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Psychological Distress in Mid-Life in the 1958 and 1970 Cohorts: The Role of Childhood Experiences and Behavioural Adjustment

Introduction

This paper addresses the levels of psychological distress experienced in mid-life (at age 42) by men and women born in 1958 and 1970. Comparing these cohorts born 12 years apart, we ask whether psychological distress has increased, and if so whether this increase can be explained by differences between the cohorts in their childhood conditions, including differences in their social and emotional adjustment during adolescence. While the huge costs to society, and to the economy, of poor mental health are undisputed (Layard, 2013), the idea that experiences of mental distress in adulthood are increasing across generations has not been much discussed, and yet if true, is of major societal and population health significance. Younger cohorts in the UK have gained years of life expectancy, but if these added years are to be lived healthily, these cohorts need to be healthier compared to older cohorts. In other words, in order for healthy life expectancy to grow faster than overall life expectancy and for morbidity to be compressed, the prevalence of major drivers of mortality needs to be lower in younger cohorts. If this does not happen, morbidity will expand and more years will be lived in poor health, a scenario with wide implications considering population ageing and the changing age structure in the UK.

A small number of previous studies have noted that psychological distress during adult life has been increasing across cohorts. For example, Spiers et al (2011) (Spiers et al., 2011) note an increase in common mental disorders in men born post- 1950 compared to pre-1950, using data from the British Adult Psychiatric Morbidity surveys. Sacker and Wiggins (Sacker & Wiggins, 2002) compare the 1958 and 1970 British birth cohorts up to age thirty, showing an increase in reports of psychological distress, particularly in men (and hence a narrowing of the gender gap in psychological distress). In this paper we use new data from the 1970 cohort to compare psychological distress across the 1958 and 1970 cohorts at age 42. Secondly, in acknowledgement that the roots of adult psychological distress often lie in childhood (Colman, Ploubidis, Wadsworth, Jones, & Croudace, 2007; Krause, Mendelson, & Lynch, 2003; Power, Stansfeld, Matthews, Manor, & Hope, 2002), we exploit the rich longitudinal data available in these studies since birth and across childhood to try to understand whether any observed major differences in childhood conditions between the cohorts appear to drive the differences in psychological distress in adult life.

Data

The National Child Development Study (NCDS) follows the lives of 17,000 people born in England, Scotland and Wales in a single week of 1958 (Power & Elliott, 2006). Also known as the 1958 Birth Cohort Study, it collects information on physical and educational development, economic circumstances, employment, family life, health behaviour, wellbeing, social participation and attitudes. Since the birth survey in 1958, there have been nine further 'sweeps' of all cohort members at ages 7, 11, 16, 23, 33, 42, 44, 46, 50 and 55. The 1970 British Cohort Study (BCS70) follows the lives of more than 17,000 people born in England, Scotland and Wales in a single week of 1970 (Elliott & Shepherd, 2006). Over the course of cohort members' lives, the BCS70 has collected information on health, physical, educational and social development, and economic circumstances among other factors. Since the birth survey in 1970, there have been eight surveys (or 'waves') at ages 5, 10, 16, 26, 30, 34, 38 and 42. Our analytic sample included participants that had at least four valid – not missing - responses on the Malaise inventory. Our analytic sample comprised of N = 9986 women and N = 9270 men. This limited the presence of participants with missing data to <5% of the overall sample. Nevertheless, we employed the Full Information Maximum Likelihood (FIML) method under a Missing at Random (MAR) assumption (Little & Rubin, 2002), which in this case implies that our estimates are valid if missingness is due to variables included in our models.

Measures

Outcome

Psychological distress was measured at age 42 in both cohorts using the Malaise inventory (Rodgers, Pickles, Power, Collishaw, & Maughan, 1999; Michael Rutter, Tizard, & Whitmore, 1970). The NCDS Age 42 survey used the full 24 item version of the scale whereas the BCS70 survey included 9 of the 24 items. Those with a score of 4 or more on the 9 items included in both cohorts were defined as showing signs of depression and/or anxiety.

