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Deindustrialisation and the post-socialist mortality crisis

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An unprecedented mortality crisis struck Eastern Europe during the 1990s, causing around seven million excess deaths. We enter the debate about the causes of this crisis by performing the first quantitative analysis of the association between deindustrialisation and mortality in Eastern Europe. We develop a theoretical framework identifying deindustrialisation as a process of social disintegration rooted in the lived experience of shock therapy. We test this theory relying on a novel multilevel dataset, fitting survival and panel models covering 52 towns and 42,800 people in 1989–95 in Hungary and 514 towns in European Russia in 1991–99. The results show that deindustrialisation was directly associated with male mortality and indirectly mediated by hazardous drinking as a stress-coping strategy. The association is not a spurious result of a legacy of dysfunctional working-class health culture aggravated by low alcohol prices during the early years of the transition. Both countries experienced deindustrialisation, but social and economic policies have offset Hungary's more immense industrial employment loss. The results are relevant to health crises in other regions, including the deaths of despair plaguing the American Rust Belt. Policies addressing the underlying causes of stress and despair are vital to save lives during painful economic transformations.

Key words: Deindustrialisation, Mortality, Stress, Eastern Europe, Multilevel modelling
JEL classifications: B520 Current Heterodox Approaches; I15 Health and Economic Development; P3 Socialist Institutions and Their Transitions

1. Introduction

An unprecedented mortality crisis hit the former socialist countries in the early 1990s—a phenomenon that Ellman (1994) famously labelled 'katastroika'. The number of excess deaths could have been around 7.3 million in Eastern Europe in 1991–99 (Stuckler, 2009, p. 7). Male life expectancy in Russia declined by seven years

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between 1988 and 1995. As Popov (2012, p. 324) noted: 'Never after the 1947 famine had Russia had, in the post-war period, mortality rates as high as those in the 1990s. Even in 1950–53, during the last years of Stalin's regime, with the high death rates in the labour camps and the [delayed] consequences of wartime malnutrition and injuries, the mortality rate was only nine to ten per 1,000, compared with 14–16 in 1994'. Hungary also suffered a significant though less dramatic mortality crisis, although worse than its Visegrad neighbours (Chenet *et al.*, 1996). Male life expectancy in Hungary declined by 1.5 years between 1988 and 1994, and death rates reached levels last observed during the Great Depression of the 1930s, that is, 14.5 per 1,000 in 1993 (Kopp *et al.*, 2007, p. 326).

Life chances have improved since the second half of the 1990s, but wide health inequalities still plague most countries. In addition to emigration and low fertility, these health problems are the primary reasons why 15 out of the 20 fastest-shrinking countries are located in Eastern Europe (United Nations, 2022). Figure 1 summarises the post-socialist mortality crisis in Russia and Hungary, compared to the average of Visegrad countries.[1]

Working-class men without a college degree suffered the most; therefore, we focus on male life expectancy and death rates. In the 1990s, blue-collar male workers had 111% higher odds of dying than those with a college degree in Hungary, a 17%

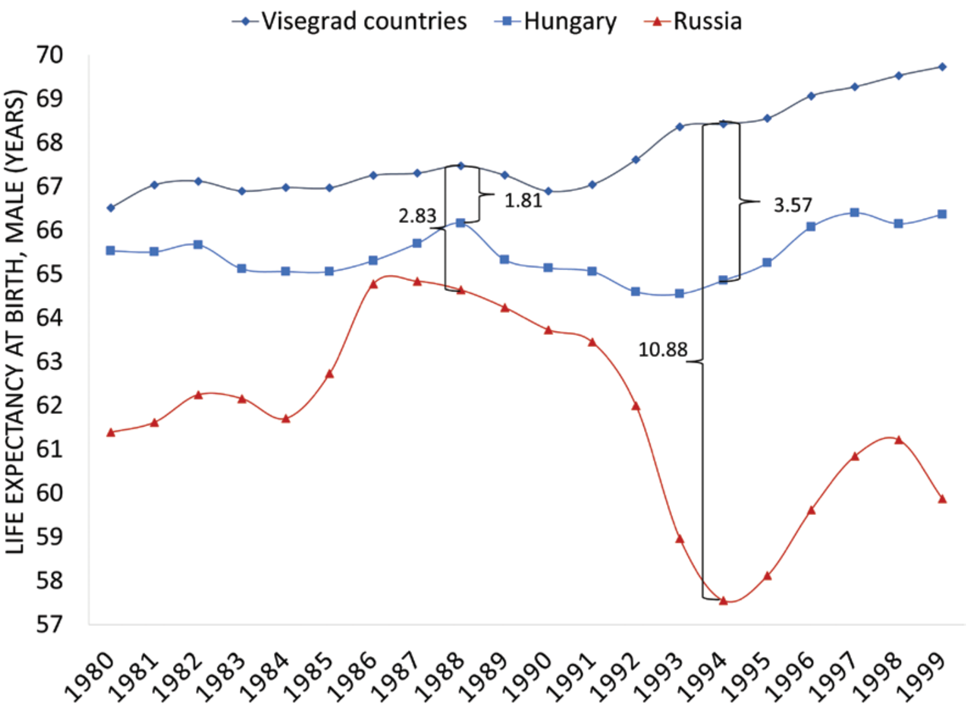

Fig. 1. *Post-socialist mortality crisis in Hungary and Russia.*
Note: Visegrad countries' average includes the Czech Republic, Poland and Slovakia (excludes Hungary).
Source: WHO (2020) Health for All Europe Database.

[1] Czech Republic, Poland and Slovakia, excluding Hungary.

increase from the 1980s, and 50% higher odds of dying in Russia, a 14% increase from the 1980s (Doniec *et al*., 2018). However, this does not mean that absolute material deprivation drove the mortality crisis. Wealthier, more urbanised, and industrialised regions in Russia suffered a more severe wave of excess deaths than impoverished regions (Walberg *et al*., 1998). Poverty-related malnutrition played no meaningful role (Ellman, 1994; Cornia and Paniccià, 2000).

In both countries, deaths considered sensitive to social disintegration and the resulting psychosocial stress and associated behaviours increased the most, including deaths due to mental disorders, homicides, alcohol (digestive system diseases, injury and poisoning), and heart disease—with heavy alcohol consumption now known to play a disproportionate role in cardiovascular deaths in Russia (Tomkins *et al*., 2012). Appendix A of Supplementary Data (patterns of the post-socialist mortality crisis, Supplementary Figure A1, Supplementary Tables A1 and A2) presents detailed data on age- and cause-specific mortality trends, further underpinning the centrality of stress-related mechanisms.

While there is widespread consensus on the proximal, or downstream, causes, especially easy access to alcohol, disagreement persists concerning the upstream, political–economic factors—what Marmot (2018) called the 'causes of causes' of ill health. Many argue that economic dislocation caused stress and despair, which led to elevated mortality. Others hypothesise that adverse lifestyles—conditioned by the socialist legacy of dysfunctional working-class health culture and aggravated by a marked increase in the supply of alcohol, especially illicit or surrogate beverages (Gil *et al*., 2009)—are the main culprits. Despite the intensity of this debate and the clear-cut policy relevance, curiously, the role of industrial employment decline has not received much attention. In a recent study, Scheiring and King (2022) qualitatively analysed the health implication of deindustrialisation. However, a quantitative assessment of this association is still lacking. This article aims to fill this gap.

Understanding the role of deindustrialisation in the post-socialist mortality crisis is crucial. First, 30 years after the fall of the Soviet Union, there is a continued interest in assessing the social costs of the transformation (Ellman, 2000; Ghodsee and Orenstein, 2021). Many connect the rise of populism in Eastern Europe to 'the failure of liberalism to deliver' (*cf.* Krastev, 2016; Scheiring, 2020). Workers' physical and mental suffering in left-behind areas is a critical correlate of anti-liberal, populist attitudes (Koltai *et al*., 2020; Kavanagh *et al*., 2021). Therefore, the insights from analysing the deindustrialisation–mortality association go beyond public health.

Second, many emerging economies are experiencing 'premature deindustrialisation' (Tregenna, 2008, 2016; Rodrik, 2016), with potentially severe health and well-being costs. Third, global competition, technological change, and the imperative to mitigate climate change will continue to put pressure on industrial employment, which could harm workers' health in rustbelt areas. Fourth, deindustrialisation appears to be a critical factor in the 'deaths of despair' plaguing the USA (Case and Deaton, 2020; Venkataramani *et al*., 2020). Better understanding the association of deindustrialisation with the wave of excess deaths hitting Eastern Europe in the 1990s promises important theoretical and policy insights for these fields (King *et al*., 2022).

We enter the debate about the causes of this crisis by performing the first quantitative analysis of the association between deindustrialisation and mortality in Eastern Europe. We develop a theoretical framework identifying deindustrialisation as a process of social disintegration rooted in the lived experience of shock therapy. Empirically, we ask whether deindustrialisation was directly and independently associated with the excess deaths and increased hazardous drinking driven by stress and despair. Second, we ask whether the effect of deindustrialisation on mortality is a spurious consequence of dysfunctional working-class health culture. We test whether hazardous drinking habits could be considered a consequence of previous socialist industrialisation (as some argue) and whether low alcohol prices acted on this dysfunctional working-class culture leading to elevated death rates. We use an innovative multilevel dataset and a retrospective cohort study originating in the Privatisation and Mortality project, the biggest exercise in individual data-gathering addressing the post-socialist mortality crisis to date (see Irdam *et al.*, 2016).

