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Does temporary employment increase length of commuting? Longitudinal evidence from Australia and Germany

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Abstract

On average, temporary jobs are far less stable than permanent jobs. This higher instability could potentially lower workers' incentives to relocate towards the workplace, thereby resulting in longer commutes. However, surprisingly few studies have investigated the link between temporary employment and commuting length. Building on the notion that individuals strive to optimize their utility when deciding where to work and live, we develop and test a theoretical framework that predicts commuting outcomes for different types of temporary workers – fixed-term, casual and temporary agency workers – and in different institutional contexts. We estimate fixed-effects regression models using 17 waves of data from the Household, Income and Labour Dynamics in Australia (HILDA) Survey and the German Socio-Economic Panel (SOEP). As expected, the results show that the link between temporary employment and commuting length varies by employment type and institutional context. Agency work is associated with longer commutes than permanent work in both countries, whereas this applies to fixed-term contracts for Germany only. For casual work, the findings suggest no commuting length differential to permanent employment. In terms of policy, our findings suggest lengthy commuting can be a side effect of flexible labour markets, with potentially negative implications for worker well-being, transportation management and the environment.

Keywords Work-related spatial mobility · Fixed-term contracts · Casual work · Temporary agency work · HILDA Survey · SOEP

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Introduction

Temporary forms of employment, such as fixed-term contracts, casual work and temporary agency work, are an entrenched feature of many Western labour markets. Across EU-28 countries, for example, the share of temporary workers among employees aged 20 to 64 averaged 10.8% in 2019, and was in excess of 15% in several countries. In Germany, however, the share was slightly below the EU average (9.3%) (Eurostat 2020). In contrast, in Australia, the comparable fraction is much higher. Our analysis of data from the Household, Income and Labour Dynamics in Australia (HILDA) Survey shows that about 29% of Australian employees were employed on either a fixed-term or casual basis in 2018 (and unlike in Europe, it is casual employment that is numerically dominant). Furthermore, work commutes have increased in terms of distance and time in many industrialised countries over recent decades (e.g., American Association of State Highway and Transportation Officials 2013; Federal Institute for Population Research 2018; Office for National Statistics 2015; Wilkins et al. 2019). In Germany, the proportion of workers commuting 60 min or longer (round trip) has increased from 20% in 1991 to 27% in 2016 (Federal Institute for Population Research 2018). In Australia, the corresponding shares are even higher, and have risen from 38% in 2001 to 47% in 2017 (Wilkins et al. 2019).

Despite the pervasiveness of both temporary employment and lengthy commutes, the connection between the two has received relatively little attention from researchers. This is surprising given previous literature has suggested that the length of employment contract, and thus perceived job security, likely affects incentives to move closer to the workplace, ultimately resulting in longer commutes among temporary workers compared to permanent workers (Abraham and Nisic 2007; Crane 1996; Parenti and Tealdi 2019).

Investigating the relationship between temporary employment contracts and commuting length is important on both an individual and societal level: At the individual level, both temporary contracts and long commutes have been linked to adverse worker outcomes. For instance, researchers have found temporary employment associated with lower wages (Kahn 2016; Laß and Wooden 2019; OECD 2015), greater job insecurity (Auer and Danzer 2016; Dawson et al. 2017; Green and Leeves 2013) and lower levels of job satisfaction (Aleksynska 2018). Also, lengthy commutes have been shown to be associated with increased stress and decreased subjective well-being (e.g., Chatterjee et al. 2020; Evans and Wener 2006; Royal Society for Public Health 2016) as well as reduced job satisfaction (Chatterjee et al. 2017; Clark et al. 2020; Novaco et al. 1990). To the extent that temporary employment and lengthy commuting occur together, we may thus observe an accumulation of disadvantages for the affected workers. At the societal level, trends in temporary employment may reinforce trends in commuting length, with potentially negative implications for public health, transportation management and the environment. According to Eurostat data on 12 EU-countries, commuting to work accounts for between 27% and 47% of individuals' total distance travelled per day, thereby constituting the most important travel purpose for urban mobility (Eurostat 2021).

Against this background, our study investigates the link between temporary employment and length of commuting in Australia and Germany. We improve on existing research in several ways, both with regards to theory and empirical analysis. First, in contrast to previous studies (e.g., Abraham and Nisic 2007; Crane 1996; Parenti and Tealdi 2019), we provide a detailed theoretical framework that generates hypotheses about expected differences in commuting length, not only between temporary and permanent workers overall but also between different types of temporary employment and within different country contexts.

Second, previous quantitative empirical studies have exclusively considered fixed-term contracts. Several of these studies have linked fixed-term contracts to longer daily commutes and/or a higher likelihood of weekend commuting compared to permanent contracts in a range of Western European countries (Abraham and Nisic 2007; Lück and Ruppenthal 2010; Parenti and Tealdi 2019; Rüger and Sulak 2017; Schneider and Meil 2008). That said, other studies have found differing results by country (Viry and Vincent-Geslin 2015), socio-economic groups (Kersting et al. 2021) or gender and family status (Wachter and Holz-Rau 2021). Our study goes beyond this narrow focus on fixed-term contracts by investigating three different types of temporary employment, namely fixed-term contracts, temporary agency work, and casual work (for Australia). This is important given these employment types show some marked differences in characteristics, and especially their level of workplace stability.

Third, in contrast to all mentioned studies, we apply fixed-effects regression analysis, allowing the elimination of the influence of time-constant unobserved worker characteristics. This is potentially of large importance given workers may select into employment types based on unobserved characteristics that may also affect commuting length, such as ability, ambition and preferences. For example, those who are more ambitious may have both a higher willingness to commute and a higher likelihood of working in permanent jobs. Not accounting for such factors may therefore result in biased estimates.

Fourth, we provide a more thorough analysis of the commuting information than previous studies. With one exception (Abraham and Nisic 2007), previous studies used binary or categorical, and thus rather crude, measures of (long-distance) commuting (Kersting et al. 2021; Lück and Ruppenthal 2010; Parenti and Tealdi 2019; Viry and Vincent-Geslin 2015; Wachter and Holz-Rau 2021). By contrast, we make full use of our detailed commuting information by specifying commuting length as a metric variable, measured in kilometres and minutes.