Mediators

The mediators we included in our analysis were divided into four groups, birth characteristics, parental characteristics, child characteristics and adult characteristics. The birth characteristics included in the models were birthweight, maternal smoking during pregnancy and breastfeeding (ever breastfeed and breastfed for more than one month). The parental characteristics were paternal social class at age 19, whether the mother of the participants was employed (birth to age 5), divorce (by age 10), and years of education of both parents of the participants. The child characteristics we include in the analysis were nocturnal enuresis (wet at night after age 5) and the Rutter mental health assessment at age 16, using the same modified version of the Rutter 'A' scale (Michael Rutter et al., 1970). The scale generates an overall behavioural adjustment score in addition to four sub-scales: conduct problems, hyperactivity, emotional problems and peer problems. Although not the main focus of this paper, we also included some adult characteristics that vary between the two cohorts and have also been linked to psychological distress. These were, participants' education at age 33, partnership status at age 33, number of children by age 42 and employment status at age 33. Descriptive statistics of all mediators in our analysis are available in the Appendix.

Statistical modelling

Psychometric modelling

The probability of endorsing a particular response on dichotomous variables as the observed items of the Malaise inventory is usually quantified with a 2 parameter logistic model

$$\Pr\left(y_{ij}=1 \mid \boldsymbol{\theta}_j\right)=\frac{\exp\left(\beta_i+\lambda_i\,\theta_j\right)}{1+\exp\left(\beta_i+\lambda_i\,\theta_i\right)}$$

where $\boldsymbol{\theta}$ is the latent (unobserved) psychological distress, which is assumed to have a normal distribution N ~ (0, 1), λ is the factor loading that captures the strength of the association between the latent variable $\boldsymbol{\theta}$ and the observed items and β is the threshold or "difficulty" parameter which quantifies the level of the latent trait that needs to be reached for a response to an observed item to switch from 0 to 1 (Rabe-Hesketh & Skrondal, 2008). Within the modern Generalised Latent Variable Modelling framework for dichotomous and ordinal responses, a latent continuous response y_{ij}^{*} that underlies the observed items is invoked (Muthen, 1984):

$$y_{ij} = \begin{cases} 1 \text{ if } y_{ij}^{*} > 0 \\ 0 \text{ otherwise} \end{cases}$$

A unidimensional factor model is then specified for y_{ij}^{*} (Muthen, 1983):

$$y_{ij}^{*} = \beta_i + \lambda_i \, \theta_{\mathrm{j}} + \varepsilon_{ij}$$

$$\boldsymbol{\theta}_{j} \sim \mathrm{N}(0, \sigma), \; \varepsilon_{ij} \; \mathrm{F} \sim (0, \delta) \text{ and } \mathrm{Cov}(\boldsymbol{\theta}_{j}, \varepsilon_{ij}) = 0$$

This specification leads to the 2 parameter logistic model if instead of the normal the logistic distribution is specified for the errors $\varepsilon_{ij,}$ or a logistic link function is used. In our latent specification we opted for a normal specification of the error term and a probit link function using a cumulative normal distribution function Φ which leads to the normal ogive model for the observed y_{ij},

$$P(y_{ij}^{*} = 1 \mid \boldsymbol{\theta}_{\mathrm{j}}) = P(y_{ij}^{*} > \tau_i \mid \boldsymbol{\theta}_{\mathrm{j}}) = 1 - P(y_{ij}^{*} \leq \tau_i \mid \boldsymbol{\theta}_{\mathrm{j}}) =$$

$$= 1 - \Phi[(\tau_i - \lambda_i \boldsymbol{\theta}_{\mathrm{j}}) \, V(\varepsilon_{ij})] = \Phi[(-\tau_i + \lambda_i \boldsymbol{\theta}_{\mathrm{j}}) \, V(\varepsilon_{ij})]$$

Standardizing to $V(\varepsilon_{ij}) = 1$ defines a probit model. With this approach measurement error in the observed Malaise inventory items is controlled since the latent dimension $\boldsymbol{\theta}_j$ captures only the common variation in these and leaves out unique to each item variance (measurement error - ε_{ij}) that is not due to latent $\boldsymbol{\theta}$. However, additional sources of error may arise in between cohort comparisons from differences in the comprehension of items and in response tendencies which may vary by cohort so their distribution as sources of error cannot be assumed to be uniform between groups (Meredith, 1993). In order to obtain a meaningful comparison between the NCDS and BCS70 with respect to psychological distress ($\boldsymbol{\theta}_j$) the measurement parameters (τ and λ) of the model need to function equivalently between the two cohorts. To empirically test this assumption we estimated a multigroup confirmatory factor analysis of the normal ogive model, where measurement model parameters (τ and λ) were not allowed to vary between the two cohorts (Diagram 1).

