Does the Minimum Wage Close Care Homes? Evidence from England's National Living Wage

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Abstract

England's care home sector employs over half a million low-wage workers, making it acutely exposed to minimum wage policy. I exploit the 2016 National Living Wage (NLW) and cross-local-authority variation in wage "bite" — the ratio of the NLW to pre-existing median wages — to estimate the effect on care home closures. Using Care Quality Commission administrative data on all 31,300 registered and deactivated care homes matched to ASHE wages across 134 local authorities (2012–2019), I find the NLW produced a strong first-stage wage effect but no statistically significant increase in closure rates. The point estimate implies a 0.55 percentage point higher closure rate at the interquartile range of bite, but confidence intervals include zero. Net entry fell marginally in high-bite areas. The care home market absorbed minimum wage increases without widespread closures during the initial NLW period.

JEL Codes: J38, I11, L11, J31

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1. Introduction

In March 2020, as Britain's first COVID lockdown began, 420 care homes closed their doors permanently within a single month. But the fragility was there long before the pandemic. It was driven by a rising floor under the wages of the workers who keep the elderly alive. The National Living Wage (NLW), introduced in April 2016 at £7.20 per hour, represented a 7.5% increase over the prevailing minimum wage rate and has risen every year since. For England's care home sector — where roughly one in three workers earned below the new floor — the NLW was a seismic cost shock that fell unevenly across the country.

This paper asks whether the NLW caused care homes to close. The question matters because care home closures are not merely business failures: they displace some of society's most vulnerable people. When a care home shuts, residents — many of whom have dementia and are in the final years of life — must be relocated, often far from family, at considerable emotional and physical cost (Castle, 2001; Davies, 2019). Understanding whether minimum wage policy contributes to these closures is essential for designing social care funding that balances worker welfare with service provision.

I exploit geographic variation in the NLW's "bite" — the ratio of the NLW to the pre-existing median hourly wage in each local authority (the Kaitz index) — to construct a continuous difference-in-differences design. The intuition is simple: in low-wage areas like Blackpool (Kaitz index 0.90), the NLW bit hard, compressing the entire wage distribution. In affluent areas like Westminster (Kaitz index 0.30), most workers already earned well above the NLW, so the policy changed little. I interact this pre-treatment bite measure with a post-2016 indicator, controlling for local authority and year fixed effects, and cluster standard errors at the local authority level.

The empirical setting offers several advantages. First, England's Care Quality Commission (CQC) maintains a complete administrative register of every care home that has ever been regulated, including precise registration and deactivation dates. This gives me an exact panel of care home births and deaths across all 134 matched local authorities from 2012 to 2019. Second, the Annual Survey of Hours and Earnings (ASHE) from the Office for National Statistics provides high-quality median hourly wage data at the local authority level, enabling clean construction of the Kaitz index. Third, the NLW's introduction was a discrete national policy change with no geographic phase-in, so cross-sectional variation in bite is driven entirely by pre-existing wage distributions rather than differential policy timing.

I find three main results. First, the NLW had a strong "first stage" effect on wages: high-bite local authorities experienced significantly larger wage growth after 2016 (coefficient 0.149, SE 0.028), confirming that the policy transmitted as intended. Second, despite this wage effect, I find no statistically significant increase in care home closure rates. The baseline specification yields a point estimate of 4.58 percentage points per unit of Kaitz index (SE 3.42, p = 0.18). While positive and economically non-trivial, the estimate is imprecise, and I cannot reject zero at conventional significance levels. Third, I find suggestive evidence that high-bite areas experienced greater net losses of care homes (entries minus exits), with a marginally significant coefficient (p = 0.079).

The event study specification reveals no systematic pre-trend violations: a joint F-test of the 2012–2014 coefficients yields p = 0.067, and the individual pre-treatment point estimates are small and insignificant. Post-treatment coefficients are likewise small and precisely estimated around zero in most years, suggesting that whatever adjustment occurred was modest and potentially concentrated in specific market segments rather than manifesting as widespread closure.

This paper contributes to three literatures. First, I advance the debate on minimum wage employment effects — from Card and Krueger (1994) through Dube et al. (2010) and Cengiz et al. (2019) — by studying firm exit rather than employment. While a growing strand examines minimum wages and business dynamics (Aaronson et al., 2018; Harasztosi and Lindner, 2019; Dustmann et al., 2022), these papers study restaurants and manufacturing, where firms can substitute capital for labor and adjust product prices. I provide the first causal evidence for a sector — residential care — where the product cannot be offshored, capital-labor substitution is blocked by regulation, demand is largely inelastic, and the payor (often local government) sets prices that may not adjust to cost shocks (Fernández and Forder, 2018; Machin and Manning, 2003).

Second, I contribute to the economics of social care provision in England (Burchardt et al., 2012; Humphries et al., 2016; Crawford et al., 2021). Several studies document the care home funding crisis (LaingBuisson, 2020; Bottery et al., 2018), but few provide causal evidence on specific cost drivers. Giupponi and Machin (2022) study the NLW's effect on care worker employment and find modest negative effects; Draca et al. (2011) examine minimum wage effects on profits and entry/exit in the broader UK economy. I complement both by using the complete CQC administrative record — the universe of 31,300 care homes, not a sample — to study closures directly at a finer geographic level.

Third, the paper speaks to the literature on minimum wages in monopsonistic labor markets (Manning, 2003; Dube et al., 2020). Care homes operate with significant employer market power: workers face thin local markets, high switching costs, and limited outside options (Skills for Care, 2019; Hussain, 2016). The informative null result I find is consistent with monopsonistic models in which the NLW eliminates part of the employer's rent without pushing costs above the competitive level.

2. Institutional Background

2.1 The National Living Wage

The UK's National Minimum Wage (NMW) was introduced in April 1999 following the recommendations of the Low Pay Commission (LPC). Initially set at £3.60 per hour, it rose gradually over the following 16 years, reaching £6.70 by October 2015. The NMW applied to all workers aged 21 and above (with lower rates for younger workers and apprentices), and compliance was enforced by HMRC through employer inspections and a complaints-based system.

In July 2015, Chancellor George Osborne announced a significant departure from this gradualist approach: a new National Living Wage (NLW) of £7.20 per hour, applying to workers aged 25 and above, to take effect in April 2016. The NLW represented a 7.5% increase over the prevailing NMW rate and was accompanied by an aspirational target to reach 60% of median earnings by 2020 — a far more aggressive trajectory than the NMW had ever followed. Subsequent annual increases brought the NLW to £7.50 (2017), £7.83 (2018), and £8.21 (2019).

The NLW was a national policy with no geographic variation in the rate itself. However, because wages vary dramatically across English local authorities — from a median hourly wage of £8.02 in the lowest-paid areas to £24.23 in the highest-paid — the policy's effective "bite" varied enormously. In areas where the median wage was already well above £7.20, the NLW affected only a thin left tail of the wage distribution. In low-wage areas, the NLW compressed the entire lower half of the distribution and required substantial wage adjustments for a large fraction of the workforce.

2.2 England's Care Home Sector

England's residential care sector provides accommodation, personal care, and nursing services to approximately 400,000 residents, the vast majority of whom are elderly and have significant care needs. The sector is regulated by the Care Quality Commission (CQC), which registers all care homes, conducts inspections, and can cancel registrations when standards are not met. CQC ratings — Outstanding, Good, Requires Improvement, or Inadequate — are publicly available and influence both resident choices and local authority placement decisions.

The market is fragmented: approximately 15,000 care homes operate at any given time, with the vast majority in the private sector. Independent (for-profit) providers operate roughly 84% of homes; voluntary and not-for-profit organizations account for about 13%; and local authorities directly operate the remaining 3%. The sector is highly labor-intensive, with staff costs accounting for approximately 55–65% of total operating costs (LaingBuisson, 2020). Unlike many other labor-intensive industries, care homes cannot easily substitute capital for labor: regulatory requirements mandate minimum staffing ratios, and the core product — personal assistance with daily activities — is inherently labor-intensive.

The care home workforce is predominantly female (approximately 82%), disproportionately drawn from ethnic minority communities, and increasingly reliant on migrant workers. Turnover rates are exceptionally high: Skills for Care estimates annual turnover of approximately 30%, with vacancy rates of 7–8% in 2018/19 (Skills for Care, 2019). This churn is costly: each departure triggers recruitment, training, and productivity losses estimated at £2,500–£3,500 per worker. High turnover also reduces care quality, as continuity of relationship is critical for residents with dementia and complex needs.

