers not exposed to any solvent. For Morgan et al. (1998), the primary analysis used results
19 for the cumulative exposure metric, and a sensitivity analysis was done with the results for the
20 peak exposure metric.

21 For Radican et al. (2008), it should be noted that the referent group is workers with no
22 chemical exposures, not just no TCE exposure. In addition, exposure group results were
23 reported separately for males and females and were combined for this assessment using
24 inverse-variance weighting, as in a fixed effect meta-analysis. Radican et al. (2008) present only
25 mortality HR estimates by exposure group; however, in an earlier follow-up of this same cohort,
26 Blair et al. (1998) present both incidence and mortality RR estimates by exposure group. The
27 mortality RR estimate based on the more recent follow-up of Radican et al. (2008) (6 deaths in
28 the highest exposure group) was used in the primary analysis, while the incidence RR estimate
29 based on similarly combined results from Blair et al. (1998) (4 cases) was used as an alternate
30 estimate in a sensitivity analysis.

31 Zhao et al. (2005) present kidney cancer RR estimates adjusted for exposure to other
32 carcinogens, because, unlike for lymphoma, this adjustment made a considerable difference.
33 Estimates of RR with this additional adjustment were selected over those without. Furthermore,
34 the kidney results were presented with and without a 20-year lag. A 20-year lag seemed
35 reasonable for kidney cancer, so the lagged estimates were preferred to the unlagged; unlagged

1 estimates were used in sensitivity analyses. In addition, the incidence results reflect more cases
2 (4 with no lag) in the highest exposure group than do the mortality results (3), so the incidence
3 result (with the 20-year lag) was used for the primary analysis, and the unlagged incidence result
4 and the mortality results were used in a sensitivity analysis. Sensitivity analyses were also done
5 using results from Boice et al. (2006) in place of the Zhao et al. (2005) RR estimate. The cohorts
6 for these studies overlap, so they are not independent studies. Boice et al. (2006) report
7 mortality RR estimates for kidney cancer by years worked as a test stand mechanic, a job with
8 potential TCE exposure, and by a measure that weighted years with potential exposure from
9 engine flushing by the number of flushes each year. No results were presented for a third metric,
10 years worked with potential exposure to any TCE, because the Cox proportional hazards model
11 did not converge. The Boice et al. (2006) estimates are adjusted for years of birth and hire and
12 for hydrazine exposure.

13 For Charbotel et al. (2006), results from the analysis that considered “only job periods
14 with a good level of confidence for TCE exposure assessment” (Table 7 of Charbotel et al.,
15 2006) were preferred, as these estimates would presumably be less influenced by exposure
16 misclassification. Estimates from the full study analysis, additionally adjusted for exposure to
17 cutting fluids and other petroleum oils, were used in a sensitivity analysis. Additionally, the high
18 cumulative dose results were preferred, but the results for high cumulative dose + peaks were
19 included in sensitivity analyses. For Pesch et al. (2000), TCE results were presented for
20 two different exposure assessments. As discussed above, estimates using the JTEM approach
21 were preferred because they seemed to represent a more comprehensive exposure assessment;
22 estimates based on the JEM approach were used in a sensitivity analysis.

23

24 **C.3.2.2. Results of Meta-Analyses**

25 Results from the meta-analyses that were conducted for kidney cancer in the highest
26 exposure groups are summarized at the bottom of Table C-8 and reported in more detail in
27 Table C-10. The pooled RR estimate from the random effects meta-analysis of the 9 studies with
28 results presented for exposure groups was 1.59 (95% CI: 1.26, 2.01) (see Figure C-7). The RR_p
29 estimate from the primary random effects meta-analysis with null RR estimates (i.e., 1.0)
30 included for Anttila, Axelson, and Hansen to address reporting bias (see above) was 1.53
31 (1.23, 1.91) (see Figure C-8). The inclusion of these 3 additional studies contributed just under
32 8% of the total weight. As with the overall kidney cancer meta-analyses, the meta-analyses of
33 the highest-exposure groups were dominated by 2 studies (Raaschou-Nielsen and Pesch), which
34 provided about 66% of the weight. No single study was overly influential; removal of individual
35 studies resulted in RR_p estimates that were all statistically significant (all with $p < 0.02$) and that

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1 ranged from 1.43 (with the removal of Raaschou-Nielsen) to 1.58 (with the removal of Boice
2 [1999] or Pesch).

3 Similarly, the RRp estimate was not highly sensitive to alternate RR estimate selections.
4 Use of the 12 alternate selections, individually, resulted in RRp estimates that were all
5 statistically significant (all with $p < 0.002$) and that ranged from 1.42 to 1.55, with all but 2 of
6 the alternate selections yielding RRp estimates in the narrow range of 1.49–1.55 (see
7 Table C-10). The lowest RRp estimates, 1.42 in both cases, were obtained when the alternate
8 selections involved the 2 large studies. One of the alternate selections was for Raaschou-
9 Nielsen, with a highest-exposure group estimate for all kidney cancer in the total cohort, rather
10 than RCC in the subcohort expected to have higher exposure levels. The latter value is strongly
11 preferred because, as discussed above, the subcohort is likely to have less exposure
12 misclassification. Furthermore, RCC is very different from other types of kidney cancer, and
13 TCE, if an etiological factor, may not be etiologically associated with all kidney cancers, so
14 using the broad category may dilute a true association with RCC, if one exists. The other
15 alternate selection with a considerable impact on the RRp estimate was for Pesch, with the
16 highest exposure group result based on the JEM exposure assessment approach, rather than the
17 JTEM approach. As discussed above, the JTEM approach is preferred because it seemed to be a
18 more comprehensive and discriminating approach, taking actual job tasks into account, rather
19 than just larger job categories. Thus, although results with these alternate selections are
20 presented for comprehensiveness and transparency, the primary analysis is believed to reflect
21 better the potential association between kidney cancer (in particular, RCC) and TCE exposure.

22 There was no observable heterogeneity across the studies for any of the meta-analyses
23 conducted with the highest-exposure groups, including those in which RR values for Anttila,
24 Axelson, and Hansen were assumed. No subgroup analyses (e.g., cohort vs. case-control studies)
25 were done with the highest exposure group results.

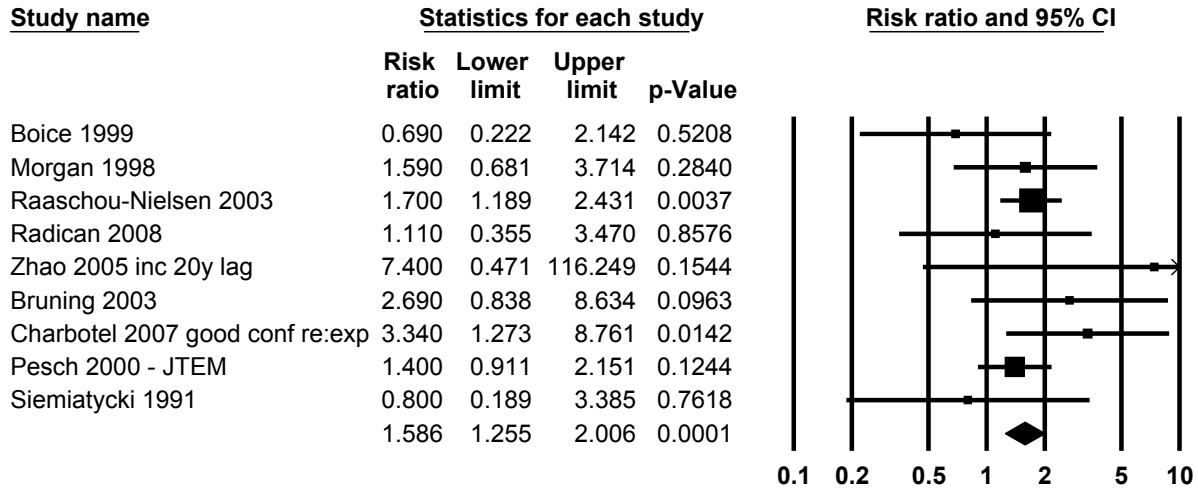
Table C-10. Summary of some meta-analysis results for TCE (highest exposure groups) and kidney cancer