2. Competing perspectives on the lived experience of shock therapy

We rely on the heterodox literature critiquing shock therapy as the theoretical backdrop for interpreting the political–economic determinants of the post-socialist mortality crisis. Several problems had plagued socialist economies; there was a consensus that the outdated industrial structure required reform. However, the modalities of this transformation have been controversial (see the review by Ellman, 1997a). The transition orthodoxy, influenced by the Washington Consensus, advocated shock therapy. Its proponents hypothesised that there was an 'enormous scope for increases in living standards in a few years, particularly as resources are shifted out of the military-industrial complex into other sectors' (Lipton *et al.*, 1992, p. 214). Heterodox, institutionalist economists and economic sociologists criticised this approach. They argued that shock therapy unnecessarily destroyed companies and weakened institutions necessary for a successful transition (Amsden *et al.*, 1994; Kolodko, 2000; King, 2003; Popov, 2012).

Scholars tend to concentrate on the macroeconomic component of shock therapy, that is, stabilisation. In some cases, macroeconomic stabilisation was necessary to bring down inflation, but this could also be achieved through heterodox stabilisation measures without radical neoliberal monetary policy instruments, such as a currency board that binds policy-makers to internal devaluation (austerity, real wage reduction) as the only option for macroeconomic adjustment. Scholars disagree whether the macro-policies implemented in Russia in the early 1990s constituted an abortive attempt at shock therapy (Aslund and Layard, 1993) or delivered an actual shock therapy package (Murrell, 1993; King, 2002). Whichever is the case, it is clear that these macroeconomic measures failed to achieve the stated goal of shock therapy, which is to reduce inflation to a modest level in Russia. Whether this was due to a flaw in the concept or its execution continues to be debated. We do not seek to resolve this dispute here, as we concentrate on the microeconomic aspects of the chosen policies.

In addition to the macro-stabilisation component, shock therapists also advocated for a 'big bang' of liberalisation and privatisation. The orthodox approach contends that the faster these microeconomic reforms are implemented, the less opportunity there is for resistance to block them, and the earlier the benefits will materialise. This microeconomic component of shock therapy (i.e. rapid privatisation, trade liberalisation, strict bankruptcy laws, and the abandonment of activist industrial policy) was particularly problematic. It neglected the inherent slowness of restructuring companies, the need to create a dense domestic economic fabric, and to build state capacity.

Advocates of neoliberal shock therapy often cite Poland or former Czechoslovakia as proof of its superiority. However, these countries steered far away from microeconomic shock therapy, avoiding unnecessary bankruptcies and building strong state institutions (Orenstein, 2001; King and Sznajder, 2006). This contrasts with Hungary's and Russia's more radical neoliberal approach to microeconomic restructuring. Consequently, deindustrialisation in Poland, Slovakia and the Czech Republic was much less severe than in Hungary or Russia (Scheiring, 2020, pp. 133–86).

Here, we add to the evidence on the lived experience of this radical industrial transformation by examining the deindustrialisation–mortality association. A large body of scholarship is devoted to the health consequences of the post-socialist transformation (see the systematic reviews by Scheiring *et al.*, 2018a, 2019). This scholarship comprises two approaches: those arguing that economic stress associated with neoliberal reforms is the main factor of the excess deaths in the 1990s (from now on, the 'dislocation-despair approach'), versus those questioning this link and emphasising the dysfunctional health habits of working-class people aggravated by the affordability of alcohol during the early transition (from now on, the 'dysfunctional culture' approach).

Followers of the dislocation-despair approach pointed to how microeconomic shock therapy was implicated in a large part of the excess deaths of the early 1990s. Unemployment and labour market turnover strongly correlated with the wave of post-socialist mortality (Walberg *et al.*, 1998; Perlman and Bobak, 2009). Even those who kept their job but experienced fear of job loss, higher workload and decreased control at work were at higher risk of dying (Lundberg *et al.*, 2007). Deindustrialisation led to social disintegration, status loss, the loss of communities and a cascade of infrastructural, social and health problems, depression and despair (Kideckel, 2008; Scheiring and King, 2022). This despair and distress correlated with increased mortality (Kopp *et al.*, 2007). A related stream of studies showed that mass privatisation was a critical economic policy factor behind the transformation-associated economic crisis (Hamm *et al.*, 2012), also driving the life expectancy decline (Stuckler *et al.*, 2009; Azarova *et al.*, 2017), and alcohol-related deaths in Russia (King *et al.*, 2009), and in Hungary (Scheiring *et al.*, 2018b).

However, proponents of the dysfunctional culture approach question the centrality of socio-economic dislocation. Cockerham (1997, p. 127) argued that 'evidence is lacking that stress per se can account for the sharp rise in male deaths throughout the region' and proposed that unhealthy lifestyles are the main culprits. Regressing death rates on a measure of socialist industrialisation, Carlson and Hoffmann (2011, p. 375) concluded that state socialist development policies emphasising industrial employment 'created anomic conditions leading to unhealthy lifestyles and self-destructive behaviour among men,' explaining the rise in mortality until the middle of the 1990s.

Other proponents of the dysfunctional culture approach suggested that the Gorbachev anti-alcohol campaign 'saved' many male lives, and these men started to die after the liberalisation reforms, as alcohol became cheaper and more accessible, allowing working-class people to indulge in their unhealthy drinking habits, leading to the wave of excess deaths observed in the early 1990s (Treisman, 2010; Bhattacharya *et al.*, 2013). However, as we have noted previously, this is inconsistent with detailed demographic analysis (Stuckler *et al.*, 2012). In sum, according to the dysfunctional culture approach, self-destructive health behaviour—in the case of the post-socialist mortality crisis primarily referring to hazardous drinking—is not a result of the radical economic policies adopted during the transition from socialism to capitalism but of a dysfunctional, anomie-laden culture inherited from the socialist past whose effect was made worse by low alcohol prices during the transition.

3. The political economy of deindustrialisation and mortality

In this section, we present a theoretical framework that captures the human dimension of neoliberal shock therapy through the lived experience of deindustrialisation. According to Kaldor's second growth law, industrialisation leads to economy-wide productivity growth through dynamic economies of scale: manufacturing is the engine of growth (Kaldor, 1967). Industry has more linkages than other sectors of the economy; thus, it has a bigger multiplier effect (Hirschman, 2013). Heterodox economists also showed that industrialisation is tightly interwoven with the development of social cohesion (Ramazzotti, 2009) and is necessary to ensure the provision of basic needs (Singh, 1979). Therefore, a high rate of loss of industrial capacity could create a cascade of economic and social problems. This deindustrialisation is driven not only by the quest for productivity gains but also by external shocks facilitated by neoliberal microeconomic policies in emerging and core capitalist countries (Rodrik, 2016; Tregenna, 2016; Felipe *et al.*, 2018).

Deindustrialisation can be captured in two dimensions, output and employment decline. As Felipe *et al.* (2018) also show, compared with employment, output change is a weak predictor of prosperity and is under less pressure in emerging economies. Industrial employment decline is particularly important for health outcomes. A severe drop in industrial employment is most likely to signify mass plant closures, while production might decline for many reasons without necessarily affecting the socio-economic fabric of the town in which the plants affected are located. If plants survive, they can respond to a future recovery by adding hours worked or re-employing those laid off. However, much more is lost when plants are closed, with large-scale job losses, as the costs, in labour and capital, of restarting are often extremely high.

Deindustrialisation entails a loss of a complex set of socio-economic linkages that are very difficult to re-establish. As capital escapes from deindustrialised areas, local infrastructures collapse, with a loss of services, such as health, education, family support, or transport that were either provided directly by the large plants or by the local authorities that they helped fund. This creates a downward spiral of social and economic disintegration, leading to a regional lock-in of rustbelts (Hassink, 2010). Deindustrialisation could lead to a cascade of social problems, such as increasing income inequalities as it creates winners and losers (Morris and Western, 1999), growth of precarious jobs and in-work poverty (Burchell, 2009), or the erosion of communities and communal identities (Ramazzotti, 2009; Rodríguez-Pose, 2018), which in turn could lead to ill health. The growth of service sector jobs is no substitute for the lost industrial capacity as 'most skills acquired in manufacturing travel very poorly to service occupations' (Iversen and Cusack, 2011, p. 326).

Deindustrialisation in post-socialist Eastern Europe was a particularly painful social process. Socialist industry played a crucial role in workers' lives, providing stable, lifetime jobs and a comparatively high salary. Industrial workers enjoyed high social status as the backbone of state socialist societies (Burawoy and Lukács, 1992). Companies also provided many free services, including healthcare, housing, holiday homes, sports- and cultural facilities. Russian enterprises spent around 3%–5% of GDP on social provision, while East European firms spent about half this amount, which is still very important for the beneficiaries (Cook, 2007, pp. 39–40). Industrial employment also contributed to social integration, vibrant work- and neighbourhood communities (Kideckel, 2008; Scheiring, 2020). Thus, although industrial working-class culture was also associated with harmful behaviours, industrial plants contributed to health and well-being in many ways. These company functions were lost with mass plant closures, with significant health implications.