Finally, we extend the Western European focus of previous research through a comparison of Germany with Australia. This comparison is promising on both theoretical and practical grounds. From a theoretical standpoint, country comparisons of social phenomena using micro-level data provide additional insights to single-country studies. Most obviously, they allow an assessment of whether results depend on the country-specific macro-structural context (Aisenbrey and Fasang 2017; Geishecker et al. 2010; Kohn 1987; Lillard 2021). Australia and Germany differ considerably with respect to labour markets, welfare regimes and geographical structure. Most importantly, employment protection legislation (EPL) for both permanent and temporary workers is considerably stricter in Germany than Australia (OECD 2020b; 2020c). Australia also stands out internationally through the explicit legal recognition and high prevalence of casual employment (Campbell and Burgess 2001; McVicar et al. 2019). Overall, we expect country differences to moderate the relationship between temporary employment and commuting. However, finding similar results in both countries, despite the institutional differences, would suggest a more robust, general link between temporary employment and commuting. From a practical standpoint, we can make use of comparable high-quality, large-scale household panel studies for both countries—the German Socio-Economic Panel (SOEP) and the HILDA Survey—enabling rigorous, longitudinal analysis.

Theory and hypotheses

Basic theoretical framework

Building on the rational choice paradigm to explain human behaviour (e.g., Becker 1965; Boudon 2003), we consider commuting to be the result of a deliberate decision about where to live and where to work: Among all possible combinations of place of residence and place of work, individuals strive for the one from which they derive the greatest (lifetime) utility (e.g., Simpson 1980; Kalter 1994; van Ommeren et al. 1999). The abundance of parameters, however, means this decision is complex: Besides the various benefits and costs associated with work location (e.g., income) and residence (e.g., amenities attached to a particular place), the commuting distance inherent in the combinations is a central factor in the underlying cost-benefit calculation. The costs of daily or weekly commuting include material costs, in the form of travel expenses or establishing a second home, and immaterial costs due to the loss of time for alternative beneficial activities, like spending time with the family, and psychological or physical strain during the commute (Chatterjee et al. 2020; van Ommeren and Fosgerau 2009). While some positive aspects of commuting are discussed in the literature (e.g., Lyons and Chatterjee 2008), in general, people dislike long commutes and therefore try to avoid them. Chatterjee et al. (2020), in reviewing the literature, found that commuting is the daily activity rated with the lowest positive and highest negative affect scores, and that longer commute duration is associated with lower commuting satisfaction. Corroborating this, German employees reported to be willing to give up 13.2% of their earnings for reducing a 45-minute commute to 15 min (Nagler et al. 2022). Moreover, workers often adjust their lives to avoid long-distance commuting. For example, research has repeatedly found that the propensity to move closer to the workplace increases with the distance between home and work (e.g., Clark et al. 2003; Petzold 2020). Furthermore, Wilkins et al. (2019) show that only 49% of Australian workers with lengthy commutes (defined as two hours or more per day) in one year are still lengthy commuters in the next. By contrast, 71% of workers with short commutes (i.e., less than one hour per day) are still short commuters in the next year.

A shorter commute can be accomplished by adjusting either the place of work or the place of residence. Such changes, in turn, typically incur other types of costs that must be considered in the overall calculation. The costs of moving the home (i.e., moving costs) are related to setting up a new (family) residence (e.g., search costs) and/or the loss of location-specific capital; for instance, if friends and relatives can no longer be seen on a daily basis (Amundsen 1985; Fischer and Malmberg 2001). The main costs of a workplace change are the loss of firm-specific human capital and privileges linked to seniority (Farber et al. 1993; Mitchell 1983; Topel 1991).

It is obvious that the complexity of the decision regarding the best combination of place of residence and place of work increases as the number of options increases. Indeed, simultaneously evaluating a virtually unlimited number of home and work locations and the associated costs of commuting and changing home or work environments would exceed human information processing capacity (Kalter 1994). Therefore, many models are based on the assumption that actors typically take one component – e.g., the workplace – as given and focus on the decision about the location of the other component – in this example, the place of residence. Many studies thus rely in their theoretical modelling on the assumption that workers decide between different residence alternatives based on a given place of work (e.g., Alonso 1964; White 1988; Clark et al. 2003). This approach has repeatedly been justified by the argument that it is usually more difficult for employed persons to find a new place of work than a new place of residence (e.g., Abraham and Nisic 2007; Pfaff 2017). In this scenario, workers decide to change their place of residence if they are aware of an alternative work-home combination that, after deducting moving costs, is expected to provide a higher net benefit than their current combination. This assessment may be based on better housing quality, lower housing costs or a shorter commute to work.

Based on this model of residential choice, some key assumptions about the influence of temporary forms of employment on commuting length can be developed. Housing quality and costs being equal, people will decide to move their home closer to their workplace if the sum of the mobility costs associated with that move (i.e., moving costs plus future commuting costs after the move) is expected to be lower than the future commuting costs without the move. Future commuting costs are, however, not only a matter of residence choice, but also depend on future changes in work location. If people have to search for a new job again after moving, the distance to the workplace is likely to increase again. More generally, the higher the probability of a future workplace change, and hence, the more short-term the expected benefits resulting from reduced commuting, the more likely it is that people will choose not to move and accept time-consuming commuting. Given temporary contracts typically reduce perceived job security, we expect such contracts to discourage workers from moving closer to their workplace, thereby resulting in longer commutes for temporary workers compared to permanent workers (see also Crane 1996).

However, one can also think of the opposite situation: individuals make decisions regarding their place of work given a place of residence (Simpson 1980; Roberts and Taylor 2016). This includes the question of whether to keep or change a job after a (non-work-related) relocation. In this scenario, the commuting costs are weighed against a variety of job-related characteristics. Crucial job characteristics are both monetary (e.g., wages) and non-monetary rewards (e.g., occupational prestige, meaningful and intrinsically satisfying work, and pleasant and safe working environments) (Daw and Hardie 2012; Kalleberg 1977). Temporary jobs are widely regarded as inferior in terms of job characteristics compared to permanent jobs (Dekker and van der Veen 2017; McGovern et al. 2004; OECD 2014). Hence, utility-oriented workers may choose jobs with temporary contracts over jobs with permanent contracts if this sacrifice of utility is compensated for by shorter commutes. With the same rationality, longer commutes could be accepted in favour of permanent contracts. This argument, which is based on the theory of compensating differentials (Smith 2007 [1776]), implies an opposite effect to the one discussed above: temporary work contracts would be associated with shorter commutes.

Overall, both mechanisms are theoretically possible. However, previous literature on this topic has been unanimous in assuming that the difference in the incentives for temporary and permanent workers to change their residence is the decisive mechanism (Abraham and Nisic 2007; Crane 1996; Kersting et al. 2021; Parenti and Tealdi 2019; Rüger and Sulak 2017). This conjecture is supported by the fact that, as mentioned, empirical studies have mostly found fixed-term contract work to be positively associated with commuting length. We thus put forward this first hypothesis:

H1: Temporary employment will be associated with longer commutes than permanent employment.