Analysis	Model	Combined RR estimate	95% LCL	95% UCL	Heterogeneity	Comments
Analysis based on reported results	Random	1.59	1.26	2.01	None obs (fixed = random)	
Primary analysis	Random	1.53	1.23	1.91	None obs	Includes assumed values for Anttila, Axelson, and Hansen (see text). Statistical significance not dependent on single study.
Alternate RR selections ^a	Random	1.52	1.22	1.90	None obs	With Blair et al. (1998) incidence RR instead of Radican mortality HR.
	Random	1.55	1.24	1.94	None obs	With Morgan peak metric.
	Random	1.42	1.15	1.75	None obs	With Raaschou-Nielsen for all kidney cancer ≥ 5 yrs in total cohort.
	Random	1.51–1.54	1.21–1.23	1.89–1.92	None obs	With Zhao incidence unlagged and mortality with and without lag.
	Random	1.53–1.54	1.23–1.24	1.91–1.92	None obs	With Boice (2006) alternates for Zhao (see text).
	Random	1.49–1.52	1.19–1.22	1.86–1.91	None obs	With Charbotel high cumulative dose + peaks in subgroup; and high cumulative dose and high cumulative dose + peaks in full study additionally adjusted for exposure to cutting fluids and other petroleum oils..
	Random	1.42	1.16	1.74	None obs	With Pesch JEM.

^aChanging the primary analysis by one alternate RR each time.

obs = observable.

TCE and Kidney Cancer - highest exposure groups

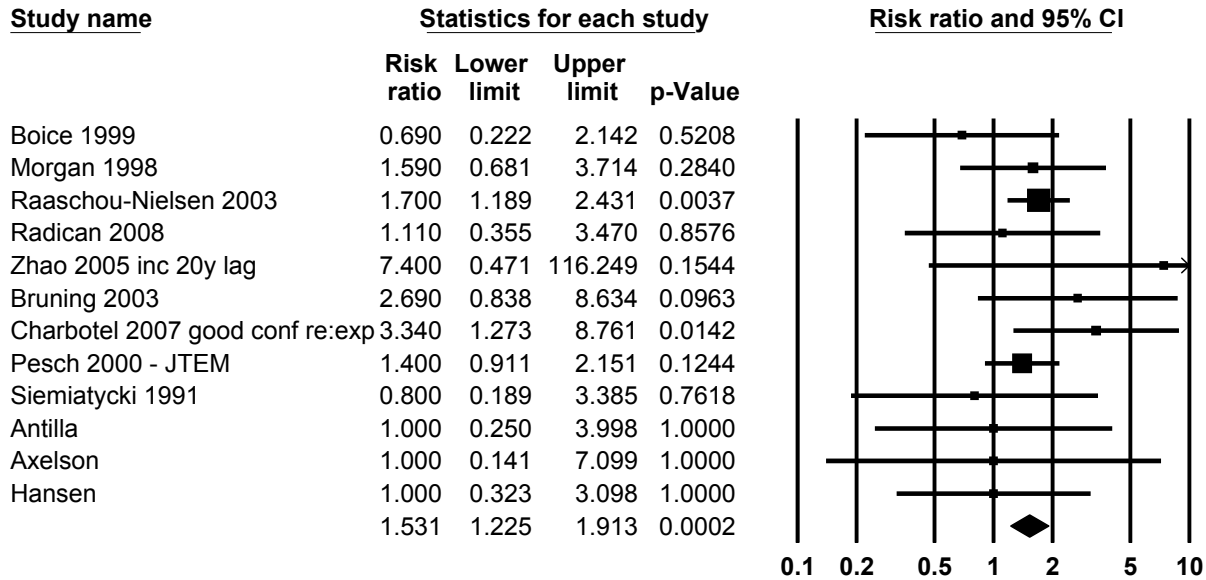


random effects model

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Figure C-7. Meta-analysis of kidney cancer and TCE exposure—highest exposure groups. The pooled estimate is in the bottom row. Symbol sizes reflect relative weights of the studies. The horizontal midpoint of the bottom diamond represents the pooled RR estimate and the horizontal extremes depict the 95% CI limits.

TCE and Kidney Cancer - highest exposure groups



random effects model; same for fixed

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2 **Figure C-8. Meta-analysis of kidney cancer and TCE exposure—highest**
3 **exposure groups, with assumed null RR estimates for Anttila, Axelson, and**
4 **Hansen (see text).**

7 C.3.3. Discussion of Kidney Cancer Meta-Analysis Results

8 For the most part, the meta-analyses of the overall effect of TCE exposure on kidney
9 cancer suggest a small, statistically significant increase in risk. The pooled estimate from the
10 primary random effects meta-analysis of the 14 studies was 1.25 (95% CI: 1.11, 1.41). Although
11 the analysis was dominated by 2–3 large studies that contribute 70–80% of the weight, the
12 pooled estimate was not overly influenced by any single study, nor was it overly sensitive to
13 individual RR estimate selections. The largest downward impacts were from the removal of the
14 Brüning study, resulting in an RR_p estimate of 1.22 (95% CI: 1.08, 1.37), and from the
15 substitution of the Pesch JTEM RR estimate with the RR estimate based on the JEM approach,
16 resulting in an RR_p estimate of 1.19 (1.07, 1.32). Thus, the finding of an increased risk of

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1 kidney cancer associated with TCE exposure is robust. Furthermore, there is no evidence of
2 publication bias in this data set.

3 In addition, there was no heterogeneity observed across the results of the 14 studies.
4 When subgroup analyses were done of cohort and case-control studies separately, there was
5 some observable heterogeneity among the case-control studies, but it was not statistically
6 significant ($p = 0.17$) and the I^2 value of 38% suggested the extent of the heterogeneity was low-
7 to-moderate. The increased risk of kidney cancer was strengthened in the case-control study
8 analysis and weakened in the cohort study analysis, but the difference between the 2 RRp
9 estimates was not statistically significant. One difference between the case-control and cohort
10 studies is that the case-control studies were of RCC and almost all of the cohort studies were of
11 all kidney cancers, including renal pelvis. As discussed above, RCC is very different from other
12 types of kidney cancer, and TCE, if an etiological factor, may not be etiologically associated
13 with all kidney cancers, so using the broad category may dilute a true association with RCC, if
14 one exists.

15 With respect to the nonsignificant heterogeneity in the 5 case-control studies, these
16 studies differ in TCE exposure potential to the underlying population from which case and
17 control subjects were identified, and this may be a source of some heterogeneity. Prevalence of
18 exposure to TCE among cases in these studies was 27% in Charbotel et al. (2006) (for
19 high-level-of-confidence jobs), 18% in Brüning et al. (2003) (for self-assessed exposure), 18% in
20 Pesch et al. (2000), 13% in Dosemeci et al. (1999) and 1% in Siemiatycki (1991). Both Brüning
21 et al. (2003) and Charbotel et al. (2006) are studies designed specifically to assess RCC and TCE
22 exposure. These studies were carried out in geographical areas with both a high prevalence and
23 a high degree of TCE exposure. Some information is provided in these and accompanying
24 papers to describe the nature of exposure, making it possible to estimate the order of magnitude
25 of exposure, even though there were no direct measurements (Cherrie et al., 2001; Brüning et al.,
26 2003; Fevotte et al., 2006). The Charbotel et al. (2006) study was carried out in the Arve Valley
27 region in France, where TCE exposure was through metal-degreasing activity in small shops
28 involved in the manufacturing of screws and precision metal parts (Fevotte et al., 2006).
29 Industrial hygiene data from shops in this area indicated high intensity TCE exposures of
30 100 ppm or higher, particularly from exposures from hot degreasing processes. Considering
31 exposure only from the jobs with a high level of confidence about exposure, 18% of exposed
32 cases were identified with high cumulative exposure to TCE. The source population in the
33 Brüning et al. (2003) study includes the Arnsberg region in Germany, which also has a high
34 prevalence of TCE exposure. A large number of small companies used TCE in metal degreasing
35 in small workrooms. Subjects in this study also described neurological symptoms previously

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1 associated with higher TCE intensities. While subjects in the Brüning et al. (2003) study had
2 potential high TCE exposure intensity, average TCE exposure in this study is considered lower
3 than that in the Charbotel et al. (2006) study because the base population was enlarged beyond
4 the Arnsberg region to areas which did not have the same focus of industry.