Borrowing from Scheiring and King (2022), we conceptualise industry as a social institution, allowing us to capture deindustrialisation's multidimensional health implications. The collapse of the industry as an institution engenders social disintegration, leading to ruptures in economic production and social reproduction. These ruptures entail job and income loss, increased exploitation, social inequality and the disruption of services previously provided by industrial companies. These ruptures in economic production affect social reproduction, leading to adverse outcomes, such as material deprivation, job strain, fatalism, increased domestic workload, anomie, community disintegration and alienation. The framework shown in Figure 2 captures the multidimensional and long-term socio-economic effects of deindustrialisation.[2]

These ruptures in social reproduction are sources of psychosocial stress, through which deindustrialisation gets embodied as dysfunctional health behaviour and ill health. Deindustrialisation is a short-term adverse life event (*acute stress* through economic deprivation, job loss). It also increases long-term strain (*chronic stress* through increased workload, loss of status, loss of communities, the stress of inequalities), requiring a significant behavioural change to adapt (Thoits, 2010, p. 45). The accumulation of stressors eventually depletes individuals' physical or psychological coping resources, negatively affecting psychological health, immune and cardiovascular

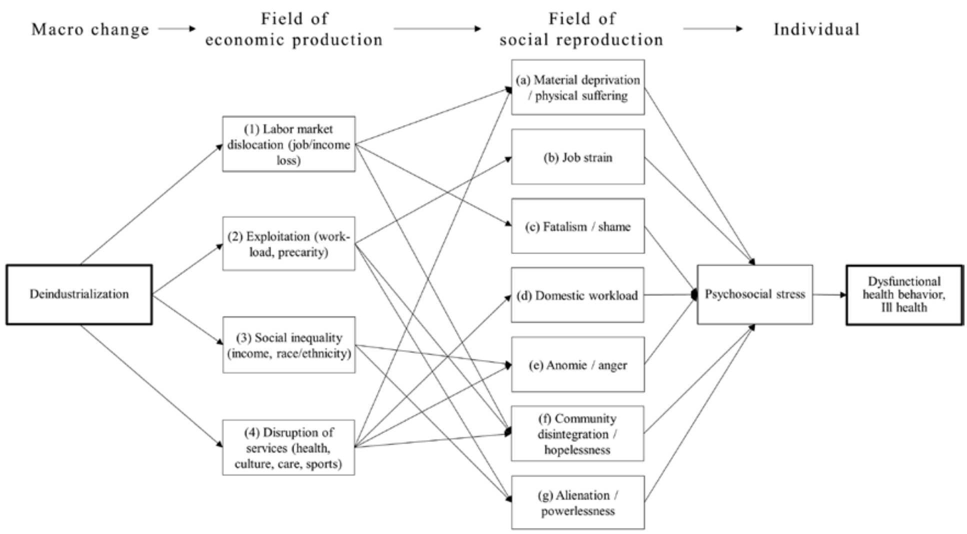

Fig. 2. *Theoretical framework: deindustrialisation, social disintegration and health.*
Note: Based on Scheiring and King (2022, p. 12).

[2] Scheiring and King (2022) combine insights from Marx, Durkheim and Polanyi, elaborating on their theoretical framework in more detail and show its usefulness analysing life-history interviews of workers affected by deindustrialisation in Hungary.

systems (McEwen and Stellar, 1993), inducing harmful health behaviour, such as hazardous drinking (Cooper *et al.*, 1992; Tomkins *et al.*, 2007).

Marmot and Bobak (2000) found that stressful situations cause a higher secretion of cortisol, endorphins, platelets, fibrinogens, fibrinolysis and other substances. These affect the level of plasma lipids, blood coagulability, blood pressure, cardiovascular reactivity, central obesity, responses to inflammation or infection, depression, coronary artery atherogenesis and a weakening of the immune system, that is, changes affecting cardiovascular mortality. Psychosocial stress has been shown to indirectly affect health via the increased use of stress relievers such as alcohol, tobacco and drugs, which influence health and social behaviours and the ability to maintain emotional balance (Cooper *et al.*, 1992). Through this stress mechanism, deindustrialisation can lead to worse self-reported health (Mitchell *et al.*, 2000), lower life expectancy (Nosrati *et al.*, 2018), and elevated drug- and alcohol-related deaths (Autor *et al.*, 2019; Venkataramani *et al.*, 2020), especially when accompanied with a mix of neoliberal policies (Walsh *et al.*, 2009), as the extant literature on Western Europe and the USA has established.

Based on these considerations, we hypothesise that (a) deindustrialisation was associated with mortality directly during the post-socialist mortality crisis, (b) indirectly mediated by increased drinking as a dysfunctional stress-coping mechanism, and (c) its effects cannot be reduced to inherited working-class culture or alcohol price policies. We test the third hypothesis derived from the main alternative explanation by examining whether hazardous drinking habits were inherited from the past due to socialist industrialisation and whether low alcohol prices activated this dysfunctional working-class culture leading to elevated death rates.

4. Data and methods

4.1 Data

The analysis relies on two datasets, one on Hungary and the other on Russia. For Hungary, we assembled a novel multilevel dataset comprising town and individual-level data as part of the Privatisation and Mortality (PrivMort) project (for a detailed description of the study protocol, see Irdam *et al.*, 2016). When analysing Hungary, we focus on 1989–95. This was when deindustrialisation and the wave of excess deaths were the most severe. There was no deindustrialisation in Hungary after 1995, and it was less significant before 1988.[3]

The first pillar of the multilevel dataset (level 1) comprises a broad range of individual-level data collected in a retrospective cohort study between January 2014 and December 2015. In each selected town, interviewers visited randomly selected addresses for face-to-face interviews. The primary screening criteria were having parents, siblings or partners living in the same settlement between 1980 and 2010. Interviewers only interviewed one respondent from each household, regardless of the family size. Following the method of retrospective cohort studies, respondents provided information on their relatives (parents, siblings and spouses). This indirect approach to collecting mortality data from relatives (also known as the 'Brass technique') produces

[3] See Figure 3. The Hungarian Central Statistical Office stopped reporting annual town-level industrial employment from 1998. Since we are interested in the effect of deindustrialisation in the mortality crisis, we do not extend the multilevel models to cover 1996–97 when the economy and life expectancy were already growing. Robustness checks using town-level panel data including 1996 and 1997 confirm the significant associations (see Supplementary Table E4).

results that are robustly consistent with mortality estimates from administrative sources in Russia (Bobak *et al*., 2003).

As respondents could not have died, we only included respondents' relatives in the analysis who lived in the surveyed towns during the 1980s and the 1990s. We further reduced the sample by excluding cases with missing information on gender. In total, 42,800 people met these criteria and were included in this analysis: 24,377 men and 18,423 women. The average number of subjects per town was 881. The average response rate was 85%. Survey questions and the resulting variables used for the analysis include the following: relationship status of the survey respondent and the subject, vital status, year of birth and death (if applicable), age in 1989, gender, education, smoking, alcohol consumption and marital status. The subjects' mean age in 1989 was 51 years. Supplementary Table B1 presents descriptive statistics of the individual-level data for Hungary. Supplementary Table B2 gives a detailed overview of the survey items and variable coding.

The second pillar of the multilevel dataset (level 2) covers town-level data. We randomly selected towns with inhabitants between 5,000 and 100,000 and industrial employment (as a share of total employment) in 1989 exceeding 30%. TÁRKI Social Research Institute, Hungary's leading polling agency, conducted individual surveys in these towns as described above. Because our primary concern is the health impact of the collapse of industrial plants, we concentrated on towns with significant industrial capacities. Predominantly agricultural small towns and villages experienced a different type of economic shock.

Our sample includes only medium-sized towns in both Hungary and Russia.[4] The main reason for this is that we rely on the PrivMort project database, whose primary aim was to analyse the association between privatisation and mortality. There are many potential privatised companies in large cities, making data collection and linkage of individuals to companies daunting. In contrast, in medium-sized towns, the five largest companies capture a large enough share of the population to detect the effect of privatisation. This dataset is not ideal, but it is still the best dataset available to analyse the impact of deindustrialisation on mortality in Eastern Europe.

Regarding economic geography, Hungary is a highly polarised country, dominated by Budapest and its agglomeration, which could skew the results; thus, the Budapest agglomeration was excluded from the sample. This way, we generated a set of 52 towns covering the entire geographical area of Hungary outside the capital. The sample represents the types of mid-sized towns where most Hungarians live and can be used to assess the impact of deindustrialisation in non-metropolitan urban areas. Supplementary Tables B3–B5 present an overview of town-level data for Hungary.

We collected data from the Hungarian Central Statistical Office on annual industrial employment in each town and calculated industrial employment as a percentage of the population. While the decline in manufacturing employment in the share of total employment might be a better measure, such data are not available at the level of towns in Hungary. The Hungarian Central Statistical Agency only published the number of persons employed in industry (manufacturing, mining, construction) at the level of towns. Total employment is only available from the censuses conducted every 10 years. It would be misleading to calculate the industrial employment share by relying on total employment in 1990. Therefore, we concentrate on industrial employment decline

[4] Hungary only has nine cities with more than 100,000 inhabitants.

as a percentage of each town's population. Additionally, we gathered data on towns' unemployment, dependency ratio, death rates, income per capita, number of general practitioners, outmigration and immigration as control variables.

The second dataset covers towns in the European part of Russia, also collected under the PrivMort project's auspices. Because Russia experienced two economic crises during the transformation in the 1990s and more prolonged deindustrialisation continuing at least until 1998, we analyse the 1991–99 period. The dataset includes 514 medium-sized towns in the European part of the country where the population exceeded 3,000 but was smaller than 200,000 in 1989. These towns are in the 49 most populous regions of Russia. The total population in the sampled towns was 20.9 million, that is, 14.1% of Russia's total population.

We assessed the quality of mortality data in the sample of 514 towns by calculating the mean crude death rates across the towns each year between 1991 and 1999 and compared them with the respective national-level statistics, as shown in Supplementary Figure B1. While at lower levels, the line for the sample of towns generally tracks the national-level trend. Therefore, we conclude that the selected towns provide a robust sample to assess the association between deindustrialisation and mortality in European Russia.