Differences by type of temporary employment

Casual work. A key characteristic of casual work is that the employer does not need to provide workers any advance commitment about the duration of the employment contract or the specific number of days or hours per week to be worked (Creighton and Stewart 2010: 198). Perceived job security can therefore be expected to be relatively low among casual workers, suggesting they have less incentive than permanent workers to move closer to the workplace. However, for several reasons, casual workers may still not be more likely to have lengthy commutes: First, casual jobs are often low-paid and low-skilled (Green and Leeves 2013; Laß and Wooden 2019), and our own analyses of HILDA Survey data show they often involve particularly short shifts. Thus, many workers may be unwilling to accept a casual job that would necessitate lengthy commutes or relocation because the job's benefits would not outweigh the mobility costs. Second, van Ham et al. (2001) argue that low-skilled jobs are less spatially dispersed than high-skilled jobs, suggesting that workers may often be able to get a casual job close to home. In this case, the compensating differentials argument discussed above could apply, meaning some workers would accept a casual contract in exchange for shorter commutes. Overall, there are arguments for both shorter and longer commutes among casual workers. We thus expect the commuting length gap between casual and permanent workers to be relatively small on average.

Fixed-term contracts. These types of contracts stipulate a date or event at which they will be terminated, which reduces job security compared to permanent contracts. In contrast to casual work, fixed-term contracts exhibit relatively high qualification requirements in the two investigated countries. In Australia, fixed-term jobs tend to be particularly highly skilled and highly paid (Laß and Wooden 2019). In Germany, both high-skilled and low-skilled workers have a higher likelihood of receiving a temporary contract than mid-skilled workers, but with high-skilled workers most affected (Gebel and Giesecke 2011). This is in contrast to many other European countries, where low-skilled workers are most likely to receive fixed-term contracts (Gebel and Giesecke 2011). Given the relatively high skill level, and correspondingly high pay, connected to fixed-term contracts in Australia and Germany, accepting these types of jobs may often be worthwhile, even if located at a considerable distance from home. Furthermore, as mentioned, high-skilled jobs are spatially more dispersed, increasing the need for work-related spatial mobility. Overall, distant fixed-term job offers can be expected to be equally likely to be accepted as distant permanent job offers; however, because of lesser job security, relocation will be less frequent—and lengthy commutes thus more frequent—for fixed-term positions. Compared to the other temporary employment types considered here, however, the workplace stability associated with fixed-term employment is relatively high, suggesting the incentive to relocate is larger for workers in fixed-term contract jobs than for those with casual or agency jobs. Whereas casual workers can be dismissed anytime, fixed-term contract workers have a reasonable expectation that their employment will last at least until the day stipulated in the contract—often several years into the future. And, in contrast to agency workers (see below), the workplace is unlikely to change over the duration of the fixed-term contract. In addition, studies on labour market transitions (Fuller and Stecy-Hildebrandt 2015; McVicar et al. 2019; Watson 2013; see also Table S2 in the Online Supplementary Material) suggest that the chances of subsequently receiving a permanent contract are higher for fixed-term workers than for casual and agency workers. Fixed-term contract workers may thus be more likely than other temporary workers to move towards the workplace in anticipation of their contract subsequently being converted to permanent.

Temporary agency work. The key characteristic of agency work is the tripartite relationship between the employer (the agency), the worker, and the host company (to which the labour service is provided). Job tenure with the agency is usually relatively short, as many agency workers are employed on either casual or (often short) fixed-term contracts (see Data and Methods section). Additionally, frequent changes of workplace due to short assignments with different host companies are common (Arrowsmith 2006; Storrie 2007). Agency workers are thus subject to an extreme type of workplace instability, and as a consequence will usually not consider relocating closer to work, but commute from their (fixed) place of residence to their various workplaces. We thus expect even longer commutes on average for agency workers than for fixed-term contract workers.

Overall, we expect the following differences by employment type:

H2 Among the group of temporary workers, temporary agency workers will have the longest commute, followed by fixed-term contract workers and then casual workers.

Country differences

Australia and Germany show substantial differences with regard to labour markets, welfare regimes and geographical conditions. These may moderate the strength of the association between employment type and commuting length through at least two channels.

The first channel works via country differences in average commuting length. The commuting length differential between workers on temporary contracts and those on permanent contracts can be expected to be larger in contexts with greater average commuting lengths. This is because if commutes are generally short, most workers will not consider moving closer to the workplace, regardless of their type of contract. As mentioned previously (see Introduction), the share of workers with substantial commuting times is much larger in Australia than in Germany. This may be due to several country differences. For example, Australia is much larger in size than Germany, which could result in Australians being more likely to accept jobs at a great distance from home. Additionally, most Australians live in vast capital cities,[1] which are associated with particularly long commutes (Wilkins et al. 2019). Furthermore, social security payments in the case of short- or mid-duration unemployment are much lower in Australia than in Germany (OECD 2020a). Australian workers may therefore feel more pressure than German workers to be employed continuously, which may increase their willingness to accept distant jobs. However, although average commutes are thus longer in Australia than in Germany, this may not necessarily result in a stronger

[1] The three largest Australian cities – Sydney, Melbourne and Brisbane – alone account for half of the Australian population (calculations based on Australian Bureau of Statistics 2019). This compares with around 8% of the German population living in Berlin, Hamburg or Munich in 2017 (calculations based on Statistisches Bundesamt 2019).

relationship between employment type and commuting length in Australia. This is because exceptionally high housing costs in Australian city centres (Smith et al. 2021) may often prohibit commute-reducing moves, for both permanent and temporary workers.

The second channel through which institutional differences may moderate the link between temporary contracts and commuting length is by affecting the (relative) levels of perceived job security and workplace stability of temporary and permanent workers. Germany and Australia differ considerably with respect to both the degree of EPL and the difference in the level of EPL between permanent and temporary workers, i.e., the "EPL gap" (e.g., Barbieri and Cutuli 2016). Over the 2001–2018 period, Germany scored relatively high on the OECD EPL index with respect to the individual dismissal of permanent workers (an average of 2.6 out of 6 points; OECD 2020b) but relatively low on EPL for temporary workers (1.2 points; OECD 2020c), resulting in an EPL gap of 1.4. This reflects relatively high job security for permanent workers compared to temporary workers. In Australia, average EPL for both permanent (1.5 points) and temporary jobs (0.9 points) has been relatively low over the period, giving an EPL gap of 0.6 and thus smaller differences in job security between permanent and temporary workers. The larger job security gap between permanent and temporary workers in Germany suggests larger differences in the utility from moving closer to the workplace, which should result in a larger gap in commuting length.