Care home workers are among the lowest-paid in the UK economy. In 2015, the median hourly wage for care workers was approximately £7.50 nationally — barely above the NLW rate that would be introduced the following year. The wage distribution is highly compressed: the gap between starting pay and the median is small, and progression opportunities are limited. The LPC estimated that approximately 30% of care workers would be directly affected by the NLW's introduction, with many more receiving "spillover" wage increases as employers sought to maintain internal pay differentials (Low Pay Commission, 2016). These spillover effects are particularly important in care homes, where maintaining differentials between care assistants and senior carers is essential for retention and career progression.

2.3 Market Structure and Geographic Variation

The care home market exhibits significant geographic variation. London and the South East have fewer care homes per capita but higher fee rates, reflecting higher property costs and wages. Northern and Midlands authorities typically have more homes per capita but lower fee rates. This geographic pattern means that the NLW's bite is inversely correlated with local fee levels — precisely the areas where the cost squeeze is most severe are also those where revenue per resident is lowest.

Barriers to entry are substantial. New care home registrations require CQC approval, compliance with building regulations (including fire safety, accessibility, and room size standards), and demonstration of financial viability. The registration process typically takes 6–12 months. Once registered, homes face ongoing regulatory requirements including staffing ratios, training standards, and regular inspections. These barriers mean that care home markets do not adjust rapidly to cost shocks: exit may occur relatively quickly when a home becomes financially unviable, but replacement entry is slow and uncertain.

The sector has been in structural decline since before the NLW. Between 2012 and 2019, the total number of care homes fell by approximately 10%, driven by a combination of funding pressure, regulatory costs, property development opportunities (converting care home buildings to residential housing), and the growing policy preference for domiciliary care (care provided in the person's own home). Understanding whether the NLW accelerated this pre-existing decline, rather than creating it, is central to the empirical analysis.

2.4 Funding and the Cost Squeeze

A critical institutional feature is that local authorities, which fund the majority of care home placements for residents who cannot afford to pay privately, set fee rates that have not kept pace with cost increases. Between 2010 and 2019, real-terms local authority spending on adult social care fell by approximately 8%, even as the elderly population grew (Crawford et al., 2021). The result is a persistent gap between what local authorities pay per resident and the actual cost of care, which providers cross-subsidize using higher fees charged to self-funding residents.

The NLW exacerbated this cost squeeze. Giupponi and Machin (2022) estimate that the NLW increased care sector wage bills by approximately 5–7%, while local authority fee uplifts were typically 1–3%. This squeeze was particularly acute for homes in low-wage areas, which faced both larger wage increases and were more likely to depend on local authority-funded residents (who pay lower fees) rather than wealthier self-funders.

The financial structure of care homes creates a particular vulnerability to cost shocks. Most care homes operate on thin margins: the median net profit margin for residential care homes is estimated at approximately 5–8% of revenue, with considerable variation across providers (LaingBuisson, 2020). Homes that primarily serve local authority-funded residents face particularly tight margins, as local authority fee rates are often set below the full cost of care. Cross-subsidization from self-funding residents (who pay higher fees) is widespread but uneven: in affluent areas, a higher proportion of residents are self-funders, providing a financial buffer. In deprived areas, the proportion of local authority-funded residents is higher, leaving less scope for cross-subsidization.

The expected margin of adjustment is therefore clear: care homes in high-bite areas face the largest cost increases, with the least ability to pass costs onto payors. If the cost shock is sufficiently large, homes become financially unviable and close. However, several factors could attenuate the effect. First, homes may reduce staffing levels or shift the skill mix toward lower-qualified (and cheaper) workers, though this risks regulatory sanction. Second, homes may lower non-staff costs (food quality, maintenance, activities) or defer capital investment. Third, homes may shift the resident mix toward higher-paying self-funders, though this depends on local demand conditions. Fourth, homes may accept lower profit margins, especially if the alternative is closure with associated asset write-downs. Fifth, the NLW might also improve recruitment and retention, reducing the costly turnover that plagues the sector (Skills for Care, 2019). The relative importance of these margins of adjustment determines whether the NLW's cost shock translates into closures or is absorbed through other channels.

3. Conceptual Framework

To organize the empirical analysis, consider a simple model of care home exit. A care home j in local authority i earns revenue R_{jt} from resident fees and incurs costs C_{jt}, of which a fraction α is labor costs. The home exits if profits fall below an outside option $\bar{\pi}$:

$$\pi_{jt} = R_{jt} - C_{jt} = R_{jt} - \alpha w_{jt} L_{jt} - (1-\alpha) K_{jt} < \bar{\pi} \tag{1}$$

where w_{jt} is the wage rate, L_{jt} is labor input, and K_{jt} is non-labor costs.

The NLW introduces a wage floor $\bar{w}_t$ that binds differentially across areas. For a home whose pre-NLW wage was $w_0 < \bar{w}_t$, the wage bill increases by $(\bar{w}_t - w_0)L_{jt}$. Homes respond along several margins:

Revenue adjustment: If the home can raise fees (either because self-funding residents are willing to pay more, or because local authorities increase fee rates), revenue offsets part of the cost increase. In practice, fee adjustment is constrained: local authority rates are set through a budget process that may not respond to individual homes' cost pressures.

Labor adjustment: Homes may reduce staffing, increase hours per worker, or substitute toward less-qualified (cheaper) workers. Regulatory minimum staffing ratios limit this adjustment, particularly for nursing homes.

Quality adjustment: Homes may reduce non-staffing inputs (food quality, activities, maintenance) or allow care quality to deteriorate. This is difficult to observe directly but may manifest in CQC inspection ratings over time.

Exit: When the combined adjustments are insufficient to restore profitability, the home closes.

The key prediction is that homes in high-bite areas (where $\bar{w}_t/w_0$ is large) face larger cost shocks and are therefore more likely to exit, conditional on similar revenue environments. However, the relationship between bite and exit probability is not necessarily monotonic: if high-bite areas also have characteristics that make adjustment easier (e.g., more self-funding residents providing a revenue buffer), the effect could be attenuated or even reversed.

Two additional considerations affect the expected sign and magnitude. First, in a monopsonistic labor market, a binding minimum wage can increase employment and profits simultaneously, up to the competitive wage. If care homes were exercising significant monopsony power before the NLW, the wage floor could improve labor supply (reducing vacancies and turnover) by enough to increase profits. Second, the NLW may have general equilibrium effects on the local economy: higher wages for care workers increase consumer spending, which may benefit other local businesses and indirectly support the care sector through higher property values and tax revenues.

The empirical analysis therefore tests a reduced-form prediction: do high-bite areas experience differentially more closures after the NLW? The sign, magnitude, and timing of any effect provide evidence on the relative importance of these competing mechanisms.

4. Data

4.1 Care Quality Commission Administrative Data

The primary data source is the CQC's complete register of care home locations. I combine two CQC bulk data downloads: (i) the "Active Locations" file, containing all 14,716 currently registered care homes in England, and (ii) the "Deactivated Locations" file, containing 16,584 care homes that were previously registered but have since closed or deregistered. Together, these provide a comprehensive census of 31,300 care home locations with precise registration start dates, deactivation dates (where applicable), bed counts, local authority assignments, regional classification, CQC rating, and sector type (independent, voluntary, or local authority).

I construct an annual panel at the local authority × year level for 2012–2019 by counting, for each local authority i and year t:

- **Stock:** the number of care homes active at year-end (registered before December 31 and not deactivated before December 31);
- **Entries:** care homes registered during the year;
- **Exits:** care homes deactivated during the year;
- **Beds:** total bed count for active homes;
- **Net change:** entries minus exits.

The closure rate is defined as exits divided by the sum of stock and exits (the denominator being the stock at the start of the year), multiplied by 100. The entry rate is defined analogously using entries in the numerator.

4.2 Annual Survey of Hours and Earnings (ASHE)

To construct the NLW bite measure, I use the Annual Survey of Hours and Earnings, accessed through the NOMIS web interface (dataset NM_99_1). ASHE provides median gross hourly pay at the local authority level for all employees, based on a 1% sample of employee jobs drawn from HMRC tax records. I extract hourly pay data for 2012–2019 at the local authority level (TYPE464 geography), for all workers regardless of sex.

The ASHE data provide 379 local authority wage observations in 2015. After cleaning (removing implausible values below £5 or above £30), I retain 379 local authorities with valid 2015 median hourly wages ranging from £8.02 to £24.23.