For Russia, gender-specific town-level death rates were unavailable, so we use the overall death rates covering men and women. In addition to the income, age structure, number of inhabitants, net migration, healthcare provision and housing conditions, we were also able to collect data on regional-level average alcohol consumption (per capita) and the regional-level average alcohol price.[5] The consumption variable does not include consumption of illicitly distilled alcohol, as the official statistics did not report these. The price of alcohol is expressed in roubles, corresponding to the average for each year in the regional capital where the town is located, deflated to 1991 by the CPI to avoid bias arising from the hyperinflation characterising the early 1990s in Russia. Even though 1991 was the last year of the Soviet command economy, deflating wages and prices to the price level in that year is a reasonable strategy, given the hyperinflation in the following periods. For example, in January 1992, the rate was as high as 245%, totalling 2509% yearly in 1992.

We use these two alcohol-related variables to check the robustness of the deindustrialisation hypothesis against the alcohol policy hypothesis, which suggested that the excess deaths in the 1990s resulted from the relative drop in alcohol price after the Gorbachev anti-alcohol campaign ended. We obtained information on industrial employment, death rates, income, age structure, number of inhabitants, migration, health care provision, housing conditions, alcohol price and total alcohol consumption from the Federal Statistics Service (Rosstat) yearbooks. Supplementary Table B6 presents summary statistics, while Supplementary Table B7 shows the variable definitions for the town-level data used in analyses covering Russia.

[5] We did not include the unemployment variable due to the unreliability of this measure in the Russian Federation during 1990s. Many formally employed workers were often sent on unpaid leave for indefinite durations or did not have gainful employment while being *de facto* employed. It has been widely recognised that these records vastly underestimated the extent of unemployment in Russia during the transition years (Standing, 1996; Grogan and van den Berg, 2001). The Hungarian data on registered unemployment is more reliable, and the practice of sending employed workers on unpaid leave was uncommon in the country in the 1990s.

4.2 Estimation strategy

For Hungary, we fit two types of models. First, we examine the relationship between deindustrialisation and mortality using multilevel discrete-time survival analysis, with individuals (level 1) nested in towns (level 2).[6] The dependent variable in the multilevel models in Hungary reflects the odds of dying in the 52 towns during the 1989–1995 period, adding dummy variables for each year (fixed effects).[7]

The primary independent variable is deindustrialisation. To capture the contextual, non-contemporaneous effect of deindustrialisation in the multilevel models, we measure deindustrialisation as a change in industrial employment (manufacturing, mining and construction) from 1989 to 1995, expressed as a percentage of each town's total population. The mean value of deindustrialisation in the 52 towns was 41%, and the median was 40%. We group towns into two categories: those where deindustrialisation equals or exceeds 50% (n = 16) and those below 50% (n = 36). This approach to defining deindustrialisation as a contextual factor is widely used in the extant literature on deindustrialisation and health (Mitchell *et al.*, 2000; Rind *et al.*, 2014).

We use a series of covariates to control for town-level heterogeneity. First, the towns' population size could also influence mortality and the effect of modernisation and industrialisation; therefore, we include a variable in our models measuring the mean number of inhabitants. The age composition of towns could also influence the primary association. To account for any effect, we use the towns' dependency ratio (ratio of those 0–14 and those over 64 to those of working age) to proxy for age structure. We also control for the number of unemployed people and use the number of general practitioners as a proxy for the towns' health infrastructure. We control for the towns' death rate in 1989 to reduce the potential for selection bias. We also assess the robustness of the deindustrialisation–mortality association against the initial level of industrial employment. We control for subjects' age, gender, education, smoking, alcohol consumption and marital status at the individual level. All models include additional controls for the subjects' relationship with the respondent to account for potential bias due to the survey design.

Complementing the multilevel models, we also fit models using only town-level administrative time-series data, with dummies for towns and years (two-way fixed effects) and standard errors clustered on towns, estimated through Ordinary Least Squares. In these panel models, the independent variable is continuous, measuring the annual deindustrialisation compared to 1989, using the same data as constructing the contextual deindustrialisation variable in the multilevel models. The dependent variable is the annual town-level all-age male death rate (number of male deaths/100,000). The panel covers the same 52 towns with 310 observations (town-years) and is strongly balanced, with only two missing data points. We use a similar set of control variables as in the case of the multilevel models. The models filter out towns' time-invariant characteristics by design, and we also included year dummies to filter out unobserved time-variant heterogeneity.

[6] We specify random intercepts and random coefficients, allowing for the coefficient of deindustrialisation to vary across towns. We rely on pseudo maximum-likelihood estimation with robust standard errors clustered at the level of towns.

[7] This is the established practice in discrete-time survival analysis (Rabe-Hesketh and Skrondal, 2012). Survival analysis uses a 'censored' outcome variable. In our case, individuals who died after 1995 or were still alive at the date of the questionnaire (2014 or 2015) were censored in 1995.

We use only ecological (regional- and town-level) data for Russia, fitting two-way fixed-effects models with clustered standard errors. The dependent variable measures town-level annual all-age deaths per 100,000 population (both genders). The primary independent variable is the cumulative change in the town-level industrial employment-to-population ratio (1991 as baseline). In our sample of 514 middle-sized towns, the median decrease in industrial employment between 1991–99 was 39%. Only 13% of towns suffered no reduction in industrial employment.

We fit models covering the period from 1991 to 1999. We control for annual real wage change measured in 1991 roubles, the number of inhabitants in 1,000s, dependency ratio, net of out and in-migration per person, per-person floor space, the number of physicians per 100,000 and hospital beds per 10,000. We conduct additional analyses to assess whether the association between industrial employment and mortality is robust to alcohol policy. First, we check the correlation between industrial employment and alcohol price with pure alcohol consumption, using the same dataset and control variables as above.

We carried out all statistical analyses using STATA 16.0. Despite the multiple robustness checks against potential selection bias and unobserved heterogeneity, establishing causality is beyond the scope of our work. We hope the correlational patterns identified will inspire future work designed to assess causal connections.

5. Results

5.1 An overview of post-socialist deindustrialisation

Figure 3 summarises the key trends of labour market transformation in Hungary and Russia. While overall employment declined by 24% in Hungary between 1986 and 1995, industrial employment fell by 43% in the same period, the years of the country's most pronounced liberalisation measures. This extreme deindustrialisation was very severe and rapid by international standards. The most deindustrialised American metropolises, such as Philadelphia, Cleveland and Chicago, lost around 30% of their manufacturing labour force between 1972 and 1987 (Wallace *et al.*, 1999, p. 115). Representing one of the worst cases globally, it took 30 years for deindustrialisation to reach 60% in West Central Scotland between 1971 and 2005 (Walsh *et al.*, 2009). In comparison, within less than a decade, almost every second person employed in manufacturing in Hungary lost their job. This represents a massive shock to the social fabric, whose wide-ranging implications have not received much attention until recently.

Russia experienced a steep decline in output and wages, but industrial employment declined over a more extended period. Total employment fell by 15% between 1987 and 1999, while industrial employment dropped by 38% in the same period. Economic adjustment to the new competitive capitalist environment took place more through the income channel in Russia, leading to a more pronounced decline in aggregate income and wages, while income in Hungary fell less, and adjustment took place more through the employment channel (Boeri and Terrell, 2002; Gimpelson and Kapelyushnikov, 2011). Russian policy-makers allowed companies to keep their workers while reducing and delaying their salaries.

5.2 Deindustrialisation and mortality in Hungary

Table 1 presents the results of multilevel survival models for Hungary, showing that deindustrialisation is robustly and statistically significantly associated with male mortality

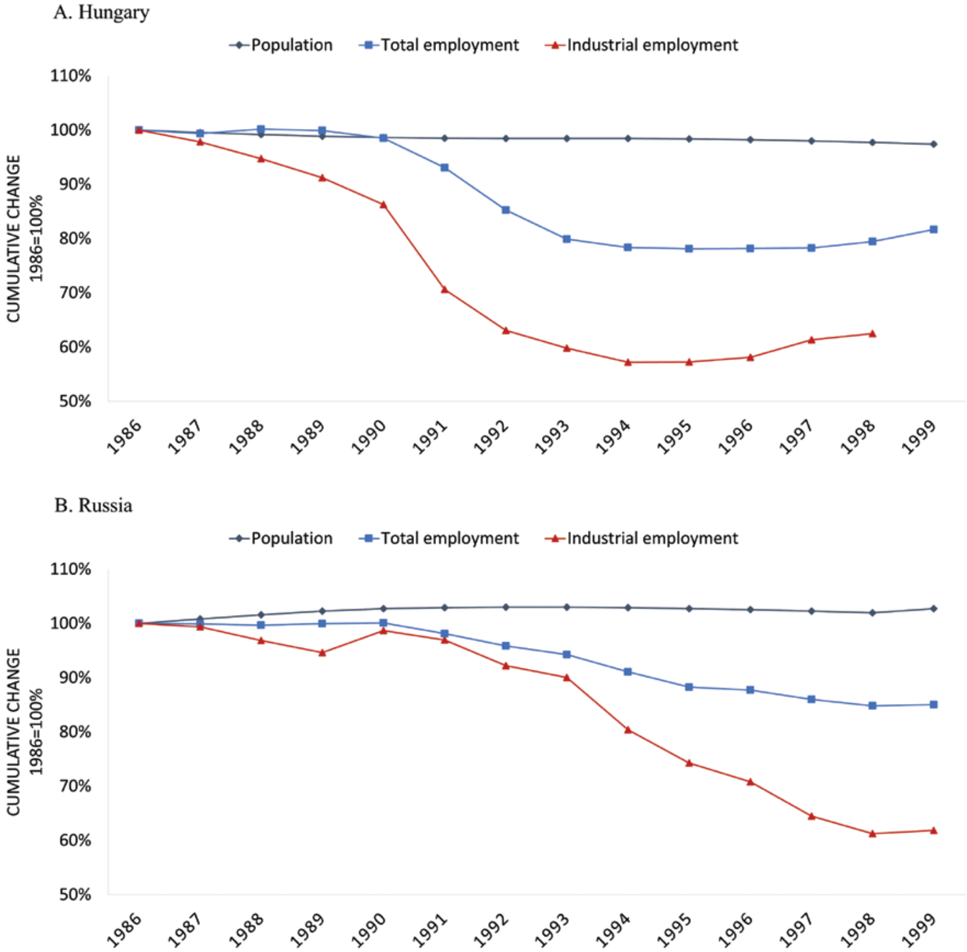

Fig. 3. *The deindustrialisation of Hungary and Russia. Panel A: Hungary. Panel B: Russia.* *Note*: Hungary: industrial employment includes mining and excludes construction. Russia: industrial employment includes mining and construction. *Source*: Hungary: Population and employment: Feenstra *et al.* (2019), manufacturing employment: Brada *et al.* (1994), and Laky (2000). Russia: Estimated from: Obzor zanyatosti Rossii (1991-2000), Issue 1, 2002. Moscow: Teis; Goskomstat SSSR. 1980–1989. Narodnoe khoziaistvo RSFSR: statisticheskii ezhegodnik. Moscow: Goskomstat.