Furthermore, in Australia, casual work is explicitly provided for in industry awards[2], and its use is legitimised further by the requirement that casual employees receive a wage premium (typically 20% prior to 2010 but gradually raised to a standardised 25% by 2014) (Laß and Wooden 2019). As a result, casual employment has become the prime vehicle for short-term labour supply in Australia. In contrast, there is no formal provision for casual employment in German law, which means short-term workers will usually be employed on fixed-term contracts. Therefore, we may expect the average duration of fixed-term contracts to be longer in Australia than in Germany, which suggests a higher utility of moving closer to work for Australian fixed-term contract workers.

Overall, differences in labour market regulation suggest larger gaps in perceived workplace stability between permanent and temporary employees in Germany than in Australia. Given the crucial role of workplace stability for the decision between relocating and commuting, these differences suggest that the gap in commuting length between permanent and temporary workers will be larger in Germany. We thus put forward the following hypothesis:

H3: The relationship between fixed-term contracts and agency work on the one hand, and commuting length on the other, will be stronger in Germany than in Australia.

Also deserving consideration is gender regimes and corresponding gender differences in both labour market and commuting behaviour. Australia and Germany are both marked by the modified male breadwinner model, in which men work full-time in their role as primary earners and women work part-time to accommodate their role as the primary carers, particularly if there are children (Pocock 2005; OECD 2017; Trappe et al. 2015). As in many other countries, Australian and German women typically spend much more time on housework than their male counterparts (Altintas and Sullivan 2016). Likely as a consequence of these different roles, numerous studies from various countries, including Australia and Germany,

[2] Industry awards are legally enforceable industry-wide determinations, made or assented to by industrial tribunals, which set minimum wages and employment conditions for most employees.

have shown that, on average, women have shorter commuting distances and times than men (e.g., Crane 2007; Fan 2017; Wilkins et al. 2019; Giménez-Nadal et al. 2022).[3] Furthermore, women are less likely than men to initiate moves to foster their own professional careers when they are living in a partnership (Abraham et al. 2010; Mincer 1978). Given these important gender differences, we carry out all analyses separately for male and female workers.

Data and methods

Data and samples

The data used come from two long-running household panel surveys, the SOEP (v34) (Liebig et al. 2019) and the HILDA Survey (General Release 18) (Department of Social Services and Melbourne Institute of Applied Economic and Social Research 2019). While the SOEP started in 1984, the HILDA Survey commenced in 2001 and was designed to mimic practices of successful household panel studies, and notably the SOEP. The two surveys thus share many common features in terms of sampling methodology and questionnaire design. For example, both surveys interview all adult household members on an annual basis and collect comprehensive information on respondents' employment situation.

For both countries, the analysis is based on a seventeen-year period, namely 2002–2018 for Australia and 2001–2017 for Germany. The first wave of the HILDA Survey (2001) was omitted due to differences in the way commuting time was collected. For the SOEP, the waves 2014 and 2016 had to be excluded given commuting distance is unavailable in these waves. The sample used here is restricted to workers aged between 18 and 64 years. This yields a sample of 44,806 employed individuals (contributing 205,430 observations) in the SOEP and 22,158 individuals (152,458 observations) in the HILDA Survey. From this general sample, we only retained those who were not undertaking any form of study or training. Analyses of the data show that students are much more likely to be employed on temporary contracts than other workers, but at the same time their choice of place of residence and workplace is more restrained as they are usually tied to the place of study. We excluded cases with missing information on commuting length as well as outliers with extremely long commutes (longer than the 99.95th percentile). In the HILDA Survey, these are daily commuting times of more than 8.3 h, and in the SOEP, commuting distances of more than 830 km. Additionally, we excluded SOEP respondents who were unable to report a commuting distance because they work in different locations. Note that our sample includes persons who work from home as well as workers who do not travel from their (primary) residence to their workplace on a daily basis (such as weekend commuters). We also excluded observations with missing information on our key predictor (employment type) and on any control variables. Finally, we excluded respondents who only contributed one observation given we apply longitudinal methods of analysis.

The final SOEP analytical sample comprises 13,934 men (contributing 79,100 observations) and 14,140 women (78,694 observations). The HILDA Survey analytical sample

[3] It should, however, be noted that the magnitude of this gender commute gap varies across countries (Giménez-Nadal et al. 2022) and with breadwinner, relationship, or parenthood status, among other factors (e.g., De Meester and van Ham 2009; Chidambaram and Scheiner 2020; Kwon and Akar 2021; Kim 2022).

consists of 7,522 men (56,086 observations) and 7,263 women (50,205 observations). Summary information on how our samples compare to the general workforce aged 18 to 64 in the SOEP and HILDA Survey is presented in Table S1 in the Online Supplementary Material. Overall, the characteristics of the analytical samples are very similar to those of the general samples. The share of permanent workers is somewhat higher in our analytical samples, which is due to the exclusion of students and (for the SOEP) of workers with changing workplaces, who are more likely to work in employment types other than permanent employment.[4] The exclusion of students also results in our samples containing fewer single and childless workers.

Method of analysis and analytical strategy

We investigate the link between temporary employment and commuting by means of fixed-effects (FE) regression. This method involves applying pooled ordinary least squares regression to transformed, demeaned data (e.g., Brüderl and Ludwig 2015). The regression equation can be written as:

$$y_{it} - \overline{y}_i = (x_{it} - \overline{x}_i)\beta + \varepsilon_{it} - \overline{\varepsilon}_{i,}$$

where y_{it} is the value of the outcome variable for individual i at time t, $\boldsymbol{x}_{it}$ is a vector of covariates of this individual at that same time point, and $\boldsymbol{\varepsilon}_{it}$ is an idiosyncratic error related to that person at that time. By contrast, $\overline{y}_i$, $\overline{\boldsymbol{x}}_i$ and $\overline{\varepsilon_i}$ are the person-specific means of these variables across all time points. $\boldsymbol{\beta}$ is the vector of parameters to be estimated. The demeaning of the data extracts the variation within persons over time and discards the variation between persons. The influence of all unobserved (time-constant) person-specific traits is thus removed, which is the main advantage of this approach. The downside is that information from workers who never change employment type over the observation period do not contribute to the estimation of the employment type coefficients. However, due to the long-run nature of the panels and the relatively short duration of temporary jobs, considerable shares of our samples are observed changing employment types (see Table S2 in the Online Supplementary Material for yearly transition rates). We run all models separately for women and men given gender differences in both labour market and commuting behaviour in both countries as outlined above. All models are unweighted.