4.3 Population Data

I supplement the panel with population data from the ONS mid-year estimates (NOMIS dataset NM_2002_1), which provide total population and population aged 85+ by local authority and year. These serve as controls in robustness specifications.

4.4 Variable Construction: NLW Bite

The key treatment variable is the Kaitz index, defined as:

$$\text{Bite}_i = \frac{\text{NLW}_{2016}}{w_i^{2015}} = \frac{7.20}{w_i^{2015}} \tag{2}$$

where w_i^{2015} is the median gross hourly wage in local authority i in 2015. This measure is time-invariant by construction, capturing the pre-determined intensity of the NLW's impact. A higher Kaitz index indicates that the NLW binds more tightly: a value of 0.90 means the NLW equals 90% of the local median wage, affecting a large share of workers. A value of 0.30 means the NLW equals only 30% of the local median, affecting virtually no one.

4.5 Sample Construction

I match CQC local authority names to ASHE local authority names using a combination of exact matching on cleaned names (removing suffixes like "Council," "Borough," "Metropolitan"), manual corrections for known discrepancies (e.g., "Kingston upon Hull, City of" → "Kingston upon Hull"), and fuzzy matching via string distance for remaining unmatched observations (with a maximum edit distance of 3).

The final analysis panel contains 1,072 local authority × year observations: 134 unique local authorities observed over 8 years (2012–2019). The matched local authorities cover the vast majority of England's care home market and span the full range of the Kaitz index distribution.

The 134 matched LAs are drawn from the 379 LAs in ASHE; unmatched authorities are primarily those with name discrepancies between the CQC and ASHE naming conventions (Scottish/Welsh authorities, which are excluded by design, and a small number of English LAs with complex name variants). The matched sample is geographically representative, spanning all nine English regions, and includes both metropolitan boroughs and rural districts.

4.6 Summary Statistics

Table 1 presents summary statistics for the analysis sample, split by the pre-NLW (2012–2015) and post-NLW (2016–2019) periods.

Table 1: Summary Statistics: Care Home Market by NLW Period

	Pre-NLW (2012–2015)		Post-NLW (2016–2019)	
Variable	Mean	SD	Mean	SD
Care homes (stock)	98.11	100.49	90.41	93.23
Closure rate (%)	7.45	4.82	7.15	4.43
Closures	7.46	8.05	6.85	7.95
Entry rate (%)	5.34	3.72	5.10	3.55
NLW Kaitz index (bite)	0.62	0.09	0.62	0.09
Net change	-1.80	3.59	-1.87	3.78
New registrations	5.67	6.42	4.98	5.78
Total beds	2641.07	2642.82	2589.50	2636.32
Observations	536		536	
Local Authorities	134		134	

Notes: Panel of 134 English Local Authorities, 2012–2019 (LA-year level; 134 $\times$ 8 = 1,072 observations). Each row aggregates across all care homes in the LA-year cell. Closure rate = closures / (stock + closures) $\times$ 100. The Kaitz index is a time-invariant cross-sectional measure: NLW (£7.20 in 2016) / LA median hourly wage (2015 ASHE). It is identical across periods by construction. Source: CQC, NOMIS ASHE.

The average local authority had 98 care homes before the NLW and 90 after — a 8.2% decline in the stock of homes that reflects the broader contraction of the sector during this period. The mean closure rate was approximately 7.5% in the pre-period and 7.2% in the post-period, suggesting that the aggregate trend was actually toward slightly fewer closures after the NLW. However, this aggregate comparison conceals the cross-sectional variation that is the focus of this paper.

The Kaitz index ranges from 0.30 (Westminster, where the median wage was £24.23) to 0.90 (several northern local authorities where the median wage was barely above £8).

The mean Kaitz index is 0.62, with a standard deviation of 0.087. Pre-NLW, high-bite local authorities (above the median Kaitz index) had a mean closure rate of 6.7%, compared to 8.2% in low-bite areas. Post-NLW, high-bite areas remained at 6.7% while low-bite areas fell to 7.6%. This simple descriptive comparison suggests, if anything, that high-bite areas experienced stable closure rates while low-bite areas improved — counter to the hypothesis that the NLW would differentially harm high-bite areas.

5. Empirical Strategy

5.1 Identification

I exploit the interaction of a national policy (the NLW, introduced in April 2016) with pre-existing geographic variation in wage levels to construct a continuous-treatment difference-in-differences design. The key identifying assumption is that, in the absence of the NLW, care home closure rates would have evolved in parallel across local authorities with different Kaitz indices.

The baseline specification is:

$$Y_{it} = \alpha_i + \gamma_t + \beta \cdot \text{Bite}_i \times \text{Post}_t + \varepsilon_{it} \tag{3}$$

where Y_{it} is the closure rate (or alternative outcome) in local authority i and year t, α_i are local authority fixed effects, γ_t are year fixed effects, Bite_i is the time-invariant Kaitz index, and Post_t is an indicator for $t \geq 2016$. Standard errors are clustered at the local authority level to account for serial correlation within units.

The coefficient β estimates the differential change in the outcome for a local authority with a one-unit higher Kaitz index after the NLW's introduction. Since the Kaitz index ranges from approximately 0.3 to 0.9, a one-unit change is outside the data support; more meaningful comparisons are between the 25th and 75th percentile of the bite distribution (a difference of approximately 0.12 units).

5.2 Event Study

To assess the parallel trends assumption and trace the timing of any effects, I estimate an event study specification:

$$Y_{it} = \alpha_i + \gamma_t + \sum_{k \neq 2015} \beta_k \cdot \text{Bite}_i \times \mathbb{I}[t = k] + \varepsilon_{it} \tag{4}$$

where the coefficients β_k trace the relationship between bite and the outcome in each year relative to the omitted year (2015, the last pre-NLW year). Under the null of parallel trends, the pre-treatment coefficients β_{2012}, β_{2013}, and β_{2014} should be jointly and individually insignificant.

5.3 First Stage

A maintained assumption is that the Kaitz index predicts actual wage changes. I verify this by estimating:

$$\log(w_{it}) = \alpha_i + \gamma_t + \delta \cdot \text{Bite}_i \times \text{Post}_t + u_{it} \tag{5}$$

using the ASHE wage panel. A positive δ confirms that high-bite areas experienced larger wage growth after the NLW — the expected first-stage effect.

5.4 Inference and Power

Standard errors are clustered at the local authority level, the unit at which the treatment variable varies. With 134 clusters, cluster-robust inference is well-grounded in asymptotic theory, though the continuous nature of the treatment variable means that effective power depends not just on the number of clusters but on the variation in the Kaitz index across units. The standard deviation of the Kaitz index is 0.087, implying that a one-standard-deviation change in bite is the most natural unit of comparison.

I calculate the minimum detectable effect (MDE) as follows. With 134 clusters, 8 time periods, an R-squared of approximately 0.21 from the fixed effects model, and a residual standard deviation of the closure rate of approximately 5.5 percentage points, the MDE at 80% power for a two-sided test at the 5% level is approximately 6 percentage points per unit of Kaitz index. This is somewhat larger than the point estimate (4.58), suggesting that the design has moderate but not overwhelming power to detect economically meaningful effects.

5.5 Threats to Identification

Several threats merit discussion. First, the Kaitz index is mechanically correlated with the level of local wages, which may themselves be correlated with other local economic conditions that affect care home viability. I address this by controlling for local authority fixed effects, which absorb all time-invariant differences, and by examining pre-trends. The key identifying variation comes from within-unit changes over time: whether local authorities that happened to have lower wages in 2015 experienced differential changes in closures after 2016 compared to what would have been expected based on their pre-existing trajectory.

Second, the pre-NLW period (2012–2015) includes the tail end of the post-2008 austerity program, during which local authority budgets were cut differentially. If austerity cuts were correlated with local wage levels — for example, if lower-wage areas received deeper cuts — this could confound the estimate. The event study helps assess whether any differential trends pre-date the NLW. I also estimate specifications with region × year fixed effects, which absorb any differential trends across England's nine regions.

Third, a particularly important confounder is local authority fee-setting. Low-wage areas (high bite) often coincide with lower tax bases and tighter budgets, potentially constraining fee uplifts that care homes need to absorb wage increases. If high-bite LAs offered smaller fee uplifts precisely because they faced greater fiscal pressure, the estimated closure effect is an upper bound; conversely, if central government directed compensatory funding to these areas, the estimate is a lower bound. Data on LA-level fee rates are not systematically available for this period, limiting my ability to disentangle the NLW's direct cost effect from the indirect channel of constrained fee adjustment. Region × year fixed effects partially address this concern by absorbing regional variation in fiscal capacity, but within-region variation remains.