Mediation

In order to investigate whether early life experiences and behavioural adjustment until age 16 account for between cohort differences in psychological distress at age 42 a formal mediation approach was employed. This was needed in order for the "indirect" effect of the cohort dummy variable on psychological distress and its standard error to be formally quantified. This indirect effect captures the proportion of the between cohort difference in psychological distress that is due to – or is transmitted by - the intermediate variables ("mediators"). In the causal mediation literature several approaches have been proposed for the estimation of direct and indirect effects with an emphasis on different aspects of mediation (Ten Have & Joffe, 2012). In this instance we employed appropriate link functions for the nature of the mediators and outcomes and report Natural Indirect Effects (NIE) as described in the causal mediation literature (Griffiths et al., 2013; Imai, Keele, & Tingley,

2010; Myers, 2003). Since we are interested in the extend to which each hypothesised mediator may explain between cohort differences, we report the NIE from "single mediator" models, whereas summaries for blocks mediators are derived from models where the total cohort effect is adjusted for all mediators. We have adjusted for potential intermediate confounders (De Stavola, Daniel, Ploubidis, & Micali, 2015) where appropriate and report adjusted NIEs.

Estimation

All our models were estimated with Mplus 7.2 (Mutthen & Muthen, 1998-2014). The measurement models were estimated with the Weighted Least Squares, Mean and Variance adjusted (WLSMV) estimator, whereas all other models were estimated using appropriate link functions (linear or logit where appropriate) with the Robust Maximum Likelihood (MLR) estimator with Monte Carlo integration.

INSTERT DIAGRAM 1 ABOUT HERE

Results

The multigroup confirmatory factor analysis model had very good fit, Comparative Fit Index – CFI = 0.986, Tucker Lewis Index – TLI = 0.983, Root Mean Square Error of Approximation - RMSEA = 0.037 (95%CI = 0.035 – 0.039) indicating the measurement equivalence of the Malaise inventory in the two cohorts. The standardised factor loadings (λ_i) were all high and ranged between 0.617 and 0.891, whereas the item thresholds (τ_i)were all > 0 and were located as expected towards the high end of the latent psychological distress continuum (0.413 – 1.687). Once measurement invariance was established, we did proceed to estimate the total effect of the cohort dummy on psychological distress (between cohort differences) and all indirect effects via the hypothesised mediators. For comparison purposes with other studies, we report the between cohort comparison from models on three versions of the outcome: the latent variable derived from the measurement invariant model, the sum score (9 point Malaise scale) and a binary version that captures participants that have scored >3 ("depressed"). Amongst those born in 1958, women were considerably more likely to be defined as showing signs of depression at 42 than men (16% compared with 10%). Twelve years later, among those born in 1970 levels of depression had increased significantly for both men and women, and although it remained the case that women were more likely to show signs of depression than men, a larger increase in levels of depression amongst men lead to the difference between men and women being reduced somewhat (20% of women compared with 16% of men).

INSERT TABLE 1 ABOUT HERE

Men

In men, the BCS70 cohort score b = 0.358 (0.308 to 0.407) standard deviations higher on the latent psychological distress variable, b = 0.476 (0.374 to 0.578) points higher on the 9 point Malaise scale and are OR = 1.833 (1.552 to 2.165) times more likely compared to the NCDS cohort to score more than 3 on the Malaise inventory. In Table 2 we present the standardised Natural Indirect Effects and the percentage of the between cohort differences that is explained by all the hypothesised mediators. From first glance it appears that some NIEs are positive while others are negative. The presence of negative NIEs, as for example the NIE for the mediating effect of own Education, b = -0.054 (-0.070 to -0.038) is evidence for inconsistent mediation or regression suppression (MacKinnon, Krull, & Lockwood, 2000). We note that these variables do not explain the observed difference between cohorts, as consistent mediators do. They inflate the observed difference when they are included in the

model, or "supress" it when they are excluded. For example the observed difference of b = 0.358 (0.308 to 0.407) between the two cohorts becomes b = 0.382 (0.325 to 0.439) when participants' education at age 33 is included in the model. Inconsistent mediation arises when the parameter estimates of the exposure (year of birth) - mediator (education at 33) and mediator - outcome (psychological distress at 42) regressions have opposite signs. In this instance, paternal social class, paternal and maternal education, peer pressure, being unemployed at age 33 and participants' education are typical inconsistent mediators that can be identified by the negative NIE.