in each specification. The results for women were not significant; see Supplementary Table C1. Therefore, we concentrate on men in the subsequent models. The constant in model 1 shows the unadjusted odds of dying in the 1989–1995 period (0.025), and the variance of the constant at the town level is 0.017, suggesting large differences across towns, underpinning the multilevel modelling strategy.[8] Next, we included the

[8] In multilevel modelling, this is called an 'empty model' or 'variance components model', regressing the odds of dying on the intercept, which is simply used to partition out the variance across the levels.

Table 1. *Deindustrialisation and male mortality in Hungary, 1989–95, multilevel survival models*

Dependent variable	Subject dying between 1989 and 1995						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Deindustrialisation 89–95 (ref.: Moderate (<50%))							
Severe (≥50%)		1.146*	1.182*	1.182*	1.141***	1.188*	1.149**
		(0.066)	(0.078)	(0.078)	(0.043)	(0.091)	(0.052)
Individual-level control variables							
Age in 1989 (centred)					1.066***		1.066***
					(0.002)		(0.002)
Education (ref.: Primary)							
Secondary					0.855***		0.853***
					(0.040)		(0.040)
Tertiary					0.678***		0.675***
					(0.060)		(0.061)
Smoking (ref.: Quit or never)							
Regularly					1.510***		1.517***
					(0.059)		(0.060)
Alcohol (ref.: Max 1–4 times a month)							
Daily or several times a week					1.229***		1.227***
					(0.039)		(0.038)
Town-level control variables							
Average population 89–95 10,000 persons (centred)						0.989	0.995
						(0.010)	(0.014)
Average unemployment 89–95 (centred)						0.985	0.972**
						(0.009)	(0.010)
Average dependency ratio 89–95 (centred)						1.008	1.011
						(0.008)	(0.007)
Death rate 89 per 1,000 (centred)						1.003	0.992
						(0.013)	(0.011)
Constant	0.025***	0.024***	0.024***	0.024***	0.017***	0.024***	0.016***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Random effects							
Town: var(Constant)	0.017*	0.013*	0.006	0.006	0.015	0.004	0.008
	(0.008)	(0.006)	(0.005)	(0.005)	(0.009)	(0.005)	(0.005)

Table 1. *Continued*

Dependent variable	Subject dying between 1989 and 1995						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Town: var(Deindustrialisation)			0.031*	0.031*	0.017*	0.028**	0.006
			(0.014)	(0.014)	(0.008)	(0.010)	(0.004)
Prob > Chi²	.	0.018	0.012	0.000	0.000	0.000	0.000
Year fixed effects	No	No	No	Yes	Yes	Yes	Yes
Type of relative	No	No	No	No	Yes	No	Yes
No. of observations (person-years)	139,211	139,211	139,211	139,211	139,211	139,211	139,211

Notes: Multilevel survival models with random coefficient for deindustrialisation. The reference category for the dependent variable is the subject not dying between 1989 and 1995. Coefficients reported as odds ratios, with cluster-robust SEs in parentheses.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

primary independent variable, deindustrialisation. In this unadjusted setting, men in severely deindustrialised towns have 14.6% higher odds of dying than men in moderately deindustrialised towns. In model 3, we also specify a random coefficient, that is, allow the effect of deindustrialisation to vary across towns. The effect size increases, and the town-level variance of the constant decreases significantly, suggesting that the random coefficient for deindustrialisation improves the models. The chi-square test also shows an improving model fit. In model 4, we add year fixed effects.

In model 5, we add individual-level controls for age, education, smoking, alcohol consumption and relationship status. In model 6, we add town-level control variables for population size, age structure, initial mortality rates and the unemployment variable.[9] These controls result in a significant reduction of town-level variance. Finally, model 7 includes each control variable simultaneously. In this fully adjusted model, men in severely deindustrialised towns have 14.9% higher odds of dying in 1989–95 than men in moderately deindustrialised towns.

It is also worth mentioning that working-class men with primary education (the reference category) have 32.5% higher odds of dying than those with a college degree, net of the effect of other factors. The only town-level control variable that reaches statistical significance is registered unemployment. A 1% increase in registered unemployment (thus the number of people receiving unemployment benefits) is associated with 2.8% lower odds of dying among men in 1989–95.[10]

Figure 4 shows the predicted marginal probability of dying in the 1989–95 period by the extent of deindustrialisation (using industrial employment change deciles) adjusted according to model 7. The association between deindustrialisation and mortality is non-linear. People living in the most severely deindustrialised decile of towns have a 28% higher probability of dying than those living in the least deindustrialised decile.

Next, we carry out a 'placebo test' to rule out the potential that pre-existing mortality differentials may have played a role. The models with town-level controls already included death rates in 1989, showing no association with mortality in 1989–95. However, our placebo test offers a more robust guarantee against selection bias. We construct a separate dataset using subjects who lived in 52 towns during the 1980s before deindustrialisation began in 1989. We use this dataset to investigate the town and individual-level determinants of subjects' death between 1985 and 1988, including the deindustrialisation level after 1989.

Supplementary Table C2 shows that we found no statistically significant pre-existing mortality differences among men living in the towns that later underwent different industrial transformations. Comparing the coefficient for deindustrialisation reported in Table 1 (model 7) and the coefficient obtained through the placebo test, we can ascertain that the 'treatment effect' is significantly different from the 'placebo effect'.[11] The Z-score for the difference between the two coefficients is 2.529, corresponding to $p = 0.006$ (one-tailed test). This means that the effect of deindustrialisation is significantly different in the 'treatment group' (1989–95) than in the 'placebo group' (1985–88).

[9] Outmigration, the number of GPs and town-level income were not significantly associated with the dependent variable, so we removed them from the analysis to avoid multicollinearity. Supplementary Table D3 presents the association between these additional town-level control variables and mortality.

[10] We offer an interpretation of this finding in the discussion section.

[11] We can use a Z-test to compare coefficients obtained through maximum-likelihood estimation (Brame *et al.*, 1998).

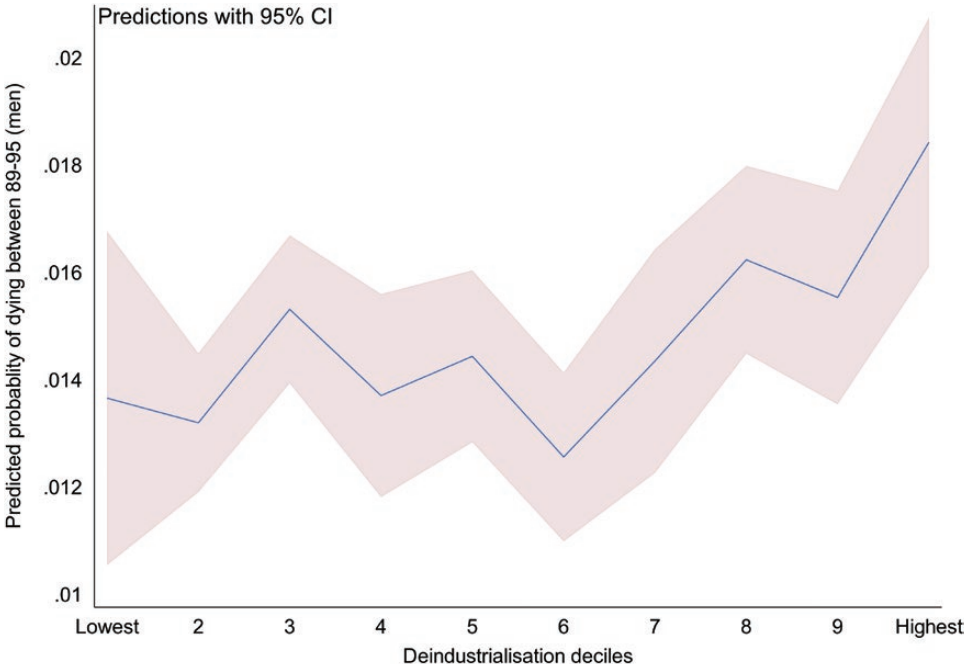

Fig. 4. *Predicted probability of men's dying by deindustrialisation in 52 towns of Hungary, 1989–95.* *Note*: Predicted probabilities from logistic regression with clustered standard errors. The dependent variable is subject dying between 1989 and 1995 (ref.: subject not dying between 1989 and 1995), the primary independent variable is deindustrialisation (in deciles), adjusted for town-level (total death rate in 1989, average unemployment in 1990–95, and average dependency ratio 1989–95), and individual-level control variables (smoking, alcohol consumption and education). Number of observations at (person-years): 139,211, number of groups (towns): 52.