5 Siemiatycki (1991), Dosemeci et al. (1999), and Pesch et al. (2000) are population-based
6 studies. Pesch et al. (2000) includes the Arnsberg area and 4 other regions. Sources of exposure
7 to TCE and other chlorinated solvents are much less well defined, and most subjects identified
8 with TCE exposure probably had minimal contact; estimated average concentrations to exposed
9 subjects were of about 10 ppm or less (NRC, 2006). Neither Dosemeci et al. (1999) nor
10 Siemiatycki (1991) describe the nature of the TCE exposure. TCE exposure potential in these
11 studies is likely lower than in the three other studies and closer to background. Furthermore, the
12 use of generic job-exposure-matrices for exposure assessment in these studies may result in a
13 greater potential for exposure misclassification bias.

14 Nine of the 14 studies categorized results by exposure level. Three other studies reported
15 results for other cancer sites by exposure level, but not kidney cancer; thus, to address this
16 reporting bias, null values (i.e., RR estimates of 1.0) were used for these studies. Different
17 exposure metrics were used in the various studies, and the purpose of combining results across
18 the different highest exposure groups was not to estimate an RRp associated with some level of
19 exposure, but rather to see the impacts of combining RR estimates that should be less affected by
20 exposure misclassification. In other words, the highest exposure category is more likely to
21 represent a greater differential TCE exposure compared to people in the referent group than the
22 exposure differential for the overall (typically any vs. none) exposure comparison. Thus, if TCE
23 exposure increases the risk of kidney cancer, the effects should be more apparent in the highest
24 exposure groups. Indeed, the RRp estimate from the primary meta-analysis of the highest
25 exposure group results was 1.53 (95% CI: 1.23, 1.91), which is greater than the RRp estimate of
26 1.25 (95% CI: 1.11, 1.41) from the overall exposure analysis. This result for the highest
27 exposure groups was not overly influenced by any single study, nor was it overly sensitive to
28 individual RR estimate selections. Heterogeneity was not observed in any of the analyses. The
29 robustness of this finding lends substantial support to a conclusion that TCE exposure increases
30 the risk of kidney cancer.

1 **C.4. META-ANALYSIS FOR LIVER CANCER**

2 **C.4.1. Overall Effect of TCE Exposure**

3 **C.4.1.1. Selection of RR Estimates**

4 The selected RR estimates for liver cancer associated with TCE exposure from the
5 epidemiological studies are presented in Table C-11. There were no case-control studies for
6 liver cancer and TCE exposure that were selected for inclusion in the meta-analysis (see
7 Appendix B, Section II-9), so all of the relevant studies are cohort studies. All of the studies
8 reported results for liver cancers plus cancers of the gall bladder and extrahepatic biliary
9 passages (i.e., ICD-7 155.0 + 155.2; ICD-8 and -9 155 + 156). Three of the studies also report
10 results for liver cancer alone (ICD-7 155.0; ICD-8 and -9 155). For the primary analysis, results
11 for cancers of the liver, gall bladder, and biliary passages combined were selected, for the sake of
12 consistency, since these were reported in all the studies. An alternate analysis was also done
13 using results for liver cancer alone for the 3 studies that reported them and the combined liver
14 cancer results for the remainder of the studies.

15 As for lymphoma and kidney cancer, many of the studies provided RR estimates only for
16 males and females combined, and we are not aware of any basis for a sex difference in the
17 effects of TCE on liver cancer risk; thus, wherever possible, RR estimates for males and females
18 combined were used. The only study of much size (in terms of number of liver cancer cases)
19 that provided results separately by sex was Raaschou-Nielsen (2003). The results of this study
20 suggest that liver cancer risk in females might be slightly higher than the risk in males, but the
21 number of female cases is small (primary liver cancer SIR: males 1.1 [95% CI: 0.74, 1.64;
22 27 cases], females 2.8 [1.13, 5.80; 7 cases]; gallbladder and biliary passage cancers SIR:
23 males 1.1 [0.61, 1.87; 14 cases]; females 2.8 [1.28, 5.34; 9 cases]). Radican et al. (2008) report
24 HRs for liver/biliary passage cancers combined of 1.36 (95% CI: 0.59, 3.11; 28 deaths) for males
25 and 0.74 (95% CI: 0.18, 2.97; 3 deaths) for females, but these results are based on fewer cases,
26 especially in females.

27

Table C-11. Selected RR estimates for liver cancer associated with TCE exposure (overall effect) from cohort studies

Study	RR	95% LCL	95% UCL	RR type	log RR	SE(log RR)	Alternate RR estimates	Comments
Anttila et al., 1995	1.89	0.86	3.59	SIR	0.637	0.333	2.27 (0.74, 5.29) for 155.0 alone	ICD-7 155.0 + 155.1; combined assuming Poisson distribution.
Axelsson et al., 1994	1.41	0.38	3.60	SIR	0.344	0.5	1.34 (0.36, 3.42) with estimated female contribution to SIR added (see text)	ICD-7 155. Results reported for males only, but there was a small female component to the cohort.
Boice et al., 1999	0.54	0.15	1.38	SMR	-0.616	0.5	0.81 (0.45, 1.33) for any potential exposure	ICD-9 155 + 156. For potential routine exposure.
Greenland et al., 1994	0.54	0.11	2.63	OR	-0.616	0.810	None	ICD-8 155 + 156. Nested case-control study.
Hansen et al., 2001	2.1	0.7	5.0	SIR	0.742	0.447	None	ICD-7 155. Male and female results reported separately; combined assuming Poisson distribution.
Morgan et al., 1998	1.48	0.56	3.91	SMR	0.393	0.495	Published SMR 0.98 (0.36, 2.13)	ICD-9 155 + 156. Unpublished RR, adjusted for age and sex (see text).
Raaschou-Nielsen et al., 2003	1.35	1.03	1.77	SIR	0.300	0.138	1.28 (0.89, 1.80) for ICD-7 155.0	ICD-7 155.0 + 155.1. Results for males and females and different liver cancer types reported separately; combined assuming Poisson distribution.
Radican et al., 2008	1.12	0.57	2.19	Mortality HR	0.113	0.343	1.25 (0.31, 4.97) for ICD-8, -9 155.0	ICD-8, -9 155 + 156, ICD-10 C22-C24. Time variable = age; covariates = sex, race. Referent group is workers with no chemical exposures.
Boice et al., 2006	1.28	0.35	3.27	SMR	0.247	0.5	1.0 assumed for Zhao et al. (2005)	ICD-9 155 + 156. Boice et al. (2006) used in lieu of Zhao et al. (2005) because Zhao et al. (2005) do not report liver cancer results. Boice (2006) cohort overlaps Zhao cohort.