With respect to birth characteristics, maternal smoking during pregnancy was the only consistent mediator that reached conventional levels of significance, b = 0.013 (0.007 to 0.019). Parental divorce, b = 0.012 (0.005 to 0.019) was the only parental characteristic with a consistent mediating effect. Conduct and emotional problems both mediated the association between year of birth and psychological distress, b = 0.012 (0.006 to 0.018) and b = 0.009 (0.002 to 0016). Partnership status at 33 was the only the adult characteristic with a consistent mediating effect b = 0.016 (0.009 to 0.022). Taken together, all consistent mediators accounted for 19.5% of the between cohort differences in psychological distress in men

INSERT TABLE 2 ABOUT HERE

Women

As expected from the descriptive analysis, the between cohort difference in psychological distress was less pronounced in women. The BCS70 scored b = 0.176 (0.125 to 0.227) standard deviations higher on the latent score, b = 0.266 (0.160 to 0.371) higher on the nine point Malaise score and were OR = 1.391 (1.207 to 1.603) times more likely to score more than 4 on the Malaise inventory. In Table 2 we present the Natural Indirect Effects, their corresponding 95% confidence intervals in the standardised latent metric for all mediators in our study and the proportion of the between cohort differences in psychological distress after we formally accounted for the effect of all consistent mediators. As in men, some mediators are consistent (positive) while others are inconsistent (negative). Maternal smoking during pregnancy, and breastfeeding for more than one month were both significant consistent mediators, b = 0.017 (0.009 to 0.022) and b = 0.018 (0.006 to 0.030). Maternal employment by age 5, b = 0.004 (0.001 to 0.006) was the only parental characteristic with a consistent mediating effect in women. Similarly, we found that conduct problems at 16 was the only consistent mediator among child characteristics, b = 0.05 (0.001 to 0.010). With respect to adult characteristics, partnership status at 33 was found to be a consistent mediator, b = 0.010 (0.005 to 0.015). Overall, all consistent mediators accounted for 30.7% of the association between year of birth and psychological distress at 42.

Discussion

We observed between cohort differences in psychological distress between the 1958 (NCDS) and 1970 (BCS70) cohorts at age 42. These differences were more pronounced in men, with the magnitude of the effect being almost twice as strong compared to women. The establishment of between cohort measurement invariance of the Malaise Inventory implies that the observed differences were due to valid between cohort variation in levels of psychological distress and not an artefact of cohort differences in response style, item comprehension or social desirability. A common bias might still have influenced the location of the latent mean of psychological distress equally in both cohorts, but not the magnitude of the observed difference. This relatively large difference in levels of psychological distress was observed despite the protective effect of the generally improved social and economic conditions that the 1970 cohort enjoyed, compared to the 1958 and older cohorts. Our formal

approach to mediation allowed us to empirically capture this protective effect in the form of "inconsistent mediation" or "regression suppression", a finding that would have been overlooked if standard methods were employed.

The observation that the more recently born 1970 cohort have higher levels of psychological distress compared to an earlier born cohort in midlife has implications for public health policy in the UK, especially considering that depression is a leading cause of Disability Adjusted Life Years (Ferrari et al., 2013). Taking into account that due to increases in life expectancy younger (more recently born) cohorts are expected to live longer, a necessary condition for compression of morbidity – in order for the added years to be lived healthily - is that controlling for age differences (or at a fixed age as in our paper), more recently born cohorts are healthier, so the average onset of morbidity is postponed (J. F. Fries, 1980; James F. Fries, Bruce, & Chakravarty, 2011). On the contrary, expansion of morbidity will most probably occur if younger cohorts are less healthy, whereas similar health levels between cohorts will most likely lead to equilibrium or mild expansion (Gruenberg, 1977; Manton, 1982; Olshansky, Rudberg, Carnes, Cassel, & Brody, 1991). Our observation for a higher average level of psychological distress in the 1970 cohort at 42 increases the likelihood of expansion of mental health related morbidity, assuming that the observed difference will persist on the same direction in the future. Therefore, unless as both cohorts age, those born in 1958 catch up and surpass the 1970 cohort on average levels of psychological distress, our findings imply that the average onset of mental health related morbidity will not be postponed in the 1970 cohort.