Proponents of the dysfunctional working-class culture thesis argue that hazardous drinking is inherited from the past due to socialist industrialisation that led to normative disorientation and unhealthy lifestyles among socialist workers (Cockerham, 1997; Carlson and Hoffmann, 2011). In Hungary, there was no anti-alcohol campaign, so the increased affordability of alcohol in the post-campaign period should not play a major role. Therefore, to assess the explanatory power of the dysfunctional working-class culture thesis, we analyse how industrial employment in 1989 influences the association. The correlation between initial industrial employment and deindustrialisation is very low ($r = -0.04$). Thus the initial level of industrial employment does not seem to explain subsequent deindustrialisation.

Next, we analyse whether industrial employment in 1989 influences mortality differentials in multilevel regression models. We split the sample of 52 towns in half, above and below the median industrial employment in 1989. According to the dysfunctional working-class culture hypothesis, a higher share of industrial employment should correlate with higher mortality in 1989–95. As we report in model 7, Supplementary Table C3, we find the opposite. Men in severely deindustrialised towns with below-median industrial employment in 1989 have *higher* odds of dying between 1989 and 1995 than men in towns with above-median industrial employment in 1989, where the effect of deindustrialisation is not significant. This means that the positive association of deindustrialisation with mortality is not a spurious effect of initial industrial employment. These results contradict the dysfunctional industrial culture hypothesis.

Figure 5 shows the unadjusted correlation between town-level mortality and deindustrialisation in Hungary. This visual inspection of the raw data reveals a potential association at the town level, with a bivariate correlation of $r = 0.314$.

Table 2 shows that the town-level two-way fixed-effects panel regressions confirm the multilevel model results. Deindustrialisation is robustly associated with male mortality in every model. Model 4—controlling for towns' size, age structure, unemployment, per capita income, outmigration and the number of GPs—shows a highly significant association ($p < 0.01$); 1% deindustrialisation correlates with 1.74 additional male deaths per 100,000 inhabitants.

5.3 Deindustrialisation and mortality in Russia

As described above, we investigate the correlation between deindustrialisation and health in Russia using town-level fixed effects modelling covering mid-sized towns in European Russia. Table 3 presents the regression results.

The association between deindustrialisation and mortality is positive and highly significant in all four models. We start by fitting regressions with year dummies and town

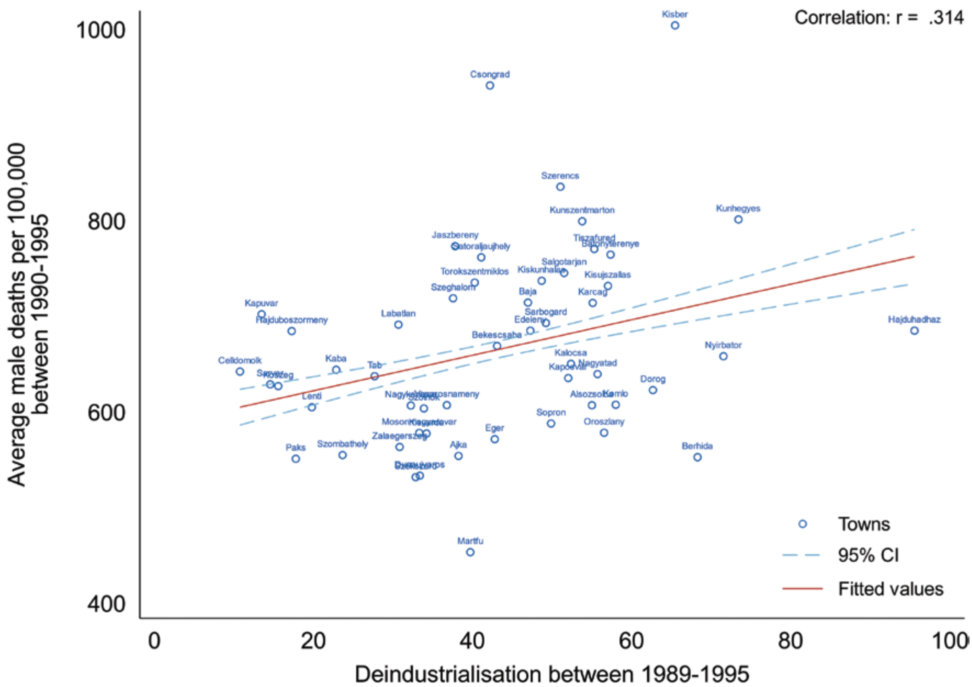

Fig. 5. *The town-level association between mortality and deindustrialisation in Hungary.*
Note: Unadjusted, bivariate correlation. Deindustrialisation is measured as the percentage decline in industrial employment ratio from 1989 to 1995. Number of observations (towns): 52.

fixed effects. Given declining trends in industrial employment, we can interpret the coefficients in model 1 as follows: a decrease in industrial employment of 1% was associated with 0.407 ($p < 0.01$) more deaths per year per 100,000 population in the 1991–99 period. This estimate is not attenuated by accounting for per capita income and square metres of accommodation in model 2. The association becomes even stronger after controlling for the number of inhabitants, dependency ratio and per capita net migration in model 3. Finally, in the fully adjusted model 4, a 1% deindustrialisation is associated with 0.516 additional deaths per year per 100,000 population ($p < 0.001$). Figure 6 shows the predicted town-level male deaths by deindustrialisation in Russia, suggesting a strong association.

The proponents of the dysfunctional culture approach questioned the idea that stress caused by economic dislocation would be an essential determinant of hazardous drinking, arguing for the centrality of dysfunctional culture amplified by the high affordability of alcohol in Russia. If this hypothesis were valid, we would expect (a) a positive association between industrial employment and alcohol consumption and (b) that controlling for alcohol price eliminates the association between deindustrialisation and alcohol consumption. Table 4 shows the coefficients and standard errors for a percentage share of industrial employment predicting alcohol consumption for this period in models 1 and 3 (full model). To facilitate cross-model comparison, we also show the effect of alcohol price on its consumption in a bivariate

Table 2. *Deindustrialisation and male mortality in Hungary, 1989–95, town-level fixed-effects panel models*

Dependent variable	Male deaths per 100,000 population			
	(1)	(2)	(3)	(4)
Cumulative deindustrialisation from 1989 (%)	1.193*	1.526*	1.697**	1.737**
	(0.580)	(0.581)	(0.578)	(0.589)
Income per capita 10,000 HUF (centred)		8.720		1.663
		(6.339)		(7.010)
Unemployment (centred)		−0.344		−0.256
		(2.480)		(2.618)
Population 10,000 persons (centred)			−1.138	−3.322
			(101.033)	(101.330)
Dependency ratio (centred)			−13.201**	−12.444*
			(4.592)	(5.343)
Outmigration % of population (centred)			1.381	1.796
			(11.701)	(11.949)
No of GPs per 10,000 inhabitants (centred)			8.496	7.559
			(11.314)	(11.896)
Constant	644.139***	878.826***	748.686***	785.867***
	(13.109)	(180.021)	(34.478)	(182.541)
Year fixed effects	Yes	Yes	Yes	Yes
R^2	0.037	0.049	0.071	0.071
No. of observations (town-years)	310	310	310	310
No. of groups (towns)	52	52	52	52

Note: Coefficient estimates from town-level two-way fixed-effects panel models, cluster-robust SEs in parentheses.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

model 2 and in a model with industrial employment share, price of alcohol and other covariates, namely wage in 1991 roubles, number of inhabitants in 1000s and per-person floor space.

Contrary to the expectations derived from the dysfunctional culture hypothesis, the association between industrial employment and alcohol consumption is significant and negative in both specifications, indicating an increase of 0.013 l per person per year with every 1% decrease in the industrial employment-per-population ratio. Thus, deindustrialisation seems to be associated with increased alcohol consumption, filtering out the effect of alcohol price. This suggests that hazardous alcohol consumption was a stress-coping strategy, above and beyond the effects of alcohol price. The correlation of price with alcohol consumption is also significant in both models; however, the coefficient has the opposite sign, as suggested by the dysfunctional culture argument. In full model 3, each ruble of increase in alcohol price correlates with 0.017 l ($p < 0.001$) more pure alcohol consumption.

Table 3 *Deindustrialisation and mortality in Russia, 1991–99, town-level fixed-effects panel models.*

Dependent variable	Deaths per 100,000 population			
	(1)	(2)	(3)	(4)
Deindustrialisation (%)	0.407*** (0.105)	0.394*** (0.106)	0.520*** (0.095)	0.516*** (0.096)
Income per capita (in 1991 RUB) (centred)		0.100** (0.034)		0.147*** (0.037)
Floor area per person (centred)		1.103 (1.425)		0.646 (1.478)
Population 10,000 persons (centred)			−37.807** (11.826)	−35.995** (11.399)
Dependency ratio (centred)			1.583*** (0.389)	1.420** (0.409)
Migration % of population (centred)			−37.355 (41.380)	−38.062 (41.216)
Physicians per 100,000 (centred)			−0.551 (0.461)	−0.410 (0.462)
Constant	1,193.056*** (6.359)	1,198.453*** (7.009)	1,205.301*** (7.334)	1,210.978*** (8.102)
R^2	0.189	0.180	0.239	0.230
No. of observations [town-years]	4,527	4,450	4,02S9	3,968
No. of groups [towns]	514	513	493	492

Notes: Deindustrialisation is defined as the cumulative change in the industrial employment-to-population ratio (1991 as baseline). All models include controls for year fixed effects (year dummies)—Cluster-robust SEs in parentheses.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

6. Robustness checks

First, we assess the stability of multilevel modelling results using a different estimation strategy. Logistic regression is sensitive to model specification and can yield biased results compared to linear models (Mood, 2009). Therefore, we re-estimate regressions from Table 1 using multilevel linear predicted probability modelling. As Supplementary Table D1 shows, the positive association between deindustrialisation and male mortality remains robust and significant in each model.