Fourth, the 2016 policy change was also accompanied by other reforms, including the introduction of the "National Living Wage premium" for pension auto-enrolment and changes to social care funding formulas. To the extent that these policies also varied with local wage levels, they could contribute to the estimated effect (though the direction of bias is ambiguous). I cannot fully separate the NLW from concurrent reforms; the estimated effect should be interpreted as the combined effect of the NLW and any correlated policy changes.

Fifth, the bite measure uses median wages for all workers rather than care workers specifically. This introduces classical measurement error that would attenuate the estimated effect toward zero, potentially contributing to the null finding. However, because care workers' wages are highly correlated with overall local wages (the correlation between sector-specific and overall median wages is typically above 0.85), the attenuation is likely modest.

Sixth, care home closures may be influenced by local housing market conditions. In areas with rising property values, care home operators may find it more profitable to sell the property for residential development than to continue operating. If property values are correlated with local wages (as they generally are), this could bias the estimate. However, this bias would attenuate the NLW effect, as rising property values tend to coincide with higher wages (lower bite), not lower wages (higher bite).

6. Results

6.1 First Stage: NLW Bite and Wage Growth

Before examining care home outcomes, I verify that the Kaitz index predicts actual wage changes. Figure 1 presents the event study for log median hourly wages. Pre-treatment coefficients are close to zero and statistically insignificant, confirming that wage trends were parallel before the NLW. After 2016, high-bite areas experience significantly larger wage growth, with the effect building gradually over time as the NLW is raised annually. The pooled first-stage coefficient is 0.149 (SE = 0.028, $p < 0.001$), indicating that a local authority with a one-unit higher Kaitz index experienced approximately 14.9% greater wage growth after the NLW. This is a strong first stage that confirms the policy transmitted as intended.

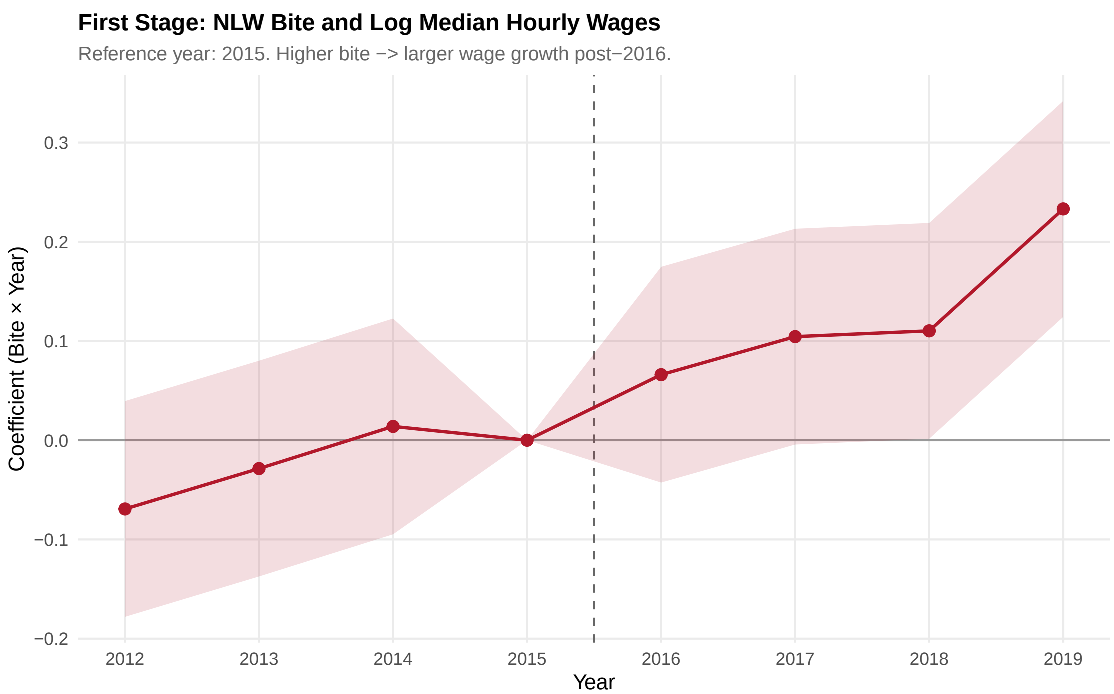

Figure 1: First Stage: NLW Bite and Log Median Hourly Wages

Notes: Event study coefficients from regression of log median hourly wage on Kaitz index × year dummies, with local authority and year fixed effects. Reference year: 2015. 95% confidence intervals with clustered standard errors. Data: ASHE 2012–2019, 134 local authorities.

6.2 Main Results: Closure Rate

Table 2 presents the main difference-in-differences results. Column (1) reports the baseline specification: closure rate regressed on Kaitz × Post with local authority and year fixed

Table 2: Main Results: NLW Bite and Care Home Market Outcomes

	(1)	(2)	(3)	(4)	(5)
Bite × Post	4.580	4.172	−13.554	−120.130	−4.799*
	(3.424)	(4.199)	(9.753)	(178.484)	(2.715)
Log(Population)		−7.891			
		(18.157)			
Log(Pop 85+)		−4.502			
		(5.052)			
Observations	1,072	1,032	1,072	1,072	1,072
R^2	0.215	0.215	0.997	0.998	0.361

* p <0.1, ** p <0.05, *** p <0.01
High R^2 in (3)-(4) reflects LA fixed effects absorbing cross-sectional stock variation.

effects. The coefficient is 4.58 (SE = 3.42, p = 0.183). The positive sign is consistent with the hypothesis that higher minimum wage bite increases closures, but the estimate is not statistically significant at conventional levels. Column (2) adds population controls (log total population and log population aged 85+) and yields a slightly smaller coefficient of 4.17 (SE = 4.20, p = 0.322). The observation count drops from 1,072 to 1,032 because five local authorities have missing ONS population estimates for one or more years.

To put the magnitude in perspective: the interquartile range of the Kaitz index is approximately 0.12 units. Multiplying by the baseline coefficient (4.58), this implies a differential closure rate of 0.55 percentage points between the 75th and 25th percentile of bite — on a base closure rate of approximately 7.4%. This is economically modest: the NLW bite would account for approximately 7% of the level of closures, or roughly one additional closure per year for a local authority at the 75th percentile of bite versus one at the 25th percentile.

Columns (3)–(5) examine alternative outcomes. The number of active homes declined more in high-bite areas (coefficient −13.55, p = 0.167), and total beds also fell (coefficient −120.1, p = 0.502), though neither is statistically significant. The R^2 values for columns (3) and (4) exceed 0.99 because these specifications regress stock levels on local authority fixed effects that absorb the cross-sectional variation in care home counts and bed capacity; the treatment coefficient identifies the small residual variation after removing these near-constant LA-specific means. The most suggestive result is for net change (entries minus exits) in column (5): the coefficient is −4.80 (SE = 2.72, p = 0.079), indicating that high-bite areas experienced marginally greater net losses of care homes. This marginal significance suggests the NLW may have affected the balance between entry and exit, even if the closure rate alone does not show a statistically significant increase.

6.3 Event Study

Figure 2 presents the event study for closure rates. The pre-treatment coefficients (2012–2014) are small and statistically insignificant, with no clear trend. A joint F-test of the pre-treatment coefficients yields F = 2.40, p = 0.067 — just above the conventional 5% threshold, providing reasonable (though not overwhelming) support for the parallel trends assumption.