A reversal of the observed differences in the future, so that the 1958 cohort report higher levels of psychological distress and an expansion of mental health related morbidity is avoided, appears – at least in theory – not likely. The association between age and mental health outcomes is at best weak, which implies that large variation/changes in within cohort levels of psychological distress are not likely as both cohorts age. Furthermore, a very strong association between early and mid-life mental health with later life mental health has been reported in the literature (Lara et al., 2009; Robins & Price, 1991; M. Rutter, 1995). Lastly, the 1970 cohort reported higher levels of psychological distress than the 1958 cohort at age 33 (Sacker & Wiggins, 2002), a finding that implies a stable trend in cohort differences over time.

We attempted to explain the observed between cohort variation by investigating the explanatory power of "consistent" mediators (rather than "suppressors") focussing on early life characteristics and behavioural adjustment. We found that the mechanism that underlies the observed between cohort differences in psychological distress at 42 appears to be gender specific. In men the observed between cohort differences were mostly due to child characteristics. Behavioural adjustment at age 16 was the strongest mediator, with the 1970 cohort reporting more conduct and emotional problems and as expected from previous literature this led to higher levels of psychological distress at 42. Parental divorce and maternal smoking during pregnancy also mediated the association between year of birth and psychological distress. A different pattern of associations emerged in women. Birth characteristics, surprisingly not birth weight, but breastfeeding for more than one month and maternal smoking during pregnancy, were the strongest mediators of the association between year of birth and psychological distress at 42. In both men and women only partnership status at 33 form the adult social and demographic characteristics that were included in our study emerged as a consistent mediator, indicating that they did not explain a substantive proportion of the between cohort differences in psychological distress.

Strengths of this study are the availability of two population based birth cohorts with various measures taken at similar stages of the life-course and our modelling strategy that allowed us to establish measurement invariance/equivalence across the two cohorts.

Limitations include our reliance on self-reported data in the assessment of psychological distress. Despite our efforts to control for measurement error, some form of common to both cohorts bias could have influenced the location of the latent psychological distress means (although not the observed between cohort differences). Furthermore, we assumed an identical missing data generating mechanism in both cohorts and our mediation estimates (Natural Indirect Effects) depend on the assumption of no unmeasured mediator – outcome confounders. Making the reasonable – we believe - assumption that the year of birth – mediators and year of birth – psychological distress associations are in theory causally identified, since it is difficult to conceptualise an omitted variable that causes both being born in 1958 or 1970 and levels of the mediators or psychological distress, we focussed on the causal identification of the mediators – psychological distress association. We performed sensitivity analysis simulating the effects of potential confounders for the mediators where a "significant" effect was observed. We found that our results were robust even in moderate and strong confounding scenarios (results not presented here, available from corresponding author).

Overall, we find that despite the secular changes that resulted in important differences between the two cohorts in early life as well as adulthood characteristics, these account for only a part (17.5% in men 30.7% in women) of the between cohort differences in psychological distress at 42. Furthermore, our analysis could not disentangle the observed differences to "cohort" or "period" effects, since only one timepoint was considered. Our main endeavour in our future work will be to consider what other factors in adulthood may explain the observed between cohort difference, employ more measurement waves in order to quantify period and cohort effects and establish whether the elevated level of psychological distress in the 1970 cohort represents a general trend.