Dependent variable: length of commute

In the SOEP, commuting length is measured as the distance (in km) between the workplace and the place of residence, which includes both daily and weekly commuting. This self-reported information is available on an annual basis (with the exception of 2014 and 2016). In the HILDA Survey, respondents are asked each year how much time they spend on travelling to and from the place of paid employment in a typical week. While in 2001, respondents could report their commuting time only in hours, the possibility of reporting minutes was added in 2002 and all later waves. As a result, commuting times for 2001 are not strictly comparable, leading us to exclude this wave from the analysis. We created a measure of

[4] Additional analyses show that only 4% of permanent and fixed-term contract workers but 8% of agency workers and 23% of the self-employed report changing work locations.

daily commuting time in minutes by dividing weekly commuting time by the usual number of days worked per week in the main job and multiplying by 60. Note that this leads to an overestimation of daily commuting times for the (small) group of multiple jobholders who commute to their second job on different days than to their main job.

In order to account for the right-skewed distribution of commuting time/distance in both countries, we took the natural logarithm of these measures to approximate a normal distribution. We added one (km/minute) to all values, which ensures that observations with commuting times/distances of zero can be retained in the analysis. Coefficients can then be interpreted as percentage changes in length of commute.

While the SOEP and HILDA Survey measures are not strictly comparable (given one is collecting commuting length as distance and the other as time), we argue that there is a strong correspondence between them. In three waves, the SOEP also collected information on the time spent travelling to work in minutes, providing the opportunity to investigate overlap between the two measures. The correlation between the untransformed values of distance and time was 0.68, and this increased to 0.82 for the logarithmised values (which are used in this analysis). A very close correspondence between measures of commuting distance and time has also been found in other contexts (e.g., Rietveld et al. 1999). These findings suggest that the two measures can be used as proxies for each other, as has previously been done in the commuting literature (e.g., Lee and MacDonald 2003; Sandow and Westin 2010).

Independent variables

Employment type. Information on contract type concerns employment in the main job and is sourced from two broadly comparable questions in the surveys. In the HILDA Survey, employees are asked whether they are employed on a fixed-term contract, a casual basis or a permanent/ongoing basis, with a fourth option “other”. The “other” category had very few cases and was thus discarded.

In the SOEP, respondents are asked whether their employment contract is permanent or fixed-term, with a third option “not applicable, do not have an employment contract”. The respondents choosing this last option mainly belong to three groups: The first and largest group are the self-employed; these are assigned to a separate category. Second, many are “Beamte”, a type of state official that is not issued with an employment contract but in the vast majority of cases hired for life. We re-classified these as permanent workers. Third, a considerable number of employees other than “Beamte” reported not to have an employment contract. These cases are difficult to classify because German law generally does not provide for the possibility of employees without employment contracts. We thus decided to keep this group in its own “no contract” category for the main analysis. However, in a robustness check, we test whether results change if no-contract workers are excluded from the analysis.

In both surveys, another question asks workers whether they are employed through a temporary employment agency. While agency workers can in principle be on any type of contract, in the SOEP, about half are on fixed-term contracts, and in the HILDA Survey, most of them are either on casual or fixed-term contracts. We classify all employees who report being employed through a labour-hire firm or temporary employment agency, regardless of their contract type, as temporary agency workers.

Finally, we create a fifth category for the self-employed and unpaid family workers.

Control variables. We account for a range of worker and job characteristics that might confound the relationship between temporary jobs and commuting length. In terms of worker characteristics, we include age and its square, parental status (represented by a dummy set indicating the age of the youngest resident own child), and relationship status. For the latter we differentiate between cohabiting and married persons (with single persons being the reference category) given previous research has shown that both the gender division of labour and the level of partnership stability differ between these partnership types (Baxter 2005; Dominguez-Folgueras 2013; Laß 2022), which may affect commuting behaviour. We also account for place of residence by including indicators for the sixteen German and eight Australian states, as well as a dummy for persons living in rural areas. We also include a dummy for a severe health condition. For the SOEP, we follow Otterbach et al. (2016) and set this dummy to one if respondents reported to have been officially assessed as severely disabled or partially incapable of work for medical reasons (with the degree of limited functioning rated at 20% or more). For the HILDA Survey, the value is set to one if respondents report the presence of a work-limiting long-term health condition.

In terms of job characteristics, we account for the number of usual weekly working hours (in a quadratic specification) and for a recent job start (by an indicator for workers with a tenure with the current employer of less than four months). We also include a dummy set representing industry with categories taken from the Cross-National Equivalence File (CNEF), which are based on the International Standard Industrial Classification (v3.1).[5] Due to the low number of miners in Germany, we combine mining and construction into one category in the SOEP. Finally, we include indicators for survey year. In the Australian models, we additionally control for multiple jobholding to account for the fact that the original measure of weekly commuting time relates to all jobs, whereas the number of working days is only known for the main job.

Mean values for all variables, differentiated by employment type and country, are provided in Table 1. With respect to commuting, the table shows that, in both countries, temporary agency workers have the longest commutes, averaging 64 min per day in Australia and 33 km one-way in Germany. Differences, however, can be seen with respect to fixed-term contract workers, who have similarly long commutes to permanent workers in Australia, but notably longer commutes in Germany. By contrast, Australian casual workers and German no-contract workers have shorter commutes than permanent workers. Regarding other worker and job characteristics, the table shows that in both countries, all types of temporary workers tend to be younger than permanent workers, are less likely to be married and are more likely to be single and childless. They also tend to work fewer hours per week than permanent workers (particularly Australian casual workers) and to have started their jobs recently. In both countries, fixed-term contract workers are more likely than permanent workers to work in the service industry, and agency workers in manufacturing. By contrast, casual workers are more likely to work in wholesale and retail trade.

[5] The CNEF is a unit-record data file that harmonises a subset of variables from a number of national household panel surveys, including both the HILDA Survey and SOEP (Frick et al. 2007).