We also re-estimate the multilevel association between mortality and deindustrialisation, operationalising the latter as a categorical variable with three and four categories. These results shown in Supplementary Table D2 again confirm the main findings. Supplementary Table D3 shows that the fully adjusted association also holds when we measure deindustrialisation as a continuous annual industrial employment ratio (mean = 17.5, standard deviation [SD] = 8.5). In this specification, the effect is even bigger; a 50% higher industrial employment ratio implies 40% lower odds of dying.

Supplementary Table D4 shows multilevel modelling estimates with additional town-level control variables (income per capita, no of GPs, industrial employment 1989). With these new covariates, the effect of deindustrialisation is even bigger (OR = 1.193, $p < 0.001$). Industrial employment in 1989 was measured as a continuous

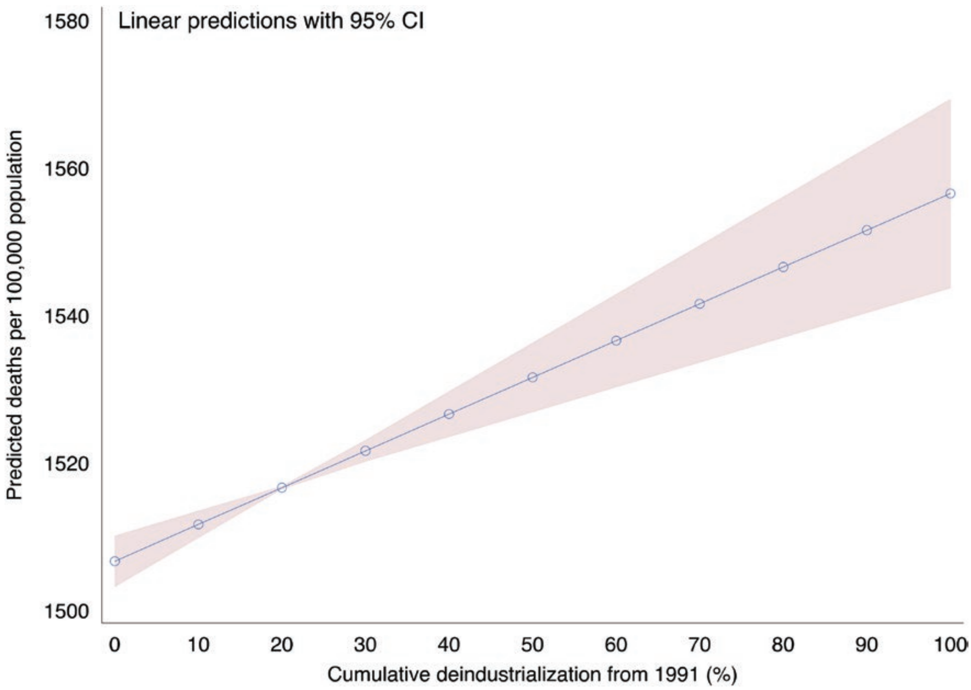

Fig. 6. *Predicted town-level male deaths by deindustrialisation in 490 towns of Russia, 1991–99.* *Note*: Linear predictions from town-level fixed-effects panel models adjusted for income per capita, floor area per person, population, dependency ratio, migration, number of GPs and unobserved time-varying heterogeneity (year fixed effects), and the dependent variable lagged by 1 year. Number of observations (town-years): 3,589, number of groups (towns): 490.

variable (min.: 0.1, max.: 0.63), expressing the share of the total population employed in industry in percentages. The association is negative but not significant, again suggesting that our main results cannot be explained by the dysfunctional culture hypothesis, according to which this association should be significantly positive.

Next, we check against another potential source of selection bias stemming from migration. Towns that lose younger or healthier people would exhibit a higher mortality rate. We test for in- and outmigration during the analytical period and following the analytical period until the individual surveys were conducted. As Supplementary Table D5 shows, none of the migration-related variables is significant, and their inclusion does not influence the significance of the effect of deindustrialisation, which is even bigger in this setting (1.143 ≤ OR ≤ 1.215).

We also test the dysfunctional culture hypothesis by splitting the town-level sample into half (above and below-median industrial employment in 1989) end estimating the panel models as specified above this way. The results in Supplementary Table D6 again are inconsistent with the dysfunctional culture thesis by showing that the effect of deindustrialisation on mortality is particularly severe in towns with lower industrial employment at the end of the 1980s. Supplementary Figure D1 confirms the same for Russia; the association between the change in death rate and initial industrial employment is flat. The superimposed regression lines demonstrate that the initial level of industrialisation does not correlate with subsequent changes in death rates.

In Supplementary Table D7, we re-estimate the panel models with a lagged dependent variable. In the fully adjusted model 6 with lagged male death rates, the

Table 4. *Industrial employment and alcohol consumption in Russia, 1991–999, town-level fixed-effects panel models*

Dependent variable	Regional-level annual alcohol consumption (litres of pure alcohol per capita)		
	(1)	(2)	(3)
Industrial employment (%)	−0.013***		−0.013**
	(0.004)		(0.004)
Alcohol price, deflated (centred)		0.023***	0.017**
		(0.006)	(0.006)
Income per capita (in 1991 RUB) (centred)			0.0005*
			(0.0002)
Population 10,000 persons (centred)			−0.039
			(0.070)
Floor area per person (centred)			0.010
			(0.08)
Constant	5.178***	4.948***	5.259***
	(0.081)	(0.037)	(0.086)
R^2	0.410	0.417	0.421
No. of observations[town-years]	4,532	4,495	4,417
No. of groups[towns]	514	514	513

Note: Deindustrialisation is defined as the annual industrial employment-to-population ratio in percent. All models include controls for year fixed effects (year dummies)—Cluster-robust SEs in parentheses.
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

association between deindustrialisation and death rates is even stronger (b = 2.342, p < 0.001). The same holds for results using data on Russian towns. Supplementary Table D8 shows that the deindustrialisation–mortality association remains significant in Russia when including a lagged dependent variable (b = 0.499, p < 0.001 in the fully adjusted model 4). Next, we measure deindustrialisation as a continuous annual industrial employment ratio (mean = 16.2, SD = 8.2). These panel models imply first differences in this case and not cumulative change. Supplementary Table D9 shows that the association between the industrial employment ratio and death rates is robustly and strongly negative in Hungary, confirming the positive association between deindustrialisation and mortality.

Finally, we test the multilevel and panel results against potentially influential outliers. First, we identify outliers in Supplementary Figure E1 for Hungary, and Supplementary Figure E2 for Russia, by checking the deviation of town-level death rates from the overall sample average. Supplementary Table E1 shows that multilevel model estimates for Hungary are robust when excluding the three towns with the highest and lowest mortalities. Supplementary Table E2 shows the same for town-level panel models. Supplementary Table E3 excludes 20 towns with the lowest mean death rates and 20 towns with the highest mean death rates in 1991–99 in Russia. A significant mortality effect remains for deindustrialisation. Supplementary Table E4 shows that Hungarian panel models remain significant and robust when including data from 1996 and 1997, while Supplementary Table E5 shows the same when excluding 1989 and 1990. In Supplementary Table E6, we retain a significant mortality effect for deindustrialisation in Russia, excluding 1998 and 1999. Finally, Supplementary Table E7 shows that the same holds when excluding 1991 and 1992.

7. Discussion and conclusions

In this article, we utilised an innovative multilevel dataset on the economic determinants of mortality in Eastern Europe. We showed that deindustrialisation was significantly associated with male death rates in Russia and Hungary in the 1990s, but social and economic policies have offset Hungary's more immense industrial employment loss. Furthermore, we showed that the association between deindustrialisation and the post-socialist mortality crisis is not a spurious result of a legacy of dysfunctional working-class health culture aggravated by low alcohol prices during the early years of the transition. Even if dysfunctional working-class culture played a role, our results suggest that researchers should not pit culture against political–economic dislocations. Hazardous drinking was, in part, a dysfunctional coping strategy in response to a highly stressful experience. To understand this process, we proposed a political–economic theory of deindustrialisation and mortality based on heterodox economics and economic sociology. This conceptual framework identifies deindustrialisation as a process of social disintegration rooted in the lived experience of neoliberal shock therapy. Deindustrialisation acts as a source of stress and despair through individual-level (income loss, job loss, precarity) and community-level channels (loss of sense of community, increased inequality, loss of company-related services), leading to reduced coping ability and ill health.

Our study has some important limitations. The individual-level sample used for the multilevel models in Hungary is constructed from the close kin of those who were still living in the specific town in 2014–15. This data collection procedure excludes those families that were in the town from 1989 to 1995 but not when the interviews were conducted. To address this weakness, our models controlled for outmigration and in-migration. Furthermore, the United Nations has validated and recommended this indirect demographic method for censuses (United Nations, 1983). Previous studies conducted in Russia using a similar method also showed results consistent with the official data and the extant literature (Bobak *et al.*, 2003; Azarova *et al.*, 2017).

In response to the much-debated study by Ruhm (2000), which found that mortality decreases during economic recessions, Arthi *et al.* (2017) showed that mortality increased during a recession that they analysed after adjusting for migration. The fact that adjusting for migration turns the recession–mortality association positive implies that if migration biases such associations, it does so because less-healthy people migrate away from affected areas. It is plausible to assume that people negatively affected by deindustrialisation migrated in search of better conditions, while those who could keep their jobs stayed in the town. While this runs contrary to the oft-quoted healthy migrant effect, it is consistent with some populations moving away from adverse economic conditions, such as the Irish in the 19th and 20th centuries (Delaney *et al.*, 2013). Although we cannot robustly rule out such a bias, we find no evidence for it, and even if it exists, it makes our estimates more conservative than they really are.