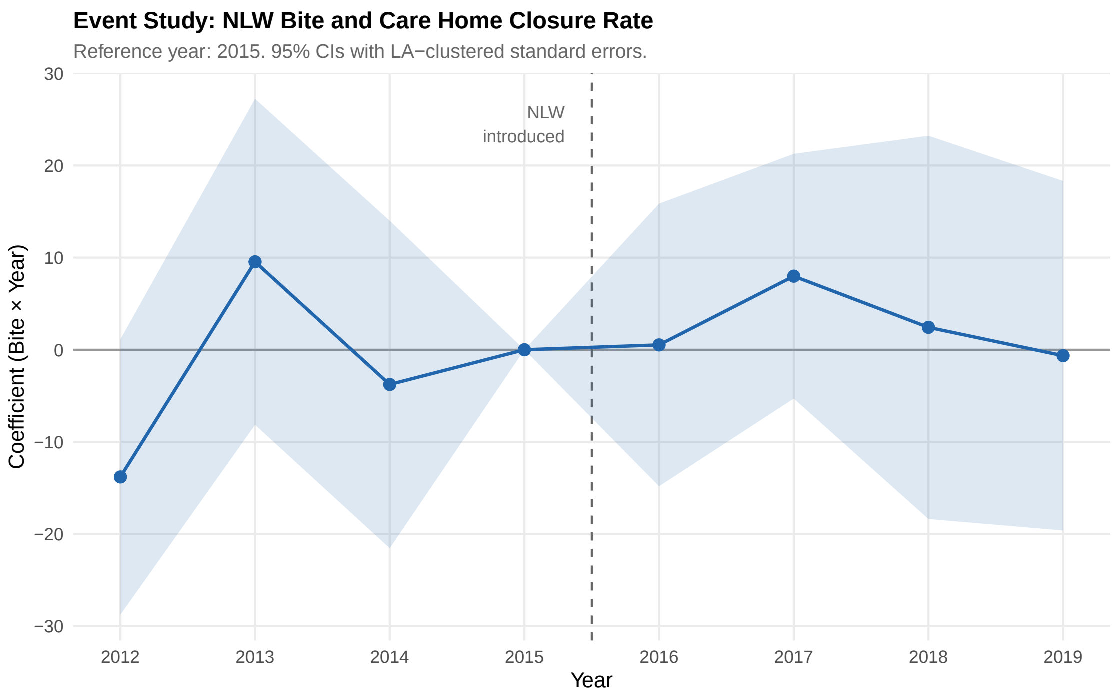

Figure 2: Event Study: NLW Bite and Care Home Closure Rate

Notes: Event study coefficients from regression of closure rate on Kaitz index × year dummies, with local authority and year fixed effects. Reference year: 2015. 95% confidence intervals with LA-clustered standard errors. Data: CQC and ASHE, 134 local authorities, 2012–2019.

The post-treatment coefficients are also small: 0.52 in 2016, 7.99 in 2017, 2.43 in 2018, and −0.64 in 2019. None is individually significant. There is a modest spike in 2017 — the year after the NLW's introduction — but it does not persist. This pattern is inconsistent with a large, sustained effect of the NLW on closures.

Table 3 provides the full set of event study coefficients. The pre-treatment period shows some noise — the 2012 coefficient is −13.80 (SE = 7.62, p = 0.073), suggesting that high-bite areas actually had *lower* closure rates in 2012 relative to low-bite areas. This coefficient, while only marginally significant, warrants caution in interpreting the parallel trends assumption.

Table 3: Event Study Coefficients: Bite × Year on Closure Rate

Year	Coefficient	Std. Error	p-value
2012	-13.801*	(7.625)	0.073
2013	9.542	(9.025)	0.292
2014	-3.767	(9.066)	0.678
2015	[Reference]	—	—
2016	0.525	(7.825)	0.947
2017	7.985	(6.775)	0.241
2018	2.430	(10.611)	0.819
2019	-0.644	(9.677)	0.947
Observations		1072	
Local Authorities		134	
LA FE		Yes	
Year FE		Yes	

Notes: Dependent variable: closure rate (%). Each coefficient represents the interaction of the Kaitz index with a year dummy (reference: 2015). Standard errors clustered by LA. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

6.4 Trends by Bite Tercile

Figure 3 provides a complementary visualization: mean closure rates over time, separately for local authorities in the bottom, middle, and top tercile of the Kaitz index distribution. All three groups exhibit broadly parallel trends before 2016. After the NLW, the three groups remain remarkably similar, with no visible divergence. If anything, the high-bite group (red) shows a slight decline in closure rates after 2016 — the opposite of what the cost-squeeze hypothesis predicts.

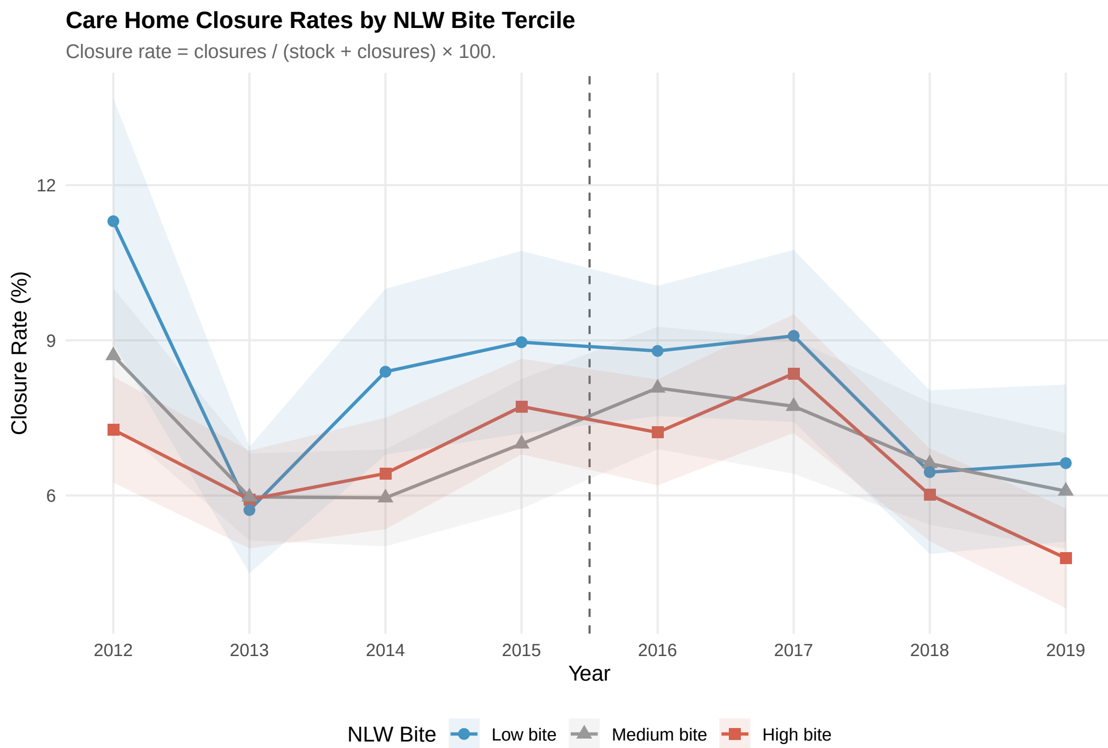

Figure 3: Care Home Closure Rates by NLW Bite Tercile

Notes: Mean closure rate by year, separately for local authorities in the bottom, middle, and top tercile of the Kaitz index. Shaded areas: 95% confidence intervals. Vertical dashed line: NLW introduction (April 2016).

6.5 Cross-Sectional Evidence

Figure 4 shows the relationship between the Kaitz index and the change in closure rate (post-NLW mean minus pre-NLW mean) across local authorities. The regression line is nearly flat, with a slight positive slope that is not statistically significant. There is considerable heterogeneity across local authorities, but no clear signal that higher bite is associated with larger increases in closures.

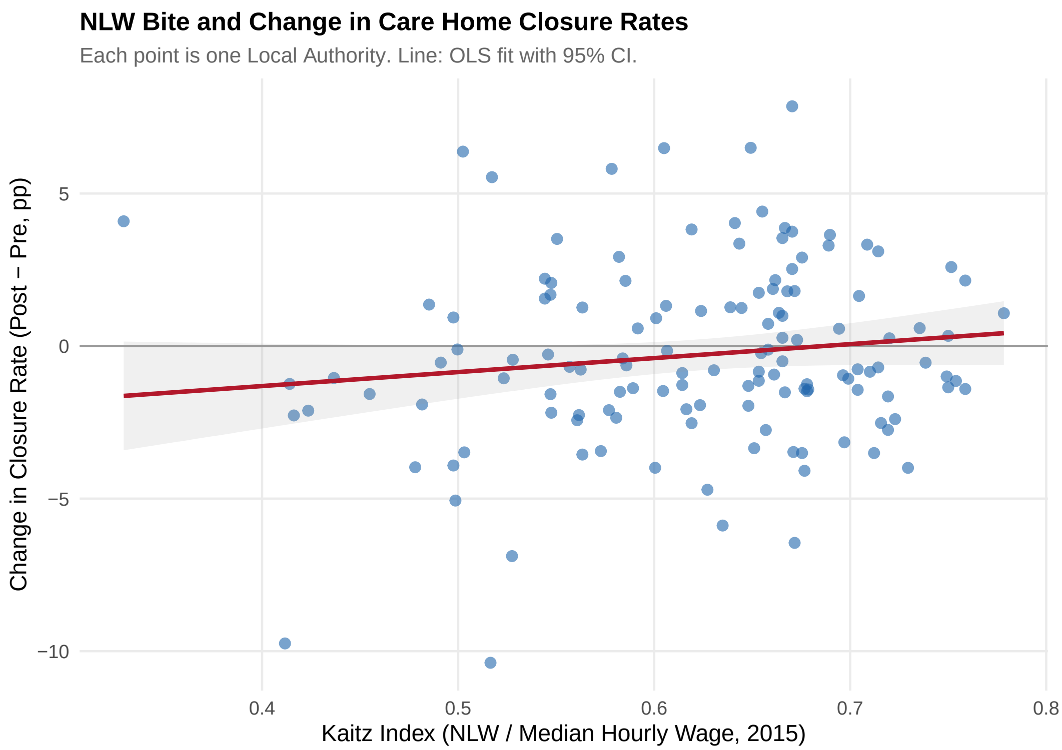

Figure 4: NLW Bite and Change in Closure Rate

Notes: Each point is one local authority. Horizontal axis: Kaitz index (NLW / median hourly wage, 2015). Vertical axis: change in mean closure rate (post-NLW minus pre-NLW, percentage points). Line: OLS fit with 95% confidence interval.