1 Most of the selections in Table C-11 should be self-evident, but some are discussed in
2 more detail here, in the order the studies are presented in the table. For Axelson et al. (1994), in
3 which a small subcohort of females was studied but only results for the larger male subcohort
4 were reported, the reported male-only results were used in the primary analysis; however, as for
5 lymphoma and kidney cancer, an attempt was made to estimate the female contribution to an
6 overall RR estimate for both sexes and its impact on the meta-analysis. Axelson et al. (1994)
7 reported that there were no cases of liver cancer observed in females, but the expected number
8 was not presented. To estimate the expected number, the expected number for males was
9 multiplied by the ratio of female-to-male person-years in the study and by the ratio of female-to-
10 male age-adjusted incidence rates for liver cancer. The male results and the estimated female
11 contribution were then combined into an RR estimate for both sexes assuming a Poisson
12 distribution, and this alternate RR estimate for the Axelson et al. (1994) study was used in a
13 sensitivity analysis.

14 For Boice et al. (1999), results for “potential routine exposure” were selected for the
15 primary analysis, because this exposure category was considered to have less exposure
16 misclassification, and results for “any potential exposure” were used in a sensitivity analysis. To
17 estimate the SE(logRR) for the alternate RR selection, it was assumed that the number of
18 exposed cases (deaths) was 15. The actual number was not presented, but 15 was the number
19 that allowed us to reproduce the reported CIs. The number suggested by exposure level in Boice
20 et al. (1999) Table 9 is 13; however, it may be that exposure level data were not available for all
21 the cases. In their published paper, Morgan et al. (1998) present only SMRs for overall TCE
22 exposure, although the results from internal analyses are presented for exposure subgroups. RR
23 estimates for overall TCE exposure from the internal analyses of the Morgan et al. (1998) cohort
24 data were available from an unpublished report (Environmental Health Strategies, 1997); from
25 these, the RR estimate from the Cox model which included age and sex was selected, because
26 those are the variables deemed to be important in the published paper. The internal analysis RR
27 estimate was preferred for the primary analysis, and the published SMR result was used in a
28 sensitivity analysis.

29 Raaschou-Nielsen et al. (2003) reported results for primary liver cancer (ICD-7 155.0),
30 gallbladder and biliary passage cancers (ICD-7 155.1), and unspecified liver cancers (ICD-7 156)
31 separately. As discussed above, RR estimates for cancers of the liver, gall bladder, and biliary
32 passages combined were preferred for the primary analysis; thus, the results for primary liver
33 cancer and gallbladder/biliary passage cancers were combined (across sexes as well), assuming a
34 Poisson distribution. The results for primary liver cancer only (similarly combined across sexes)
35 were used in an alternate analysis. The results for unspecified liver cancers (ICD-7 156) were

1 not included in any analyses because, under the ICD-7 coding, 156 can include secondary liver
2 cancers. For Radican et al. (2008), the Cox model hazard ratio (HR) from the 2000 follow-up
3 was used. In the Radican et al. (2008) Cox regressions, age was the time variable, and sex and
4 race were covariates. It should also be noted that the referent group is composed of workers with
5 no chemical exposures, not just no exposure to TCE.

6 Zhao et al. (2005) did not present RR estimates for liver cancer; thus, results from Boice
7 et al. (2006) were used in the primary analysis. The cohorts for these studies overlap, so they are
8 not independent studies. Zhao et al. (2005), however, was our preferred study for lymphoma and
9 kidney cancer results; thus, in a sensitivity analysis, a null value (RR = 1.0) was assumed for
10 Zhao et al. (2005) to address the potential reporting bias. The SE estimate for kidney cancer
11 (incidence with 0 lag) was used as the SE for the liver cancer. (It is not certain that there was a
12 reporting bias in this case. In the “Methods” section of their paper, Zhao et al. [2005] list the
13 cancer sites examined in the cohort, and liver was not listed; it is not clear if the list of sites was
14 determined *a priori* or *post hoc*.) Also, on the issue of potential reporting bias, the Siemiatycki
15 (1991) study should be mentioned. This study was a case-control study for multiple cancer sites,
16 but only the more common sites, in order to have greater statistical power. Thus, NHL and
17 kidney cancer results were available, but not liver cancer results. Because no liver results were
18 presented for any of the chemicals, this is not a case of reporting bias.

19 20 **C.4.1.2. Results of Meta-Analyses**

21 Results from some of the meta-analyses that were conducted on the epidemiological
22 studies of TCE and liver cancer are summarized in Table C-12. The pooled estimate from the
23 primary random effects meta-analysis of the 9 studies was 1.33 (95% CI: 1.09, 1.64) (see
24 Figure C-9). As shown in Figure C-9, the analysis was dominated by one large study
25 (contributing about 57% of the weight). That large study was critical in terms of statistical
26 significance of the RRp estimate. Without the large Raaschou-Nielsen study, the RRp estimate
27 does not change noticeably, but it is no longer statistically significant (RRp = 1.31; 95% CI:
28 0.96, 1.79). No other single study was overly influential; removal of any of the other individual
29 studies resulted in RRp estimates that were all statistically significant and that ranged from 1.29
30 (with the removal of Anttila) to 1.39 (with the removal of Boice [1999]).

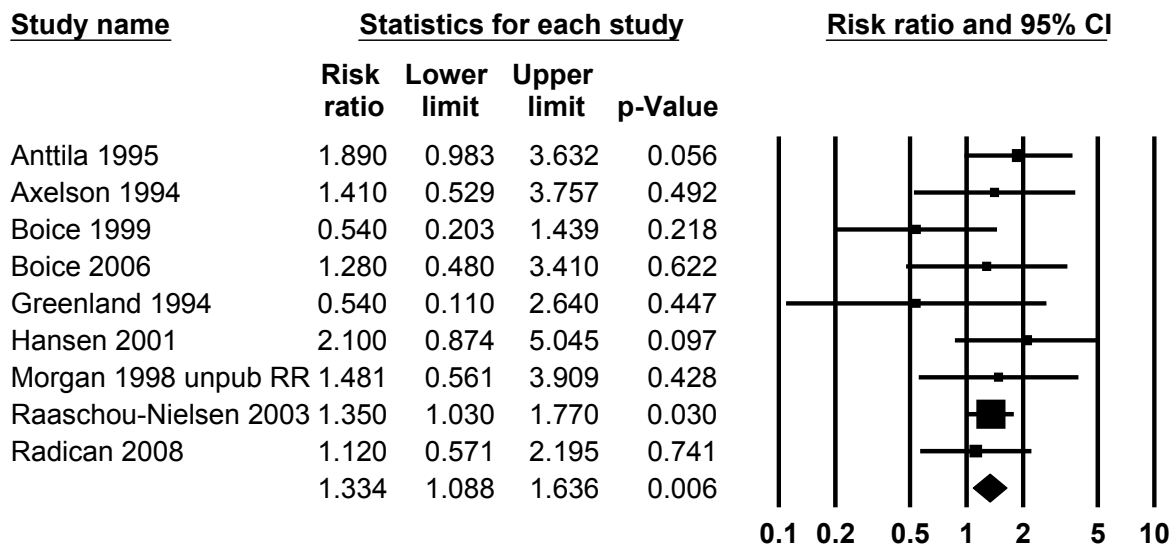
Table C-12. Summary of some meta-analysis results for TCE and liver cancer

Analysis	# of studies	Model	Combined RR estimate	95% LCL	95% UCL	Heterogeneity	Comments
All studies (all cohort studies)	9	Random	1.33	1.09	1.64	None obs (fixed = random)	Statistical significance not dependent on single study, except for Raaschou-Nielsen, without which $p = 0.08$. No apparent publication bias.
		Fixed	1.33	1.09	1.64		
All studies; liver cancer only, when available	9	Random	1.31	1.02	1.67	None obs	Used RR estimates for liver cancer alone for the 3 studies that presented these; remaining RR estimates are for liver and gall bladder/biliary passage cancers.
Alternate RR selections ^a	9	Random	1.33	1.08	1.63	None obs	With 1.0 assumed for Zhao in lieu of Boice (2006) (see text).
	9	Random	1.29	1.06	1.56	None obs	With Boice (1999) any potential exposure rather than potential routine exposure.
	9	Random	1.33	1.09	1.63	None obs	With estimated female contribution to Axelson.
	9	Random	1.30	1.07	1.59	None obs	With Morgan published SMR.
Highest exposure groups	6	Random	1.32	0.93	1.86	None obs	
	8	Random	1.28	0.93	1.77	None obs	Primary analysis. Using RR = 1 for Hansen and Zhao (see text).
	7–8	Random	1.24–1.26	0.88–0.91	1.73–1.82	None obs	Using alternate selections for Morgan and Raaschou-Nielsen and excluding Axelson. ^a

^aChanging the primary analysis by one alternate RR each time.

obs = observable.