Table 1 Sample Characteristics by Employment Type (Mean Values)

	Australia					Germany				
Variable	PER	FIX	CAS	TAW	SE	PER	FIX	NOC	TAW	SE
Commuting distance (km)	--	--	--	--	--	19.40	23.36	9.92	33.04	11.96
Commuting time (min)	57.54	57.48	50.77	64.45	46.63	--	--	--	--	--
Male	0.53	0.45	0.38	0.58	0.68	0.51	0.41	0.27	0.57	0.61
Age	41.40	39.68	39.03	39.10	45.78	43.20	37.37	43.23	39.82	45.88
Relationship status										
Single	0.26	0.29	0.37	0.35	0.16	0.21	0.33	0.20	0.30	0.18
Cohabiting	0.18	0.21	0.18	0.21	0.15	0.12	0.17	0.08	0.15	0.11
Married	0.56	0.50	0.45	0.45	0.69	0.67	0.50	0.72	0.55	0.71
Age of the youngest (resident) child										
No children	0.33	0.41	0.36	0.42	0.19	0.26	0.36	0.18	0.32	0.21
0–3 years	0.14	0.12	0.12	0.13	0.15	0.12	0.15	0.10	0.15	0.13
4–7 years	0.08	0.08	0.09	0.08	0.10	0.11	0.13	0.12	0.09	0.12
8–12 years	0.10	0.09	0.10	0.08	0.12	0.12	0.11	0.15	0.10	0.12
13–16 years	0.08	0.07	0.07	0.07	0.09	0.09	0.07	0.10	0.08	0.09
17 years and older or only non-resident children	0.27	0.23	0.27	0.23	0.34	0.30	0.18	0.34	0.26	0.33
Living in rural areas	0.29	0.31	0.40	0.24	0.36	0.33	0.35	0.41	0.36	0.32
Severe health condition	0.07	0.07	0.13	0.09	0.11	0.06	0.04	0.06	0.05	0.04
Working hours (week)	39.73	39.17	25.23	36.39	40.45	37.99	36.09	23.47	36.12	45.07
Multiple jobholder	0.06	0.09	0.15	0.10	0.08	--	--	--	--	--
Tenure < 4 months	0.05	0.11	0.18	0.29	0.04	0.02	0.13	0.06	0.12	0.02
Industry										
Agriculture	0.01	0.01	0.04	0.02	0.10	0.01	0.01	0.02	0.01	0.05
Energy	0.01	0.01	0.00	0.02	0.00	0.01	0.01	0.00	0.02	0.00
Mining	0.02	0.02	0.01	0.05	0.00	--	--	--	--	--
Construction	0.06	0.04	0.07	0.07	0.20	--	--	--	--	--
Mining or construction	--	--	--	--	--	0.12	0.07	0.11	0.17	0.09
Manufacturing	0.11	0.06	0.07	0.16	0.08	0.17	0.11	0.14	0.23	0.07
Trade	0.13	0.10	0.18	0.06	0.12	0.13	0.15	0.29	0.10	0.18
Transport	0.05	0.03	0.05	0.06	0.04	0.05	0.04	0.04	0.04	0.03
Bank, Insurance	0.05	0.03	0.01	0.05	0.02	0.04	0.02	0.01	0.02	0.04
Other services	0.55	0.67	0.56	0.49	0.41	0.41	0.54	0.30	0.34	0.47
Missing industry	0.03	0.03	0.01	0.03	0.01	0.05	0.06	0.09	0.09	0.08
N (observations)	66,930	8311	12,816	2180	16,054	127,050	10,328	5083	3265	12,068

PER = permanent; FIX = fixed-term; CAS = casual; TAW = temporary agency work; NOC = no contract

Results

Main models

Table 2 Association between Employment Type and Commuting Length – Results from Linear FE Regression

	Australia		Germany	
Variable	Men	Women	Men	Women
Employment type (ref. = permanent contract)				
Fixed-term contract	0.011	0.006	0.062**	0.036*
Casual contract	-0.025	-0.019	--	--
No contract	--	--	-0.073*	-0.101**
Temporary agency work	0.081*	0.125**	0.155**	0.060*
Self-employed	-0.292**	-0.845**	-0.792**	-0.726**
Age	0.044	0.024	0.007	0.012
Age squared	-0.000**	-0.000**	0.000	-0.000
Relationship status (ref. = single)				
Cohabiting	0.062*	0.004	0.080**	0.082**
Married	0.068**	-0.014	0.088**	0.057*
Age of the youngest resident child (ref. = no children)				
0–3 years	0.029	-0.167**	0.022	-0.138**
4–7 years	0.008	-0.190**	0.030	-0.170**
8–12 years	0.036	-0.154**	0.018	-0.180**
13–16 years	0.022	-0.094	-0.006	-0.143**
17 years and older or only non-resident children	-0.007	-0.112	-0.012	-0.118*
Living in rural areas	-0.294**	-0.449**	0.423**	0.240**
Severe health condition	0.014	-0.039	0.008	-0.003
Working hours	0.008**	0.016**	0.019**	0.015**
Working hours squared	-0.000**	-0.000**	-0.000**	-0.000**
Multiple jobholder	0.170**	0.282**	--	--
Tenure < 4 months	0.030	0.015	0.063*	0.006
Industry (ref. = other services)				
Agriculture	-0.144	-0.356**	-0.127*	0.022
Energy	0.134	0.250	0.053	0.061
Mining	0.175**	-0.026	--	--
Construction	0.181**	-0.113	--	--
Mining or construction	--	--	0.071**	-0.036
Manufacturing	0.001	-0.106**	0.041	0.015
Trade	-0.062*	-0.113**	0.025	0.014
Transport	-0.048	-0.061	0.055	0.132**
Bank, Insurance	0.047	0.006	0.021	0.123*
Missing industry	0.018	-0.036	0.030	-0.004
Constant	2.490**	3.016**	1.799**	1.404**
N (observations)	56,086	50,205	79,100	78,694

Models also control for state and calendar year. Bootstrapped standard errors. ** and * denote statistical significance at the 0.01 and 0.05 levels, respectively.

Table 2 presents the results from the FE regression. It shows for both genders in both countries that agency work is linked to significantly longer commutes than permanent employment. For Australian men, for example, working for a temporary employment agency is associated with 8.4% longer commutes compared to having a permanent contract, and for

German men, agency work is associated with an increase in commuting length of 16.8%.[6] For Germany, we also find that fixed-term contracts involve extended commutes compared to permanent contracts — 6.4% longer for men and 3.6% longer for women. By contrast, in Australia, fixed-term contracts are not associated with longer commutes for either gender. Likewise, casual work is not linked to significantly longer commutes than permanent employment; if anything, the sign of the coefficients suggests casual work may be associated with shorter, rather than longer commutes. Self-employment, by contrast, is connected to significantly shorter commutes in all cases.

Additional analyses

We conducted several additional analyses to test the robustness of the results presented in Table 2, both with respect to sample selection as well as choice of control variables (see Table 3 for key results; complete results available in Tables S3 to S6 in the Online Supplementary Material).