References

Colman, I., Ploubidis, G. B., Wadsworth, M. E. J., Jones, P. B., & Croudace, T. J. (2007). A longitudinal typology of symptoms of depression and anxiety over the life course. *Biological Psychiatry, 62*(11), 1265-1271. doi: 10.1016/j.biopsycli.2007.05.012

De Stavola, B. L., Daniel, R. M., Ploubidis, G. B., & Micali, N. (2015). Mediation analysis with intermediate confounding: structural equation modeling viewed through the causal inference lens. *Am J Epidemiol, 181*(1), 64-80. doi: 10.1093/aje/kwu239

Elliott, J., & Shepherd, P. (2006). Cohort Profile: 1970 British birth cohort (BCS70). *International Journal of Epidemiology, 35*(4), 836-843.

Ferrari, A. J., Charlson, F. J., Norman, R. E., Patten, S. B., Freedman, G., Murray, C. J. L., . . . Whiteford, H. A. (2013). Burden of Depressive Disorders by Country, Sex, Age, and Year: Findings from the Global Burden of Disease Study 2010. *PLoS Med, 10*(11), e1001547. doi: 10.1371/journal.pmed.1001547

Fries, J. F. (1980). Aging, natural death, and the compression of morbidity. *New England Journal of Medicine, 303*(3), 130-135. doi: 10.1056/nejm198007173030304

Fries, J. F., Bruce, B., & Chakravarty, E. (2011). Compression of morbidity 1980-2011: a focused review of paradigms and progress. *Journal of aging research, 2011*, 261702-261702. doi: 10.4061/2011/261702

Griffiths, L. J., Cortina-Borja, M., Sera, F., Pouliou, T., Geraci, M., Rich, C., . . . Dezateux, C. (2013). How active are our children? Findings from the Millennium Cohort Study. *BMJ Open, 3*(8). doi: 10.1136/bmjopen-2013-002893

Gruenberg, E. M. (1977). Failures of success. *Milbank Memorial Fund Quarterly-Health and Society, 55*(1), 3-24. doi: 10.2307/3349592

Imai, K., Keele, L., & Tingley, D. (2010). A General Approach to Causal Mediation Analysis. *Psychological Methods, 15*(4), 309-334. doi: 10.1037/a0020761

Krause, E. D., Mendelson, T., & Lynch, T. R. (2003). Childhood emotional invalidation and adult psychological distress: the mediating role of emotional inhibition. *Child Abuse & Neglect, 27*(2), 199-213. doi: 10.1016/s0145-2134(02)00536-7

Lara, C., Fayyad, J., de Graaf, R., Kessler, R. C., Aguilar-Gaxiola, S., Angermeyer, M., . . . Sampson, N. (2009). Childhood Predictors of Adult Attention-Deficit/Hyperactivity Disorder: Results from the World Health Organization World Mental Health Survey Initiative. *Biological Psychiatry, 65*(1), 46-54. doi: 10.1016/j.biopsych.2008.10.005

Layard, R. (2013). Mental health: the new frontier for labour economics. *IZA Journal of Labor Policy, 2*(1), 1-16.

Little, R. J. A., & Rubin, D. B. (2002). *Statistical Analysis with Missing Data* (Second Edition ed.). Chichester: Willey.

MacKinnon, D. P., Krull, J. L., & Lockwood, C. M. (2000). Equivalence of the Mediation, Confounding and Suppression Effect. *Prevention science : the official journal of the Society for Prevention Research, 1*(4), 173.

Manton, K. G. (1982). Changing concepts of morbidity and mortality in the elderly population. *Milbank Memorial Fund Quarterly-Health and Society, 60*(2), 183-244. doi: 10.2307/3349767

Meredith, W. (1993). Measurement invariance, factor analysis and factorial invariance. *Psychometrika, 58*(4), 525-543.

Muthen, B. (1983). Latent Variable Structural Equation Modeling with Categorical-Data. *Journal of Econometrics, 22*(1-2), 43-65.

Muthen, B. (1984). A General Structural Equation Model with Dichotomous, Ordered Categorical, and Continuous Latent Variable Indicators. *Psychometrika, 49*(1), 115-132.

Mutthen, L. K., & Muthen, B. O. (1998-2014). *Mplus User's Guide. Seventh Edition*. Los Angeles, CA.