TCE and Liver Cancer



random effects model; same for fixed

Figure C-9. Meta-analysis of liver cancer and TCE exposure. The pooled estimate is in the bottom row. Symbol sizes reflect relative weights of the studies. The horizontal midpoint of the bottom diamond represents the pooled RR estimate and the horizontal extremes depict the 95% CI limits.

As discussed in Section C.4.1.1, all of the 9 studies presented results for liver and gall bladder/biliary passage cancers combined, and these results were the basis for the primary analysis discussed above. An alternate analysis was performed substituting, simultaneously, results for liver cancer alone for the 3 studies for which these were available. The RRp estimate from this analysis was slightly lower than the one based entirely on results from the combined cancer categories (1.31; 95% CI: 1.02, 1.67). This result was driven by the fact that the RR estimate from the large Raaschou-Nielsen et al. (2003) study decreased from 1.35 for liver and gall bladder/biliary passage cancers combined to 1.28 for liver cancer alone.

Similarly, the RRp estimate was not highly sensitive to other alternate RR estimate selections. Use of the 4 other alternate selections, individually, resulted in RRp estimates that were all statistically significant (all with $p < 0.02$) and that ranged from 1.29 to 1.33 (see

1 Table C-12). In fact, as can be seen in Table C-12, only one of the alternates had notable impact.
2 The Boice (2006), Zhao, and Axelson original values and alternate selections were associated
3 with very little weight and, thus, have little influence in the RRp. Using the Boice (1999)
4 alternate RR estimate based on any potential exposure rather than potential routine exposure
5 decreased the RRp slightly from 1.33 to 1.29. The alternate Boice (1999) RR estimate is actually
6 larger than the original value (0.81 vs. 0.54); however, use of the less discriminating exposure
7 metric captures more liver cancer deaths, causing the weight of that study to increase from about
8 4.3% to almost 15%.

9 There was no apparent heterogeneity across the nine studies, i.e., the random effects
10 model and the fixed effect model gave the same results. Furthermore, all of the liver cancer
11 studies were cohort studies, so no subgroup analyses examining cohort and case-control studies
12 separately, as was done for lymphoma and kidney cancer, were conducted. No alternate
13 quantitative investigations of heterogeneity were pursued because of database limitations and, in
14 any event, there is no evidence for heterogeneity of study results in this database.

15 As discussed in Section C.1, publication bias was examined in several different ways.
16 The funnel plot in Figure C-10 shows little relationship between RR estimate and study size, and,
17 indeed, none of the other tests performed found any evidence of publication bias. Duval and
18 Tweedie's trim-and-fill procedure, for example, suggested that no studies were missing from the
19 funnel plot, i.e., there was no asymmetry to counterbalance. Similarly, the results of a
20 cumulative meta-analysis, incorporating studies with increasing SE one at a time, shows no
21 evidence of a trend of increasing effect size with addition of the less precise studies. The
22 Raaschou-Nielsen study contributes about 57% of the weight. Including the 2 next most precise
23 studies, the RRp goes from 1.35 to 1.42 to 1.38 and the weight to 76%. With the addition of
24 each of the next 3 most precise studies, the RRp estimate is 1.42. Further addition of the 3 least
25 precise studies gradually brings the RRp back down to 1.33. Thus, if anything, the evidence is
26 somewhat suggestive of an *inverse* relationship between SE and effect size, contrary to what
27 would be expected if publication bias were occurring.

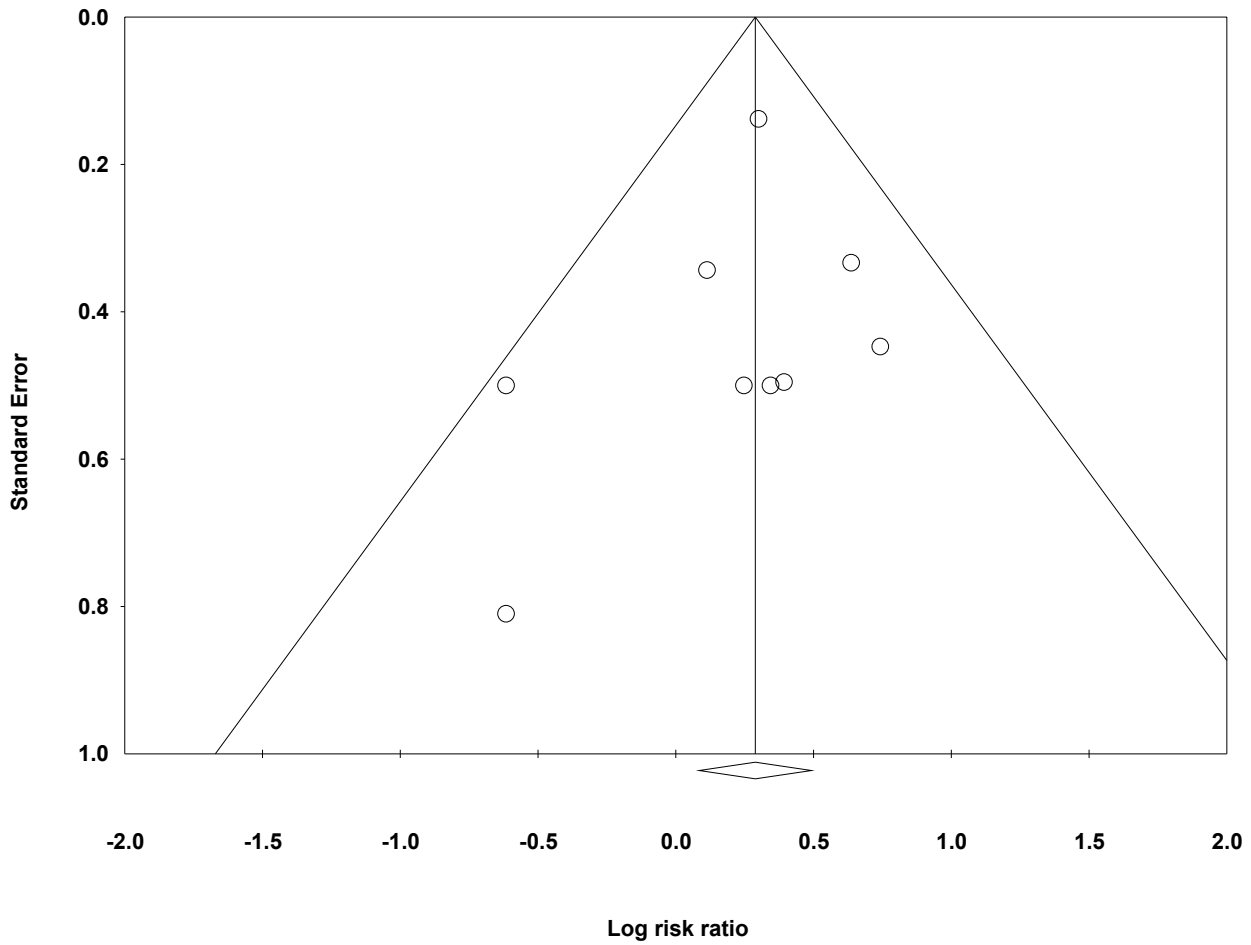
28 **C.4.2. Liver Cancer Effect in the Highest Exposure Groups**

29 **C.4.2.1. Selection of RR Estimates**

30 The selected RR estimates for liver cancer in the highest TCE exposure categories, for
31 studies that provided such estimates, are presented in Table C-13. Six of the 9 cohort studies
32 reported liver cancer risk estimates categorized by exposure level. As in Section C.4.1.1 for the
33 overall risk estimates, estimates for cancers of the liver and gall bladder/biliary passages
34 combined were preferentially selected, when presented, for the sake of consistency, and,
35 wherever possible, RR estimates for males and females combined were used.