With respect to sample modifications, we first excluded observations reporting having no weekly commute at all (i.e., commuting time of 0 min in HILDA) or who report their home and workplace to be in the same building (SOEP) (Model 2). In this analysis, the coefficients for fixed-term contracts in Germany and temporary agency work of both genders in both countries attenuate. The reduction is particularly visible for German women, where both coefficients become insignificant. This finding indicates that the positive association of fixed-term and agency work with commuting length in our main model is not solely due to a reduced propensity to relocate among these workers. Rather, it suggests that this association is at least partly due to permanent jobs more often being pursued from home than temporary jobs. Similarly, the negative coefficients for Australian casual work become larger and, among women, statistically significant in Model 2, which suggests that a lower likelihood of working from home pulls commuting times upwards for casual workers as well. By contrast, the negative coefficients of self-employed workers (and of no-contract workers in Germany) attenuate, suggesting that part of the negative association is due to an increased likelihood of working from home.

Second, we re-ran the analyses after bringing the students, trainees and apprentices back into our sample and adding an indicator for this group to the model (Model 3). Thereby, the negative coefficient for Australian casual work becomes notably larger and statistically significant at the 5%-level for both genders. In Germany, the positive coefficients for fixed-term and agency work become slightly larger for both genders.

Third, we tested whether exclusion of multiple jobholders alters the results for Australia (Model 4). This may potentially be of relevance given our measure of overall commuting time included time spent travelling to all jobs, yet we could only consider the number of working days in the main job when calculating daily commuting times. As can be seen from Table 1, temporary workers are more likely than permanent workers to be multiple jobholders. We find that for men, the negative association between casual employment and commuting length becomes more pronounced upon exclusion of multiple jobholders, but remains statistically insignificant. The results for the other groups of temporary workers remain substantially unaffected.

[6] We calculate commuting length differentials as 100*(exp(β)-1), where β is the estimated coefficient.

Fourth, we tested how the results for Germany change if no-contract workers are excluded from the analysis (Model 5). In this model, the coefficients for fixed-term contracts and agency work change only marginally.

With respect to choice of control variables, we first re-ran the estimations using a more parsimonious model to check for potential over-control (Model 6). Precisely, we omitted the measures of relationship status, the age of the youngest child, as well as the indicator for a rural residence, as these characteristics could themselves be influenced by the employment type and at the same time affect commuting behaviour. This omission of controls did not result in substantial changes in the results.

Second, we omitted job characteristics (i.e., working hours and industry) (Model 7). Again, results remained very similar. The only exception is the negative coefficient for casual work among women, which, due to the strong link between female casual employment and part-time work, becomes larger and statistically significant.

Third, we extended our initial model by adding controls (Model 8). This model additionally accounts for whether a range of different life events that could promote a change in both employment and residence (i.e., getting married, moving in together, partnership dissolution, death of the partner/child, birth of a child, and involuntary job loss) took place since the last wave. Again, no substantial changes in the results were found.[7]

Fourth, we tested whether selective attrition from the panel may influence the results (Model 9). To this end, we re-ran the models after including indicators for whether respondents are non-responding in the next wave and whether respondents have left employment in the next wave. Although some of the coefficients on these indicators were statistically significant, the coefficients on the employment type dummies did not change substantially.

Discussion and conclusion

This article has investigated the link between temporary employment and length of commute in Australia and Germany. We developed and tested a theoretical framework that suggests that temporary jobs should be associated with longer commutes because higher perceived workplace instability reduces the expected benefits of relocating closer to the job. To provide a more comprehensive picture of the link between temporary jobs and commuting than previous studies, we investigated commuting outcomes for three different types of temporary employment in two countries. To the best of our knowledge, this study is also the first to control for the impact of time-constant unobserved worker traits linking contract type and commuting length by applying FE regression to data from two large-scale panel studies.

The results for Germany are in line with our general expectation that temporary employment is connected to lengthier commutes than permanent employment (H1). These results thus confirm and extend those of previous cross-sectional studies from Western Europe linking fixed-term contracts to longer daily commutes (e.g., Parenti and Tealdi 2019; Schneider and Meil 2008). In Australia, however, only one subtype of temporary employment, agency work, was found to involve longer commutes. In contrast, the commuting time linked to

[7] Additional analyses show that the small differences that can be seen between Model 1 and Model 8 are primarily due to the reduction in sample sizes as a result of the exclusion of observations without valid information on the life events in Model 8. Re-running Model 1 with the smaller sample from Model 8 produces very similar results to Model 8 itself.

Table 3 Sensitivity Analyses – Selected Estimates from Linear Fixed-Effects Regression Models

	Sample modifications					Variable modifications			
	Model 1 Main specification	Model 2 Workers with positive commutes	Model 3 Including students	Model 4 Excluding multiple job holders	Model 5 Excluding no-contract workers	Model 6 Without potential mediators	Model 7 Without job characteristics	Model 8 Plus life events	Model 9 Plus indi-cators for attrition
Australian men									
Fixed-term contract	0.011	0.005	0.014	0.005		0.010	0.010	0.011	0.006
Casual contract	-0.025	-0.034	-0.059*	-0.049		-0.031	-0.027	-0.026	-0.020
Temporary agency work	0.081*	0.076*	0.105**	0.073		0.085*	0.079	0.086*	0.074
Self-employed	-0.292**	0.037	-0.308**	-0.293**		-0.295**	-0.299**	-0.289**	-0.279**
N	56,086	53,329	64,097	52,387		56,086	56,086	55,598	51,923
Australian women									
Fixed-term contract	0.006	0.008	0.003	0.028		0.001	0.008	0.009	0.010
Casual contract	-0.019	-0.050**	-0.051*	-0.025		-0.024	-0.048*	-0.018	-0.026
Temporary agency work	0.125**	0.105**	0.095**	0.124**		0.123**	0.123**	0.122**	0.117*
Self-employed	-0.845**	-0.154**	-0.869**	-0.894**		-0.862**	-0.897**	-0.842**	-0.839**
N	50,205	47,569	60,531	45,625		50,205	50,205	49,784	46,445
German men									
Fixed-term contract	0.062**	0.050*	0.070**		0.056*	0.060*	0.055**	0.054*	0.059*
No contract	-0.073*	-0.033	-0.067			-0.070*	-0.118**	-0.076*	-0.072
Temporary agency work	0.155**	0.137**	0.167**		0.148**	0.155**	0.146**	0.137**	0.157**
Self-employed	-0.792**	-0.142**	-0.787**		-0.810**	-0.792**	-0.796**	-0.769**	-0.771**
N	79,100	75,100	86,229		77,715	79,100	79,100	75,840	74,380
German women									
Fixed-term contract	0.036*	0.015	0.049**		0.034	0.035*	0.037*	0.031	0.035*
No contract	-0.101**	-0.052**	-0.102**		--	-0.101**	-0.140**	-0.108**	-0.096**
Temporary agency work	0.060*	0.038	0.062*		0.065*	0.061*	0.060*	0.048*	0.050