Second, the retrospective cohort method might also be prone to recall bias. People might err a few years when recollecting when their relative died. We also found subjects' answers about the occupation and place of work of their relatives unreliable. Thus, we excluded these variables from the analysis. Because of recall bias, we could not generate an annual death variable for fine-tuned year-on-year mortality risk analysis and had to rely on an indicator variable encompassing the 1989–95 period. Therefore, we could not generate a panel dataset using individual data. However, our town-level fixed-effects panel models using administratively collected death rates from two countries confirmed the multilevel models.

Third, the towns involved in the analysis might differ in unobservable characteristics in both countries. We tackled this limitation by controlling for pre-existing health differentials, suggesting that the towns did not differ in terms of health before deindustrialisation. Furthermore, the fixed-effects models, by definition, filter out unobserved time-invariant heterogeneity. We also controlled for time-variant global factors by including year fixed and tested our models against the inclusion of a lagged dependent variable. We could also control for the most critical characteristics identified in the literature. Nevertheless, unobserved heterogeneity might still be an issue, and the associations cannot be interpreted as causal links.

Theoretically, it is possible to handle selection and heterogeneity bias statistically. However, designing such a quasi-experiment for our study faces insurmountable challenges. There is either no administrative data available for the pre-treatment period or what is available is not reliable. For example, neither the Eurostat nor the IMF publishes comparative national-level data on social spending before 1995 for most countries in post-socialist Eastern Europe. The same data availability limitation is even more severe at the town level. Fine-grained digitalised statistics from the socialist era are a rarity, and if they exist, collecting them from non-digital archives is costly. While mortality data from the socialist period are reliable, the quality of information on town-level social and economic characteristics used in the present paper is hard to ascertain.

All in all, our research question, and the data available to answer it severely limit the viability of credible quasi-experimental designs. However, we do not think that quasi-experimental methods are magic bullets, as they have some important limitations (see Cook *et al.*, 2002). First, what quasi-experiments gain in internal validity, they frequently lose in external validity. The more they resemble an experiment, the more they can identify causality, but the less generalisable they become outside the study population. Second, quasi-experiments rest on assumptions about some exogenous source of 'quasi-random' treatment variation that cannot be proven statistically. Third, researchers' pre-analytical visions and 'meta-scientific' commitments irreconcilably influence the questions they ask, the variables they define and the associations they seek. Fourth, quasi-experiments, even in the most highly ranked journals, are at risk of p-hacking[12] that distorts the overall picture emerging from the research, leaving much room for researchers' preferences to influence the results (Brodeur *et al.*, 2020). As Kuhn (2012) famously argued, these pre-analytical visions are so important that science typically does not change by gradually accumulating knowledge but through scientific revolutions as incommensurable paradigms succeed each other.

Altogether, while quasi-experimental methods bring significant added value to research, they should not devalue studies of phenomena for which it is very difficult, if not impossible, to design a credible quasi-experiment. As Angus Deaton (2020, p. 1) emphasises in his criticism of randomised control trials: 'It is a mistake to put method ahead of substance…. Methodological prejudice can only tie our hands. Context is always important, and we must adapt our methods to the problem at hand'.

The fourth limitation is that our samples are restricted to medium-sized towns in both countries, curtailing the generalisability of our results. For example, the detrimental effect of deindustrialisation could be buffered in big cities due to the broader availability of jobs in non-industrial branches of the economy. Smaller towns and villages experienced a different shock related to the collapse of agriculture, which would require a different sample. Also, working-class culture likely differs across cities of various sizes. If people living in large cities, who are omitted from our sample, are more prone to hazardous drinking, then our models underestimate its effect. However, if hazardous drinking is more severe in small- and medium-sized industrial towns, then our models magnify its effect compared to the whole population of the two countries.

Even though such differences across the towns might exist, we showed that the mortality trend in our sampled towns in Russia mirrors the national mortality trend. This parallel trend suggests that large cities omitted from our sample did not experience significantly different mortality shocks. We also took account of the towns' population size in every regression model to reduce the bias arising from the size of the towns. Nevertheless, our results cannot be directly generalised to small villages and the largest cities. Notwithstanding the limited generalisability from our sample of medium-sized towns, we find it remarkable that the models behave consistently in the Hungarian and Russian samples.

Finally, during the transformation from socialism to capitalism, many policies and institutions changed simultaneously. This makes it difficult to isolate deindustrialisation, even if this was desirable, given the implausibility of a single causal factor. However, several of these policies, such as political–institutional change, did not vary between towns, only between countries, and thus should not bias our results. Following the extant literature, we controlled for health service variables, which were not significant and

[12] P-hacking 'occurs when researchers try out several statistical analyses and/or data eligibility specifications and then selectively report those that produce significant results' (Head *et al.*, 2015, p. 1).

did not influence the primary association (Brainerd and Cutler, 2005). Yet, this should not be interpreted as suggesting that the post-socialist health system was functioning well, which it was not (Andreev *et al.*, 2003). Other elements of shock therapy—such as bankruptcy policies, rapid import opening, capital account liberalisation and rapid mass privatisation—could have played a role, and deindustrialisation likely picks up some of their effect (Amsden *et al.*, 1994; Tregenna, 2008).

Notwithstanding these limitations, our results align with several bodies of research. First, we found that deindustrialisation was correlated only with men's health in the short run, echoing the existing literature on the health effect of privatisation in Russia (Azarova *et al.*, 2017). The gendered expectations of the male breadwinner model could have made men more vulnerable to labour market upheaval, especially because men were overrepresented in industrial jobs. However, men might benefit more from the new jobs replacing old industrial ones in the long run. Women in towns with more privatisation had higher odds of dying compared to women in towns with higher state ownership in the 1995–2004 period in Hungary (Scheiring *et al.*, 2018b).

Second, concerning the weakness of the alcohol affordability argument, others also found that the increase in the price of vodka led to a rise in the consumption of illicit and surrogate alcohol (Goryakin *et al.*, 2015). A recent study by Azarova *et al.* (2021) also showed that the correlation between the Gorbachev anti-alcohol campaign in the 1980s, alcohol prices in the 1990s, and mortality reported by Bhattacharya *et al.* (2013) and Treisman (2010) is likely spurious. Furthermore, death rates improved in the second half of the 1980s and worsened in the 1990s in several countries, such as Hungary, that did not have a Soviet-type alcohol campaign that led to a large alcohol price increase in the 1980s and a subsequent fall in the 1990s (Ellman, 1997b, p. 358).

Third, the negative association between registered unemployment and mortality likely reflects the cushioning impact of unemployment benefits. The maximum unemployment benefit duration in Hungary between 1989 and 1993 was 24 months, with the unemployed receiving 70% of their previous gross earnings during the first 6 months of unemployment (Vodopivec *et al.*, 2005). Hungary's privatisation strategy was also more gradual and orderly than Russia's rapid mass privatisation and the attendant disintegration of bureaucratic capacities. These social and economic policy differences likely played an important role in offsetting deindustrialisation's health effect in Hungary.

The negative association between registered unemployment and mortality also supports our hypothesis that the adverse health effect of deindustrialisation goes beyond job loss. There were many people in Hungary (and in other countries in the region) who lost their job and did not register as unemployed. Registered unemployment only reflects the net balance of those who are actively looking for a job through formal state institutions, thus overlooking the issue of high labour market turnover, high inactivity, as well as the community-level implications of deindustrialisation. Industrial employment loss captures these multidimensional ripple effects, while registered unemployment does not.

In contrast to Hungary, the Russian welfare state was less developed, and its bureaucratic capacity was especially severely affected by mass privatisation and transformation (Hamm *et al.*, 2012; Irdam *et al.*, 2015). The unemployment insurance system was practically non-existent (Clarke, 1998, pp. 50–52). The average unemployment benefit in 1993 was 12% of average pay, 18% of average pay in 1994, and the average length of benefit was 5 months. As a result, only 12% of unemployed people received any help in Russia in 1992. The welfare state's weakness, the chaotic mass privatisation programme, the precarity of employment, and the more severe income loss could explain why Russia experienced a more severe wave of excess deaths.

Of course, these two countries were already different before the transition in terms of initial conditions, such as the extent of industrial distortions, the size of the economy and non-state sector, and the structure of safety nets, among others. Future research should look in more detail at how these pre-existing differences conditioned policy choices that influence how deindustrialisation affects health. As King *et al.* (2022) argue, comparative historical research has much to add to our understanding of the causes of mortality crises. We hope our results will inspire further work to compare countries within the region and explore the similarities and differences between the post-socialist mortality crisis and deaths of despair plaguing the North American rustbelt.

The insights emerging from our results are relevant for other countries beyond post-socialist Eastern Europe. As Case and Deaton (2020, p. 108) noted, 'it is no exaggeration to compare the long-standing misery of these Eastern Europeans with the wave of despair that is driving suicides, alcohol, and drug abuse among less-educated white Americans'. Our study adds an important element to such comparisons. Robust welfare provisions, active labour market policies and community regeneration programmes are necessary to offset the adverse health effect of deindustrialisation (Venkataramani *et al.*, 2021). Well-designed strategies, such as regionally targeted industrial policies, can create jobs and contribute to the health of vulnerable populations (Rodrik, 2004). A green new deal and just transition policies (Newell and Mulvaney, 2013) can steer society towards a lower carbon future underpinned by equity, justice and workers' health in rustbelt areas.

Supplementary data

Supplementary data are available at *Cambridge Journal of Economics* online.

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