Myers, J. (2003). Exercise and Cardiovascular Health. *Circulation, 107*(1), e2-e5. doi: 10.1161/01.cir.0000048890.59383.8d

Olshansky, S. J., Rudberg, M. A., Carnes, B. A., Cassel, C. K., & Brody, J. A. (1991). Trading Off Longer Life for Worsening Health: The Expansion of Morbidity Hypothesis. *Journal of Aging and Health, 3*(2), 194-216. doi: 10.1177/089826439100300205

Power, C., & Elliott, J. (2006). Cohort profile: 1958 British Birth Cohort (National Child Development Study). *International Journal of Epidemiology, 35*(1), 34-41. doi: 10.1093/ije/dyi183

Power, C., Stansfeld, S. A., Matthews, S., Manor, O., & Hope, S. (2002). Childhood and adulthood risk factors for socio-economic differentials in psychological distress: evidence from the 1958 British birth cohort. *Social Science & Medicine, 55*(11), 1989-2004. doi: 10.1016/s0277-9536(01)00325-2

Rabe-Hesketh, S., & Skrondal, A. (2008). Classical latent variable models for medical research. *Statistical Methods in Medical Research, 17*(1), 5-32. doi: 10.1177/0962280207081236|issn 0962-2802

Robins, L. N., & Price, R. K. (1991). ADULT DISORDERS PREDICTED BY CHILDHOOD CONDUCT PROBLEMS - RESULTS FROM THE NIMH EPIDEMIOLOGIC CATCHMENT-AREA PROJECT. *Psychiatry-Interpersonal and Biological Processes, 54*(2), 116-132.

Rodgers, B., Pickles, A., Power, C., Collishaw, S., & Maughan, B. (1999). Validity of the Malaise Inventory in general population samples. *Social psychiatry and psychiatric epidemiology, 34*(6), 333-341.

Rutter, M. (1995). RELATIONSHIPS BETWEEN MENTAL-DISORDERS IN CHILDHOOD AND ADULTHOOD. *Acta Psychiatrica Scandinavica, 91*(2), 73-85. doi: 10.1111/j.1600-0447.1995.tb09745.x

Rutter, M., Tizard, J., & Whitmore, K. (1970). *Education, health and behaviour*: Longman Publishing Group.

Sacker, A., & Wiggins, R. (2002). Age–period–cohort effects on inequalities in psychological distress, 1981–2000. *Psychological Medicine, 32*(06), 977-990. doi: doi:10.1017/S0033291702006013

Spiers, N., Bebbington, P., McManus, S., Brugha, T. S., Jenkins, R., & Meltzer, H. (2011). Age and birth cohort differences in the prevalence of common mental disorder in England: National Psychiatric Morbidity Surveys 1993–2007. *The British Journal of Psychiatry, 198*(6), 479-484.

Ten Have, T. R., & Joffe, M. M. (2012). A review of causal estimation of effects in mediation analyses. *Statistical Methods in Medical Research, 21*(1), 77-107. doi: 10.1177/0962280210391076