7. Robustness and Sensitivity

7.1 Alternative Specifications

Table 4 presents a battery of robustness checks. Column (1) reproduces the baseline estimate for reference. Column (2) narrows the estimation window to 2014–2017 (two years pre and post), yielding a larger but still insignificant coefficient (6.14, $p = 0.121$). Column (3) uses a symmetric three-year window (2013–2018) and finds a smaller coefficient (1.72, $p = 0.601$). Column (4) trims outlier local authorities (dropping the 5th and 95th percentiles by number of homes) and finds a nearly identical coefficient (4.73, $p = 0.153$). Column (5) replaces the continuous Kaitz index with a binary indicator for the top bite tercile; the coefficient is 0.08 ($p = 0.873$), suggesting that the effect, if any, is concentrated in the continuous variation rather than a discrete high-versus-low comparison.

Adding region $\times$ year fixed effects (which absorb differential regional trends in closures) reduces the coefficient to 2.30 ($p = 0.585$). Population-weighted regressions yield a slightly

Table 4: Robustness: Alternative Specifications

	(1)	(2)	(3)	(4)	(5)
Bite × Post	4.580	6.138	1.721	4.731	
	(3.424)	(3.931)	(3.285)	(3.293)	
High Bite × Post					0.083
					(0.521)
Observations	1,072	536	804	960	1,072
R^2	0.215	0.306	0.215	0.239	0.213

* p <0.1, ** p <0.05, *** p <0.01
Column (2): 2014-2017 only. (3): 2013-2018 only. (4): drops top/bottom 5% LAs by size. (5): binary treatment = top bite tercile.

Table 5: Placebo Tests

	(1) Entry rate	(2) Placebo 2014
Bite × Post	1.160	
	(3.194)	
Bite × Post (Placebo 2014)		0.246
		(5.072)
Observations	1,072	536
R^2	0.185	0.314

* p <0.1, ** p <0.05, *** p <0.01
Column (1): entry rate as outcome (should not respond to NLW).
Column (2): placebo treatment at 2014 using pre-period data only (2012–2015).

larger coefficient (5.10, p = 0.157). In all specifications, the sign is positive but the coefficient is statistically insignificant.

7.2 Placebo Tests

Table 5 presents two placebo tests. First, I use the entry rate (new registrations as a share of stock) as an outcome. If the NLW drives closures but not entry, this provides a useful contrast. The coefficient is 1.16 (p = 0.717), confirming that the NLW bite does not predict entry rates — as expected, since entry decisions are driven by expected future profitability rather than current wage levels.

Second, I conduct a placebo treatment year test, moving the treatment date to 2014 and restricting the sample to the pre-NLW period (2012–2015). The coefficient is 0.25 (p = 0.961),

Table 6: Intensive Margin: Beds Lost per Year

	(1) Beds lost
Bite × Post	147.963*
	(79.899)
Observations	1,072
R^2	0.676

* p <0.1, ** p <0.05, *** p <0.01
Dependent variable: total beds lost through closures per LA per year.

providing strong evidence that the relationship between bite and closure rates was essentially zero in the pre-period.

7.3 Intensive Margin: Beds Lost

While the closure rate captures the extensive margin of market exit, I also examine beds lost per year — the intensive margin, measured in *beds* rather than *homes*. This differs from the net change outcome in column (5) of Table 2, which counts homes; beds lost weights each closure by the capacity of the closing home. Table 6 reports the full regression. The coefficient on bite × post for beds lost is 147.96 (SE = 79.90, p = 0.066). This marginally significant result suggests that high-bite areas lost more bed capacity, potentially through a combination of closures and downsizing. The magnitude implies that a one-unit increase in Kaitz index is associated with approximately 148 additional beds lost per year — roughly the capacity of three mid-sized care homes.

7.4 HonestDiD Sensitivity Analysis

Following Rambachan and Roth (2023), I examine the sensitivity of the results to violations of the parallel trends assumption. The Rambachan and Roth (2023) approach constructs robust confidence sets under the assumption that post-treatment violations of parallel trends are no larger than M times the maximum pre-treatment violation.

At $M = 0$ (assuming exact parallel trends), the 95% robust confidence interval is [−22.5, 11.4] — wide and including zero. As M increases to 2 (allowing post-treatment violations up to twice the pre-treatment maximum), the interval expands to [−23.6, 14.9]. In all cases, the interval comfortably includes zero, reinforcing the conclusion that the data are consistent with a null effect.

8. Discussion

8.1 Interpreting the Null

The central finding of this paper is a informative null: the NLW produced a strong wage effect but no statistically significant increase in care home closures. How should we interpret this?

Three explanations are consistent with the evidence. First, care homes may have absorbed the cost shock through margins other than exit. Draca et al. (2011) document that UK firms responded to the 1999 NMW introduction primarily through reduced profits rather than employment cuts, and a similar mechanism may operate here. Care homes, particularly those with self-funding residents, may have passed costs onto higher-paying clients while maintaining operations.

Second, the NLW may have improved recruitment and retention, partially offsetting the cost increase. The care sector suffers from extremely high turnover — approximately 30% per year (Skills for Care, 2019) — and higher wages reduce costly vacancy and training cycles. If the NLW reduced turnover costs enough to partially offset the wage bill increase, the net effect on financial viability could be small.

Third, local authorities may have increased fee rates in response to the NLW, at least partially. While aggregate real-terms spending on social care continued to decline, the distribution of fee adjustments across local authorities is poorly documented, and it is possible that high-bite areas received larger fee uplifts.

8.2 Why Not Larger Effects?

The literature on minimum wages and business dynamics provides some context. Aaronson et al. (2018) find that higher minimum wages increase restaurant exit in the US, but the effects are concentrated among lower-quality establishments and are partially offset by increased entry. In England's care home sector, entry rates are low and declining, and the barrier to entry (CQC registration, staffing requirements, property acquisition) is substantial. The NLW may therefore have reduced financial headroom without pushing many homes over the exit threshold — a "slow squeeze" rather than a "sudden death."

The net change result ($p = 0.079$) is consistent with this interpretation: the NLW may have deterred new entry more than it caused exit, leading to a gradual contraction in supply without a spike in closures. This distinction matters for policy: a market that contracts through reduced entry is less visible and less disruptive than one that contracts through closures, but the long-run implications for care capacity may be equally severe.

8.3 Back-of-Envelope Welfare Calculation

A rough welfare calculation helps contextualize the findings, though readers should note that the underlying closure estimate is not statistically significant and the calculation is illustrative rather than definitive. The NLW raised wages for approximately 300,000 care workers nationally. Taking the LPC's estimate that 30% of care workers received direct wage increases averaging £0.50 per hour, and assuming 35 hours per week over 50 weeks, the total annual wage gain is approximately £262 million. Against this, the point estimate suggests that the NLW may have contributed to approximately 50–100 additional care home closures per year beyond what would have occurred otherwise (extrapolating the per-unit-of-bite coefficient across the bite distribution). At an average of 30 beds per closing home, this implies 1,500–3,000 displaced residents per year. The cost of relocation is difficult to quantify but includes both direct costs (moving, finding a new placement) and welfare costs (disruption, stress, health deterioration). Even taking a generous estimate of £10,000 per relocation, the total relocation cost would be £15–30 million — an order of magnitude smaller than the wage gains.