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Funnel Plot of Standard Error by Log risk ratio



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Figure C-10. Funnel plot of SE by log RR estimate for TCE and liver cancer studies.

Table C-13. Selected RR estimates for liver cancer risk in highest TCE exposure groups

Study	RR	95% LCL	95% UCL	Exposure category	log RR	SE(log RR)	Alternate RR estimates	Comments
Anttila et al., 1995	2.74	0.33	9.88	100+ $\mu\text{mol/L}$ U-TCA ^a	1.008	0.707		SIR. ICD-7 155.0 (liver only).
Axelson et al., 1994	3.7	0.09	21	100+ mg/L U-TCA	1.308	1.000	Exclude study	SIR. ICD-7 155. 0 cases observed in highest exposure group (i.e., ≥ 2 y and 100+ U-TCA), so combined with <2 y and 100+ subgroup and females, estimating the expected numbers (see text).
Boice et al., 1999	0.94	0.36	2.46	≥ 5 yr exposure	-0.062	0.490	None	Mortality RR. ICD-9 155 + 156. For potential routine or intermittent exposure. Adjusted for date of birth, dates 1 st and last employed, race, and sex. Referent group is workers not exposed to any solvent.
Hansen et al., 2001				≥ 1080 mos \times mg/m ³			1.0 assumed	Reported high exposure group results for some cancer sites but not liver.
Morgan et al., 1998	1.19	0.34	4.16	High cumulative exposure score	0.174	0.639	0.98 (0.29, 3.35) for med/high peak vs. low/no	Mortality RR. ICD-9 155 + 156. Adjusted for age and sex.
Raaschou-Nielsen et al., 2003	1.2	0.7	1.9	≥ 5 yrs	0.182	0.243	1.1 (0.5, 2.1) ICD-7 155.0 (liver only)	SIR. ICD-7 155.0 + 155.1. Male and female results presented separately and combined assuming a Poisson distribution.
Radican et al., 2008	1.49	0.67	3.34	> 25 unit-yr	0.399	0.411	None (see text)	Mortality HR. ICD-8, -9 155 + 156, ICD-10 C22-C24. Male and female results presented separately and combined (see text). Time variable = age, covariate = race. Referent group is workers with no chemical exposures.
Zhao et al., 2005				High exposure score			1.0 assumed	No liver results reported.

^aMean personal trichloroacetic acid in urine. 1 $\mu\text{mol/L}$ = 0.1634 mg/L.

1 Two of the 9 cohort studies (Hansen et al., 2001; Zhao et al., 2005) did not report liver
2 cancer risk estimates categorized by exposure level even though these same studies reported such
3 estimates for selected other cancer sites. To address this reporting bias (as discussed above,
4 Zhao et al. [2005] did not present any liver results, and it is not clear if this was actual reporting
5 bias or an *a priori* decision not to examine liver cancer in the cohort.), attempts were made to
6 obtain the results from the primary investigators, and, failing that, alternate analyses were
7 performed in which null estimates (RR = 1.0) were included for both studies. This alternate
8 analysis was then used as the main analysis, e.g., the basis of comparison for the sensitivity
9 analyses. For the SE (of the logRR) estimates for the null estimates, SE estimates from other
10 sites for which highest-exposure-group results were available were used. For Hansen et al.
11 (2001), the SE estimate for NHL in the highest exposure group was used, because NHL was the
12 only cancer site of interest in this assessment for which highest-exposure-group results were
13 available. For Zhao et al. (2005), the SE estimate for kidney cancer in the highest-exposure
14 group (incidence with 0 lag) was used. (Note that Boice et al. [2006], who studied a cohort that
15 overlapped that of Zhao et al. [2005], also did not present liver cancer results by exposure level.)

16 For Axelson et al. (1994), there were no liver cancer cases in the highest exposure group
17 (≥ 2 years and 100+ mean urinary-trichloroacetic acid [U-TCA] level), so no log RR and
18 SE(log RR) estimates were available for the meta-analysis. Instead, the < 2 years and ≥ 2 years
19 results were combined, assuming expected numbers of cases were proportional to person-years,
20 and 100+ U-TCA (with any exposure duration) was used as the highest exposure category. The
21 female contribution to the expected number was also estimated, again assuming proportionality
22 to person-years, and adjusting for the difference between female and male age-adjusted liver
23 cancer incidence rates. The estimated RR and SE values for the combined exposure times and
24 sexes were used in the primary analysis. In an alternate analysis, the Axelson et al. (1994) study
25 was excluded altogether, because we estimated that less than 0.2 cases were expected in the
26 highest-exposure category, suggesting that the study had low power to detect an effect in the
27 highest-exposure group and would contribute little weight to the meta-analysis.

28 For Boice et al. (1999), only results for workers with “any potential exposure” (rather
29 than “potential routine exposure”) were presented by exposure category, and the referent group is
30 workers not exposed to any solvent. For Morgan et al. (1998), the primary analysis used results
31 for the cumulative exposure metric, and a sensitivity analysis was done with the results for the
32 peak exposure metric. For Raaschou-Nielsen et al. (2003), unlike for NHL and RCC, liver
33 cancer results for the subcohort with expected higher exposure levels were not presented, so the
34 only highest-exposure group results were for duration of employment in the total cohort. Results

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1 for cancers of the liver and gall bladder/biliary passages combined were used for the primary
2 analysis and results for liver cancer alone in a sensitivity analysis.

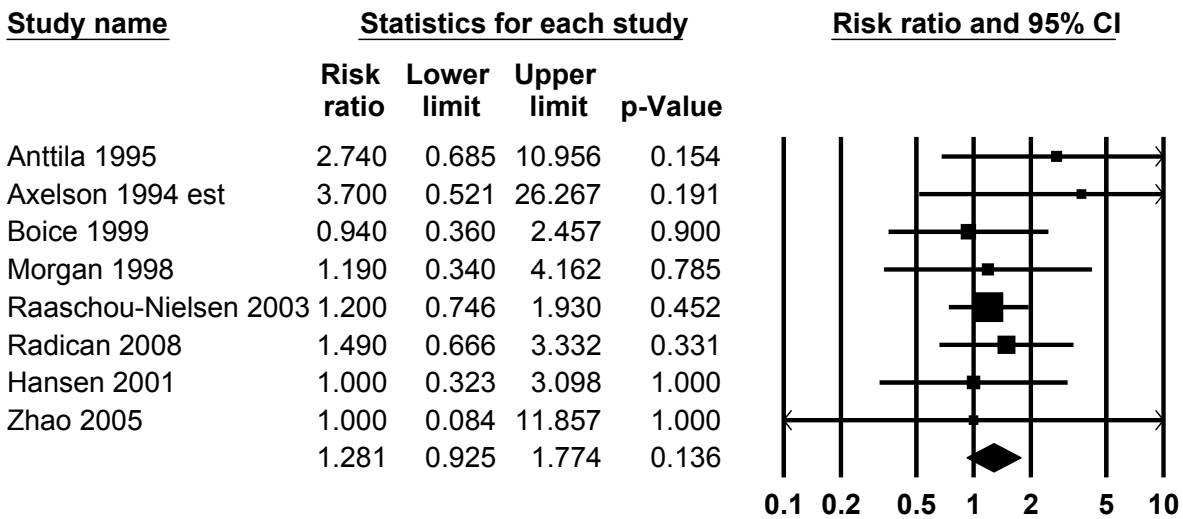
3 For Radican et al. (2008), it should be noted that the referent group is workers with no
4 chemical exposures, not just no TCE exposure. Furthermore, exposure group results were
5 reported separately for males and females and were combined for this assessment using
6 inverse-variance weighting, as in a fixed effect meta-analysis. In addition to results for biliary
7 passage and liver cancer combined, Radican et al. (2008) present results for liver only by
8 exposure group; however, there were no liver cancer deaths in females and the number expected
9 was not reported, so no alternate analysis for the highest-exposure groups with an RR estimate
10 from Radican et al. (2008) for liver cancer only was conducted. Radican et al. (2008) present
11 only mortality HR estimates by exposure group; however, in an earlier follow-up of this same
12 cohort, Blair et al. (1998) present both incidence and mortality RR estimates by exposure group.
13 As with the Radican et al. (2008) liver cancer only results, however, there were no incident cases
14 for females in the highest-exposure group in Blair et al. (1998) (and the expected number was
15 not reported). Additionally, there were more biliary passage/liver cancer deaths (31) in Radican
16 et al. (2008) than incident cases (13) in Blair et al. (1998) overall and in the highest-exposure
17 group (14 vs. 4). Thus, we elected to use only the Radican et al. (2008) mortality results from
18 this cohort and not to include an alternate analysis based on incidence results from the earlier
19 follow-up.
20