Diagram 1. Measurement model of the Malaise Inventory

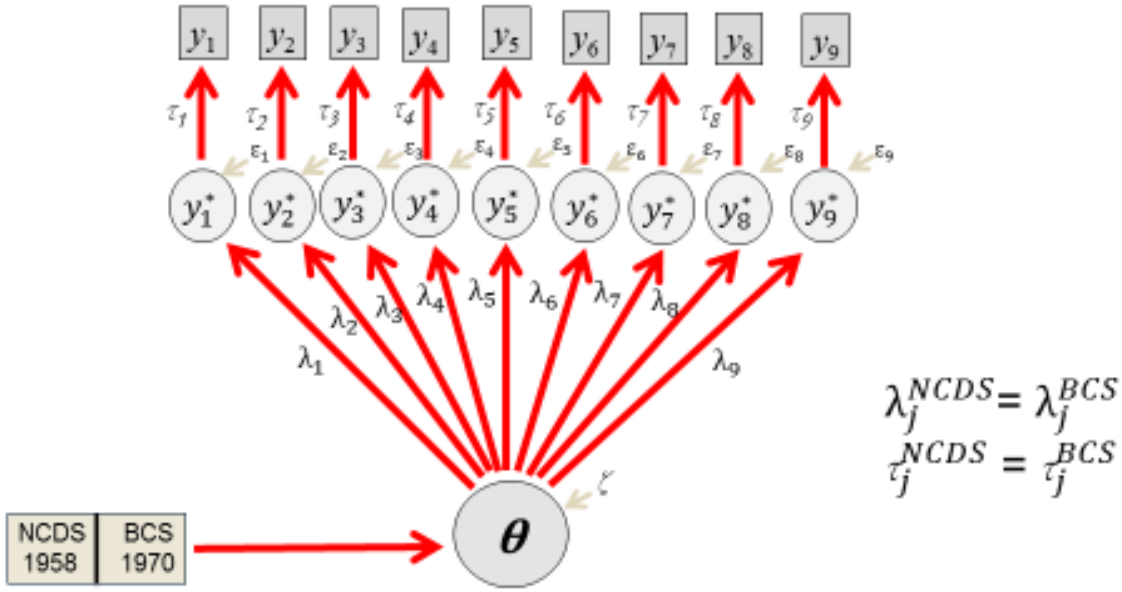

Table 1. Descriptive statistics of the Malaise Inventory

	Males	
	1958	**1970**
Depressed (Malaise > 3)	9.6% (8.9 – 10.5)	16.0% (14.7 – 17.1)
Malaise sum score mean	1.23 (1.19 – 1.27)	1.65 (1.60 – 1.72)
	Females	
	1958	**1970**
Depressed (Malaise > 3)	16.3% (15.5 – 17.4)	20.2% (19.1 - 21.4)
Malaise sum score mean	1.80 (1.75 – 1.86)	2.05 (1.99 – 2.11)

Table 2. Model parameters (Total and Natural Indirect Effects) and 95% confidence intervals.

		Men			Women	
	b	95% CI		b	95% CI	
Cohort Total Effect - Ref:1958 cohort	**0.358**	**0.308**	**0.407**	**0.176**	**0.125**	**0.227**
Cohort via Birth characteristics						
Birthweight	0.001	-0.002	0.002	-0.001	-0.002	0.001
Maternal smoking during pregnancy	**0.013**	**0.007**	**0.019**	**0.017**	**0.009**	**0.022**
Breastfeeding	0.001	-0.014	0.016	0.010	-0.004	0.023
Breastfeeding > 1 month	0.006	-0.019	0.007	**0.018**	**0.006**	**0.030**
Cohort via Parental characteristics	b	95% CI		b	95% CI	
Paternal Social Class (10/11)	**-0.009**	**-0.015**	**-0.003**	**-0.023**	**-0.030**	**-0.015**
Mother working (birth to 5)	0.002	-0.001	0.004	**0.004**	**0.001**	**0.006**
Parental divorce (by age 10)	**0.012**	**0.005**	**0.019**	0.005	-0.001	0.011
Years of education - Father	**-0.008**	**-0.014**	**-0.001**	**-0.011**	**-0.017**	**-0.005**
Years of education - Mother	**-0.011**	**-0.020**	**-0.003**	**-0.011**	**-0.017**	**-0.005**
Cohort via Child characteristics	b	95% CI		b	95% CI	
Enuresis - wet at night after 5 years	0.008	-0.001	0.017	0.006	-0.001	0.010
Total Rutter score (age 16)	0.008	-0.001	0.016	-0.001	-0.009	0.007
Conduct	**0.012**	**0.006**	**0.018**	**0.005**	**0.001**	**0.010**
Hyperactivity	**-0.004**	**-0.008**	**-0.001**	-0.004	-0.008	0.001
Emotional	**0.009**	**0.002**	**0.016**	0.007	-0.001	0.016
Peer	**-0.012**	**-0.018**	**-0.006**	**-0.004**	**-0.007**	**-0.001**
Cohort via Adult characteristics	b	95% CI		b	95% CI	
Unemployed (age33)	**-0.008**	**-0.013**	**-0.003**	-0.001	-0.003	0.001
Education - has a degree (age33)	**-0.054**	**-0.070**	**-0.038**	**-0.082**	**-0.100**	**-0.063**
Married or Cohabiting (age 33)	**0.016**	**0.009**	**0.022**	**0.010**	**0.005**	**0.015**
> 2 children (by age 42)	0.004	-0.001	0.007	0.002	-0.001	0.005
Mediation (%) by all mediators	17.5%	12.2%	23.6%	30.7%	24.7%	42.4%

*Highlighted parameters are significant at $p<0.001$