Table 3 (continued)

	Sample modifications					Variable modifications			
	Model 1 Main specification	Model 2 Workers with positive commutes	Model 3 Including students	Model 4 Excluding multiple job holders	Model 5 Excluding no-contract workers	Model 6 Without potential mediators	Model 7 Without job characteristics	Model 8 Plus life events	Model 9 Plus indi-cators for attrition
Self-employed	-0.726**	-0.209**	-0.715**		-0.768**	-0.731**	-0.729**	-0.729**	-0.735**
N	78,694	74,651	86,128		74,996	78,694	78,694	75,731	73,366

Model 1 as in Table 2. All other models are variants of model 1: Model 2 excludes workers without commutes. Model 3 includes workers in education and training plus an indicator for this group. Model 4 excludes workers reporting to have more than one job. Model 5 excludes employees reporting to have no employment contract. Model 6 omits controls for partnership status, presence of children/age of the youngest child, and living in a rural area. Model 7 omits controls for working hours and industry. Model 8 includes controls for a set of life events that happened in the past year. Model 9 includes controls for non-response in the next wave and for not being employed in the next wave. Bootstrapped standard errors. ** and * denote statistical significance at the 0.01 and 0.05 levels, respectively.

fixed-term employment is very similar to that of permanent employment, and casual work, if anything, appears to be associated with shorter commutes than permanent employment.

The weaker association between temporary employment and commuting length in Australia was expected (H3) and is in line with national differences in EPL, with the EPL gap between German temporary and permanent workers being much larger than that in Australia. In Australia, fixed-term contracts can be of longer duration than in Germany, and may thus be connected to greater perceived job security, rendering Australian fixed-term contract workers more inclined to move closer to their workplace. The weaker results for Australia are also in line with the assumption that relatively expensive housing in Australian city centres may prohibit many workers, regardless of contract type, from moving closer to their workplace.

We also found that within each country and gender, the strength of the relationship varied considerably between the different types of temporary employment, confirming H2. In both countries, we found the strongest positive association in connection with temporary agency work, which on average is accompanied by an increase in commuting length between 5 and 17%. This strong association can be explained by the fact that agency workers are subject to particularly high levels of job insecurity and frequent changes of their hosts and thus workplaces, rendering the option of moving closer to the workplace impractical in most cases.

For fixed-term contracts, the commuting length gap to permanent employment is smaller than for agency work, with the association ranging between no significant difference (Australian men and women) and about a 6% increase (German men). In contrast to agency workers, fixed-term contract workers may under certain conditions be willing to move to their place of work, particularly if their contract is of relatively long duration or they expect their contract to be renewed or converted to permanent in the future.

By contrast, among the three forms of temporary employment considered for the Australian context, casual work was associated with the shortest commutes. These results are in line with our reasoning that casual jobs are often not sufficiently attractive for workers to accept lengthy commutes and can often be found in the proximity of the home and thus do not require lengthy commutes. Our results for casual employment provide some guidance as to what commuting outcomes one would expect for low-skilled temporary workers in other countries, many of whom are employed on fixed-term contracts outside of Australia and Germany.

Overall, our findings suggest that there is no general effect of temporary employment on length of commute. Rather, the association varies significantly with the employment type and the institutional context. However, temporary agency work proved to be consistently associated with longer commutes in both of these different country contexts.

Our findings also have implications for theory. The results mainly support the theoretical considerations which predicted higher workplace instability will lead to longer commutes among temporary workers due to a lower likelihood of relocation, but also show that this perspective needs to be extended and qualified. First, we show that the degree to which this mechanism applies varies by the degree of (relative) workplace instability (with agency workers being more affected than fixed-term contract workers and German workers more affected than Australian workers). Also, workplace instability discouraging relocations does not seem to be the only relevant factor – our results for casual workers suggest that this type of temporary employment is associated with a lower likelihood of accepting a distant job in the first place, a possibility that has not previously been considered in the literature on temporary employment and commuting. Furthermore, our findings from the additional models suggest that a higher likelihood of permanent workers to work from home also plays a role. Insider-outsider theory (e.g., Bentolila and Dolado 1994; Lindbeck and Snower 2001) would imply differences in bargaining power by contract type, with higher job security putting permanent workers in a better position than temporary workers to bargain for benefits such as telework. This appears to be another mechanism influencing the permanent-temporary commuting length gap that has not been considered previously.

Our study has limitations. First, data availability forced us to use different measures of commuting length for the two countries. Although the two measures are highly correlated and were used as proxies for each other in previous literature, we cannot rule out the possibility that some of the country differences found in our analysis could be due to this difference in measures. Second, while use of FE regression allows us to account for time-constant unobserved worker traits, it still does not enable causal inferences to be made with any confidence as our results may still be affected by selection on (time-varying) unobservables. Nevertheless, our results were robust to the inclusion of additional variables identifying a number of life events, suggesting that omitted variables bias is not a large concern. Third, panel data is associated with attrition, raising the possibility of selection bias. That said, addition of controls for attrition had very little impact on the estimated coefficients on our employment type variables. Fourth, we did not model joint labour market decisions of couples or households.

The findings of this study have implications for policy and practice. First, they suggest that one potential policy avenue for reducing work-related travel (in addition to other measures, such as telework programmes) could be to foster transitions from temporary to permanent employment. Second, in terms of worker well-being, particularly for agency workers, but in part also for fixed-term contract workers, the adverse effects associated with the temporary nature of their employment (such as lower wages and job security) appear to accumulate with the adverse effects of long commutes (such as increased stress and reduced psychological well-being). Temporary workers with lengthy commutes should thus be targeted more strongly by occupational health measures (e.g., for stress prevention and coping).

With respect to implications for transportation research, future studies on the causes and consequences of long commutes should account for workers' form of employment. Conversely, when studying temporary employment, research should consider that some of the adverse effects of temporary employment could be due to lengthy commutes. Future research could also investigate directly the roles of relocations and working from home as drivers of the commuting length gap between temporary and permanent workers. Additionally, potential effect heterogeneities (e.g., with respect to life course phase, living arrangements, and socio-demographic groups) merit further investigation. Furthermore, identifying the role of specific contextual factors in explaining the cross-country differences found in our analysis (such as the EPL gap, country size and housing costs) will be a fruitful avenue for future research. Finally, and related to this point, a replication of this study with other countries would be desirable to enable more general conclusions about the effects of specific types of temporary employment, institutions, and opportunity structures on commuting. Thereby, future comparative studies would benefit from the collection of harmonised panel data on employment and commuting across countries.

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