21 **C.4.2.2. Results of Meta-Analyses**

22 Results from the meta-analyses that were conducted for liver cancer in the highest
23 exposure groups are summarized at the bottom of Table C-12. The pooled RR estimate from the
24 random effects meta-analysis of the 6 studies with results presented for exposure groups was
25 1.32 (95% CI: 0.93, 1.86). As with the overall liver cancer meta-analyses, the meta-analyses of
26 the highest-exposure groups were dominated by one study (Raaschou-Nielsen), which provided
27 about 52% of the weight. The RRp estimate from the primary random effects meta-analysis with
28 null RR estimates (i.e., 1.0) included for Hansen and Zhao to address (potential) reporting bias
29 (see above) was 1.28 (95% CI: 0.93, 1.77) (see Figure C-11). The inclusion of these 2 additional
30 studies contributed about 10% of the total weight. No single study was overly influential
31 (removal of individual studies resulted in RRp estimates that ranged from 1.23 to 1.36) and the
32 RRp estimate was not highly sensitive to alternate RR estimate selections (RRp estimates with
33 alternate selections ranged from 1.24 to 1.26; see Table C-12). In addition, there was no
34 observable heterogeneity across the studies for any of the meta-analyses conducted with the
35 highest-exposure groups. However, none of the RRp estimates was statistically significant.

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TCE and Liver Cancer - highest exposure groups



random effects model; same for fixed

Figure C-11. Meta-analysis of liver cancer and TCE exposure—highest exposure groups, with assumed null RR estimates for Hansen and Zhao (see text).

Furthermore, the RRp estimates for the highest-exposure groups were all less than the significant RRp estimate for an overall effect on liver cancer (1.33; 95% CI: 1.09, 1.64; see Section C.4.2.2 and Table C-12). This contradictory result is driven by the fact that the RR estimate for the highest-exposure group was less than the overall RR estimate for Raaschou-Nielsen, which contributes the majority of the weight to the meta-analyses. The liver cancer results are relatively underpowered with respect to numbers of studies and number of cases, and the Raaschou-Nielsen study, which dominates the analysis, uses duration of employment as an exposure-level surrogate for liver cancer, and duration of employment is a notoriously weak exposure metric. Thus, the contradictory finding that the RRp estimates for the highest-exposure groups were all less than the RRp estimate for an overall effect does not rule out an effect of TCE on liver cancer; however, it certainly does not provide additional support for such an effect.

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1 C.4.3. Discussion of Liver Cancer Meta-Analysis Results

2 For the most part, the meta-analyses of the overall effect of TCE exposure on liver (and
3 gall bladder/biliary passages) cancer suggest a small, statistically significant increase in risk.
4 The pooled estimate from the primary random effects meta-analysis of the 9 (all cohort) studies
5 was 1.33 (95% CI: 1.09, 1.64). The analysis was dominated by one large study that contributed
6 about 57% of the weight. When this study was removed, the RRp estimate did not change much,
7 but it was no longer statistically significant (RRp = 1.31; 95% CI: 0.96, 1.79). The pooled
8 estimate was not overly influenced by any other single study, nor was it overly sensitive to
9 individual RR estimate selections. The largest downward impacts were from the removal of the
10 Anttila study, resulting in an RRp estimate of 1.29 (95% CI: 1.04, 1.59), and from the
11 substitution of the Boice (1999) RR estimate for potential routine exposure with that for any
12 potential exposure, resulting in an RRp estimate of 1.29 (1.06, 1.56). Substituting the RR
13 estimates for liver/gall bladder/biliary passage cancers with those of liver cancer alone for the
14 3 studies that provided these results yielded an RRp estimate of 1.31 (1.02, 1.67). There was no
15 evidence of publication bias in this data set, and there was no observable heterogeneity across the
16 study results.

17 Six of the 9 studies provided liver cancer results by exposure level. Two other studies
18 reported results for other cancer sites by exposure level, but not liver cancer; thus, to address this
19 reporting bias, null values (i.e., RR estimates of 1.0) were used for these studies. Different
20 exposure metrics were used in the various studies, and the purpose of combining results across
21 the different highest exposure groups was not to estimate an RRp associated with some level of
22 exposure, but rather to see the impacts of combining RR estimates that should be less affected by
23 exposure misclassification. In other words, the highest exposure category is more likely to
24 represent a greater differential TCE exposure compared to people in the referent group than the
25 exposure differential for the overall (typically any vs. none) exposure comparison. Thus, if TCE
26 exposure increases the risk of liver cancer, the effects should be more apparent in the highest
27 exposure groups. However, the RRp estimate from the meta-analyses of the highest exposure
28 group results were less than the RRp estimate from the overall exposure analysis. This
29 anomalous result is driven by the fact that, for Raaschou-Nielsen, which contributes the majority
30 of the weight to the meta-analyses, the RR estimate for the highest-exposure group, although
31 greater than 1.0, was less than the overall RR estimate.

32 Thus, while there is the suggestion of an increased risk for liver cancer associated with
33 TCE exposure, the statistical significance of the pooled estimates is dependent on one study,
34 which provides the majority of the weight in the meta-analyses. Removal of this study does not
35 change the RRp estimate; however, it becomes nonsignificant ($p = 0.08$). Furthermore, meta-

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1 analysis results for the highest-exposure groups yielded *lower* RRp estimates than for an overall
2 effect. These results do not rule out an effect of TCE on liver cancer, because the liver cancer
3 results are relatively underpowered with respect to numbers of studies and number of cases and
4 the overwhelming study in terms of weight uses the weak exposure surrogate of duration of
5 employment for categorizing exposure level; however, at present, there is only modest support
6 for such an effect.

8 **C.5. DISCUSSION OF STRENGTHS, LIMITATIONS, AND UNCERTAINTIES IN** 9 **THE META-ANALYSES**

10 Meta-analysis provides a systematic way of objectively and quantitatively combining the
11 results of multiple studies to obtain a summary effect estimate. Use of meta-analysis can help
12 risk assessors avoid some of the potential pitfalls in overly relying on a single study or in making
13 more subjective qualitative judgments about the apparent weight of evidence across studies.
14 Combining the results of smaller studies also increases the statistical power to observe an effect,
15 if one exists. In addition, meta-analysis techniques assist in systematically investigating issues
16 such as potential publication bias and heterogeneity in a database.