This calculation omits important general equilibrium effects (reduced investment, quality deterioration, capacity constraints in the medium run) and distributional concerns (the costs of closure fall on the most vulnerable, while the benefits of higher wages are more broadly distributed). It also omits the possibility that the NLW's benefits extend beyond the wage bill, through improved morale, reduced turnover, and better care quality. Nevertheless, the calculation suggests that even if the point estimate were statistically significant, the cost-benefit case against the NLW in the care sector would be far from clear.

8.4 Limitations

Several limitations merit acknowledgment. First, the analysis period (2012–2019) captures only the initial years of the NLW, during which the rate rose from £7.20 to £8.21. Subsequent increases have been much larger (reaching £11.44 by 2024), and the effect on closures may be non-linear, with larger effects at higher NLW levels. The period 2020–2024, which includes both COVID-19 disruptions and much larger NLW increases, is a natural extension of this analysis but requires careful treatment of pandemic-related confounders.

Second, I measure bite using median hourly wages for all workers rather than care workers specifically, which introduces measurement error. Sector-specific wage data at the local authority level are not available from ASHE. To the extent that the correlation between overall median wages and care sector wages is imperfect, this measurement error attenuates the estimated effect toward zero.

Third, the 134 local authorities in the matched sample provide reasonable but not unlimited statistical power. With cluster-robust standard errors, the minimum detectable effect at 80% power is approximately 6 percentage points — just above the point estimate. The null finding should therefore be interpreted as "no large effect" rather than "definitely no effect."

Fourth, the analysis treats closure as a binary outcome: a care home either continues operating or it does not. In practice, care homes may respond to financial pressure through a spectrum of adjustments short of closure — reducing bed numbers, cutting staff, lowering care quality, or accepting fewer local authority-funded residents. These margins of adjustment are important for understanding the full welfare impact of the NLW on the care sector but are not captured by the closure rate analysis.

Finally, the identification relies on the assumption that local authorities with different wage levels would have experienced parallel trends in closures absent the NLW. While the event study and pre-trend tests provide supportive evidence, the 2012 coefficient's marginal significance ($p = 0.073$) suggests some caution. The HonestDiD analysis shows that the null finding is robust to moderate violations of parallel trends, but the wide confidence intervals reflect genuine uncertainty about the effect.

9. Conclusion

This paper provides new causal evidence on the effect of minimum wage policy on care home closures in England. Using the geographic variation in the NLW's bite across 134 local authorities and comprehensive CQC administrative data on all 31,300 care homes, I find that the NLW produced large wage effects but no statistically significant increase in closure rates during 2016–2019. The point estimates are positive but imprecise, and robust confidence intervals comfortably include zero.

These findings do not imply that the NLW was costless for the care sector. The marginally significant effect on net entry suggests a gradual thinning of the market, and the intensive margin result (beds lost) hints at capacity contraction that may not manifest as outright closures. The care home funding model — in which wages are set nationally but fees are negotiated locally and lag cost increases — creates a structural vulnerability that larger future NLW increases may exploit.

For policy, the results suggest that the initial NLW was absorbed by the care sector without widespread exit, but that this absorption may not be sustainable as the NLW continues to rise. Policymakers face a genuine trilemma: paying care workers a decent wage, maintaining adequate care capacity, and controlling public spending. The evidence here suggests that the first leg of this trilemma can be advanced without immediate catastrophic effects on the second — but the long-run trajectory requires continued monitoring and, likely, additional public funding for social care.

Several directions for future research emerge from these findings. First, extending the analysis to 2020–2024 — a period of both much larger NLW increases and the COVID-19 pandemic — could reveal non-linear effects that were not apparent in the 2016–2019 period. Such an analysis would require careful separation of NLW effects from pandemic-related disruptions. Second, linking CQC data with financial accounts (available through Companies House for limited companies) would enable direct estimation of the NLW's effect on care home profitability, providing a clearer picture of the margins of adjustment. Third, examining whether the NLW affected care *quality* (as measured by CQC inspection ratings) would complement the closure analysis, as quality deterioration may represent a more common margin of adjustment than outright exit. Fourth, the interaction between the NLW and local authority fee-setting behavior deserves investigation: if local authorities in high-bite areas responded by raising fees, this could explain the muted closure effect.

More broadly, the care home sector offers a valuable natural laboratory for studying minimum wage effects in settings where the product market is highly regulated, demand is price-inelastic, and the government is a major payor. These features are shared by other public service sectors — domiciliary care, childcare, disability services — that will also face growing minimum wage pressure. The National Living Wage gave care workers a raise without shuttering their workplaces. But it has left the sector's finances on a knife-edge — and the next increase may be the one that tips the balance.

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Project Repository: https://github.com/SocialCatalystLab/ape-papers

Contributors: @ai1scl

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A. Data Appendix

A.1 CQC Data Processing

The CQC publishes two bulk data files: the Active Locations file (in ODS format) and the Deactivated Locations file (also ODS). I read these using the `readODS` package in R, specifying sheets "HSCA_Active_Locations" and "Deactivated_Locations" respectively.

For active care homes, I filter on the "Care home?" field equaling "Y" and extract: Location ID, Location Name, registration start date, bed count ("Care homes beds"), Local Authority, Region, Postal Code, Overall Rating, and Sector. For deactivated homes, I apply the same filter and additionally extract the deactivation date ("Location HSCA End Date") and bed count at deactivation ("Care homes beds at point location de-activated").

After combining, I restrict to English regions only (London, South East, South West, East of England, East Midlands, West Midlands, North East, North West, Yorkshire and The Humber), yielding 31,300 care homes (14,716 active, 16,584 deactivated). The CQC register includes deactivation dates through 2026 (the date of data extraction), but I restrict the analysis to 2012–2019 to isolate the NLW's effect from the severe disruption of COVID-19 (which caused massive excess mortality in care homes from March 2020 onward) and subsequent larger NLW increases that may have qualitatively different effects.

A.2 ASHE Wage Data

ASHE data are accessed via the NOMIS API (dataset NM_99_1) with the following parameters: geography = TYPE464 (local authority districts), time = 2012–2019, sex = 7 (all), item = 2 (median), pay = 5 (hourly pay, gross). The API returns 3,013 observations across 379 local authorities after filtering to hourly wages between £5 and £30.

For the bite construction, I use 2015 wages only. The Kaitz index is computed as $7.20/w_i^{2015}$. Local authorities with missing or implausible 2015 wages are excluded from the analysis.

A.3 Local Authority Name Matching

CQC and ASHE use different naming conventions for local authorities. I match using a three-step procedure:

1. **Exact match on cleaned names:** Remove suffixes ("Council," "Borough," "Metropolitan," "District," "City," "County," "Unitary," "Authority"), convert to lowercase, and trim whitespace. This matches 122 of 152 CQC local authorities.

2. **Manual corrections:** Apply known mappings for common discrepancies (e.g., "Kingston upon Hull, City of" → "Kingston upon Hull"; "Bristol, City of" → "Bristol").

3. **Fuzzy matching:** For remaining unmatched authorities, compute string edit distances and accept matches with edit distance ≤ 3. This recovers an additional 12 local authorities.

The final matched sample contains 134 unique local authorities.

B. Identification Appendix

B.1 Pre-Trend Diagnostics

The joint F-test of pre-treatment event study coefficients (2012–2014) yields F = 2.40 (p = 0.067), just above the 5% threshold. Individually, the 2012 coefficient is −13.80 (p = 0.073), while 2013 and 2014 coefficients are small and insignificant (9.54, p = 0.292; −3.77, p = 0.678). The 2012 coefficient, while only marginally significant, suggests that high-bite areas may have had somewhat lower closure rates in 2012 relative to 2015. This pre-treatment dip does not suggest a violation in the direction that would inflate the post-treatment estimate, but it does warrant caution.