17 While meta-analysis can be a useful tool for analyzing a database of epidemiological
18 studies, the analysis is limited by the quality of the input data. If the individual studies are
19 deficient in their abilities to observe an effect (in ways other than low statistical power, which
20 meta-analysis can help ameliorate), the meta-analysis will be similarly deficient. A critical step
21 in the conduct of a meta-analysis is to establish eligibility criteria and clearly and transparently
22 identify all relevant studies for inclusion in the meta-analysis. For the TCE database, a
23 comprehensive qualitative review of available studies was conducted and eligible studies were
24 identified, as described in Appendix B, Section II-9.

25 Identifying all relevant studies may be hampered if publication bias has occurred.
26 Publication bias is a systematic error that can arise if statistically significant studies are more
27 likely to be published than nonsignificant studies. This can result in an upward bias on the effect
28 size measure, i.e., the relative risk estimate. To address this concern, potential publication bias
29 was investigated for the databases for which meta-analyses were undertaken. For the studies of
30 kidney cancer and liver cancer, there was no evidence of publication bias. For the studies of
31 lymphoma, there was some evidence of potential publication bias. It is uncertain whether this
32 reflects actual publication bias or rather an association between SE and effect size (as discussed
33 in Section C.1, a feature of publication bias is that smaller studies tend to have larger effect
34 sizes) resulting for some other reason, e.g., a difference in study populations or protocols in the
35 smaller studies. Furthermore, if there is publication bias in this data set, it may be creating an

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1 upward bias on the relative risk estimate, but this bias does not appear to account completely for
2 the finding of an increased lymphoma risk (see Section C.2.1.2).

3 Another concern in meta-analyses is heterogeneity across studies. Random-effects
4 models were used for the primary meta-analyses in this assessment because of the diverse nature
5 of the individual studies. When there is no heterogeneity across the study results, the
6 random-effects model will give the same result as a fixed-effect model. When there is
7 heterogeneity, the random-effects model estimates the between-study variance. Thus, when
8 there is heterogeneity, the random-effects model will generate wider confidence intervals and be
9 more “conservative” than a fixed-effect model. However, if there is substantial heterogeneity, it
10 may be inappropriate to combine the studies at all. In cases of significant heterogeneity, it is
11 important to try to investigate the potential sources of the heterogeneity.

12 For the studies of kidney cancer and liver cancer, there was no apparent heterogeneity
13 across the study results, i.e., random- and fixed-effects models gave identical summary
14 estimates. For the lymphoma studies, there was heterogeneity, but it was not statistically
15 significant ($p = 0.10$). The I^2 value was 33%, suggesting low-to-moderate heterogeneity. When
16 subgroup analyses were done for the cohort and case-control studies separately, there was some
17 heterogeneity in both groups, but in neither case was it statistically significant. Further attempts
18 to quantitatively investigate the heterogeneity were not pursued because of limitations in the
19 database. The sources of heterogeneity are an uncertainty in the database of studies of TCE and
20 lymphoma. Some potential sources of heterogeneity, which are discussed qualitatively in
21 Section C.2.3, include differences in exposure assessment or in the intensity or prevalence of
22 TCE exposures in the study population and differences in lymphoma classification.

23 The joint occurrence of heterogeneity and potential publication bias in the database of
24 studies of TCE and lymphoma raises special concerns. Because of the heterogeneity, a
25 random-effects model should be used if these studies are to be combined; yet, the random-effects
26 model gives relatively large weight to small studies, which could exacerbate the potential
27 impacts of publication bias. For the lymphoma studies, the summary relative risk estimates from
28 the random-effects and fixed-effect models are not very different (RRp = 1.23 [95% CI: 1.04,
29 1.44] and 1.19 [1.06, 1.34], respectively); however, the confidence interval for the fixed-effect
30 estimate does not reflect the between-study variance and is, thus, overly narrow.

31 32 **C.6. CONCLUSIONS**

33 The strongest finding from the meta-analyses was for TCE and kidney cancer. The
34 summary estimate from the primary random-effects meta-analysis of the 14 studies was
35 RRp = 1.25 (95% CI: 1.11, 1.41). There was no apparent heterogeneity across the study results

1 (i.e., fixed-effect model gave same summary estimate), and there was no evidence of potential
2 publication bias. The summary estimate was robust across influence and sensitivity analyses; the
3 estimate was not markedly influenced by any single study, not was it overly sensitive to
4 individual RR estimate selections. The findings from the meta-analyses of the highest exposure
5 groups for the studies that provided results categorized by exposure level were similarly robust.
6 The summary estimate was $RR_p = 1.53$ (95% CI: 1.23, 1.91) for the 12 studies included in the
7 analysis. There was no apparent heterogeneity in the highest-exposure group results, and the
8 estimate was not markedly influenced by any single study, nor was it overly sensitive to
9 individual RR estimate selections. In sum, these robust results support a conclusion that TCE
10 exposure increases the risk of kidney cancer.

11 For the most part, the meta-analyses of the overall effect of TCE exposure on lymphoma
12 also suggest a small, statistically significant increase in risk. The summary estimate from the
13 primary random-effects meta-analysis of the 16 studies was 1.23 (95% CI: 1.04, 1.44). This
14 result was not overly influenced by any single study, nor was it overly sensitive to individual RR
15 estimate selections, although use of one alternate RR estimate considered clearly inferior
16 narrowly eliminated statistical significance of the summary estimate ($p = 0.050$). There is some
17 evidence of potential publication bias in the lymphoma study data set; however, it is uncertain
18 that this is actually publication bias rather than an association between SE and effect size
19 resulting for some other reason, e.g., a difference in study populations or protocols in the smaller
20 studies. Furthermore, if there is publication bias, it does not appear to account completely for the
21 findings of an increased lymphoma risk. There was some heterogeneity across the results of the
22 16 studies, but it was not statistically significant ($p = 0.10$). The I^2 value was 33%, suggesting
23 low-to-moderate heterogeneity. The source(s) of this heterogeneity remains an uncertainty. The
24 summary estimate from the meta-analysis of the highest exposure groups for the 12 studies
25 which provided results categorized by exposure level was $RR_p = 1.57$ (95% CI: 1.27, 1.94).
26 This result for the highest exposure groups was not overly influenced by any single study, nor
27 was it overly sensitive to individual RR estimate selections, and heterogeneity was not observed
28 in any of the relevant analyses. The robustness of the finding of an increased lymphoma risk for
29 the highest exposure groups strengthens the more moderate evidence from the meta-analyses for
30 overall effect.

31 The meta-analyses of the overall effect of TCE exposure on liver (and gall bladder/biliary
32 passages) cancer also suggest a small, statistically significant increase in risk, but the study
33 database is more limited. The pooled estimate from the primary random-effects meta-analysis of
34 the 9 (all cohort) studies was 1.33 (95% CI: 1.09, 1.64). The analysis was dominated by one
35 large study that contributed about 57% of the weight. When this study was removed, the RR_p

1 estimate did not change much, but it was less precise (RRp = 1.31; 95% CI: 0.96, 1.79). The
2 pooled estimate was not overly influenced by any other single study, nor was it overly sensitive
3 to individual RR estimate selections. There was no evidence of publication bias in this data set,
4 and there was no observable heterogeneity across the study results. However, the findings from
5 the meta-analyses of the highest-exposure groups for the studies that provided results categorized
6 by exposure level do not add support to the overall effect findings. The summary estimate was
7 RRp = 1.28 (95% CI: 0.93, 1.77) for the 8 studies included in the analysis, which is *lower* than
8 the summary estimate for the overall effect. This contradictory result is driven by the fact that
9 the RR estimate for the highest-exposure group in the individual study which contributes the
10 majority of the weight to the meta-analyses, although greater than 1.0, was less than the overall
11 RR estimate for the same study. In sum, these results do not rule out an effect of TCE on liver
12 cancer, because the liver cancer results are relatively underpowered with respect to numbers of
13 studies and number of cases and the overwhelming study in terms of weight uses the weak
14 exposure surrogate of duration of employment for categorizing exposure level; however, at
15 present, there is only modest support for an increased risk of liver cancer.

16

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