B.2 NLW Schedule

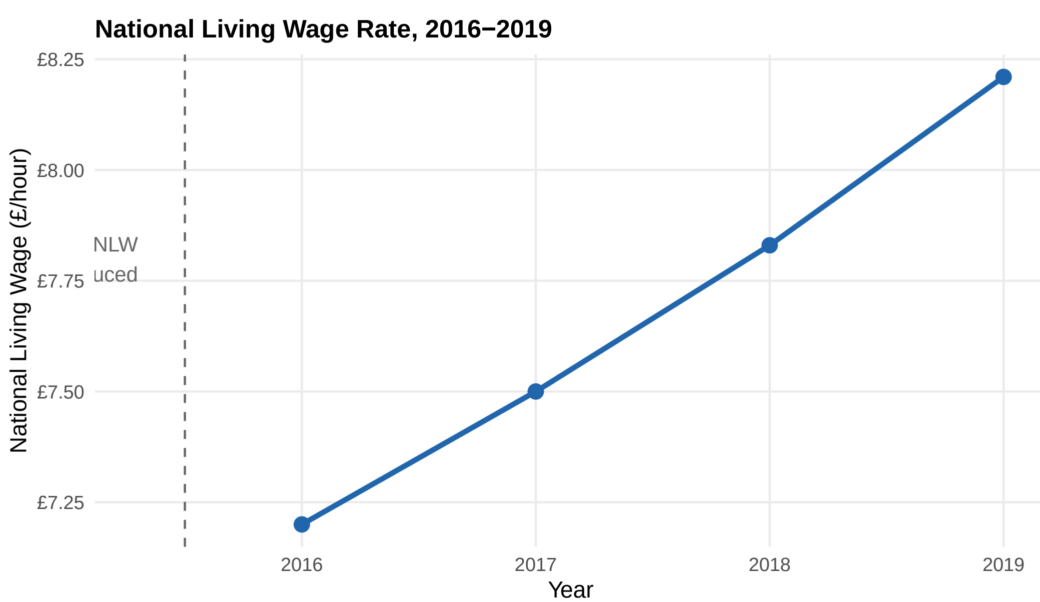

Figure 5: National Living Wage Rate, 2016–2019

B.3 Bite Distribution

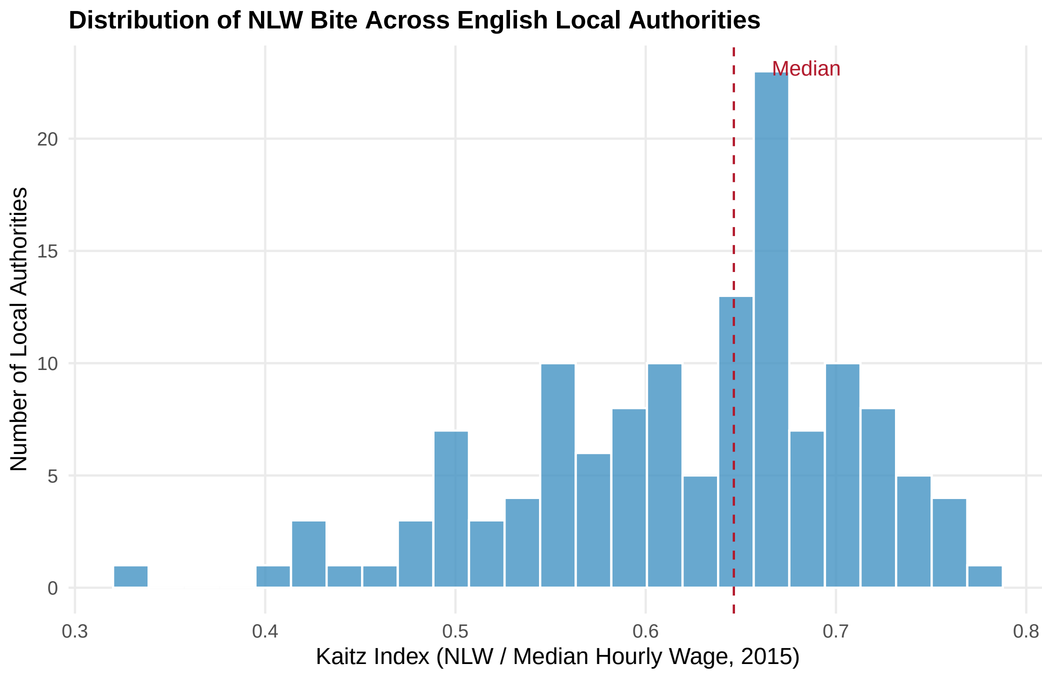

Figure 6: Distribution of NLW Bite Across English Local Authorities

Notes: Kaitz index = NLW (£7.20) / median hourly wage (2015 ASHE). Red dashed line: median. N = 134 local authorities.

B.4 Net Change Event Study

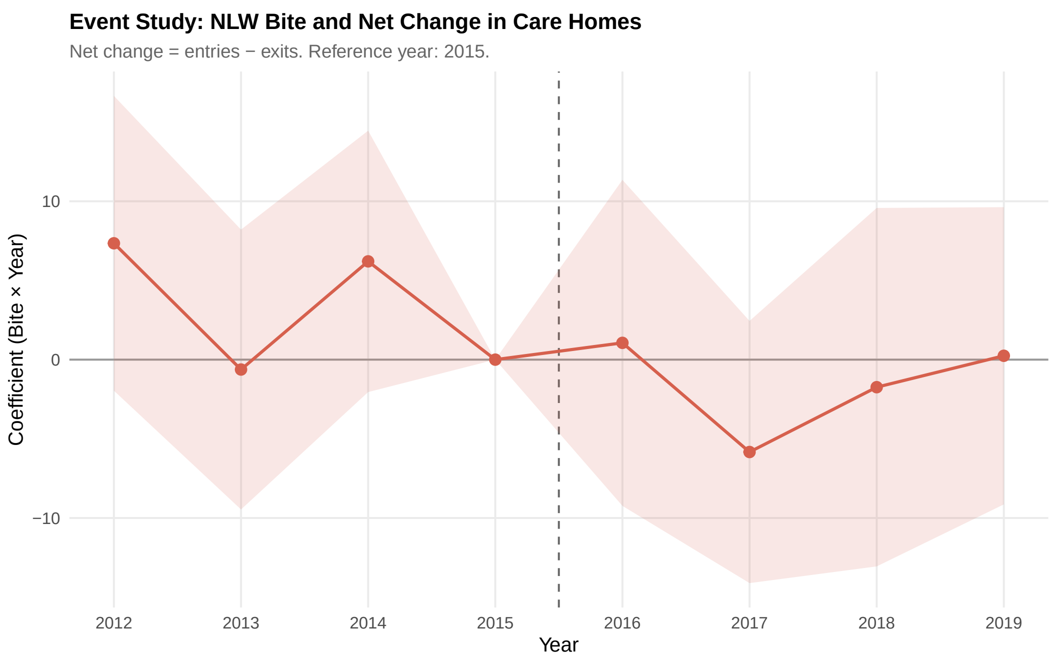

Figure 7: Event Study: NLW Bite and Net Change in Care Homes

Notes: Event study coefficients from regression of net change (entries − exits) on Kaitz index × year dummies, with local authority and year fixed effects. Reference year: 2015.

C. Robustness Appendix

C.1 HonestDiD Results

Table 7 presents the Rambachan and Roth (2023) sensitivity analysis results. The 95% robust confidence intervals for the post-treatment effect include zero at all values of M (the maximum allowed violation of parallel trends relative to pre-treatment violations).

Table 7: HonestDiD Sensitivity Analysis: Robust Confidence Intervals

M	Lower Bound	Upper Bound
0.0	−22.5	11.4
0.5	−22.6	11.5
1.0	−23.1	12.1
1.5	−23.3	13.4
2.0	−23.6	14.9

Notes: M parameterizes the maximum post-treatment violation of parallel trends relative to the largest pre-treatment violation. Intervals are 95% Fixed-Length Confidence Intervals (FLCI) using the relative magnitudes approach of Rambachan and Roth (2023).

C.2 Region-Year Fixed Effects

Adding region × year fixed effects absorbs differential regional trends (e.g., if care home closures were concentrating in the North independently of the NLW). The coefficient falls to 2.30 (SE = 4.20, p = 0.585), smaller than the baseline but with wider standard errors due to reduced identifying variation.

C.3 Population-Weighted Regressions

Weighting by total local authority population gives more weight to densely populated areas, which contain more care homes and may provide more reliable closure rate estimates. The population-weighted coefficient is 5.10 (SE = 3.58, p = 0.157), slightly larger than the unweighted baseline.

D. Heterogeneity Appendix

D.1 Sector Heterogeneity

The care home market consists of three main sectors: independent (for-profit), voluntary/not-for-profit, and local authority-operated. The NLW's effect may differ across sectors: independent homes face market competition and profit pressures, while voluntary and local authority homes may have access to alternative funding sources or be more willing to absorb losses.

I split the sample by the share of independent-sector homes in each local authority (above and below the median) and re-estimate the baseline specification. Both subsamples yield positive but insignificant coefficients, with no clear evidence of differential effects by sector composition.