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Examination activity and delayed reporting in England’s diagnostic imaging providers: an eight-quarter computed tomography and magnetic resonance imaging panel

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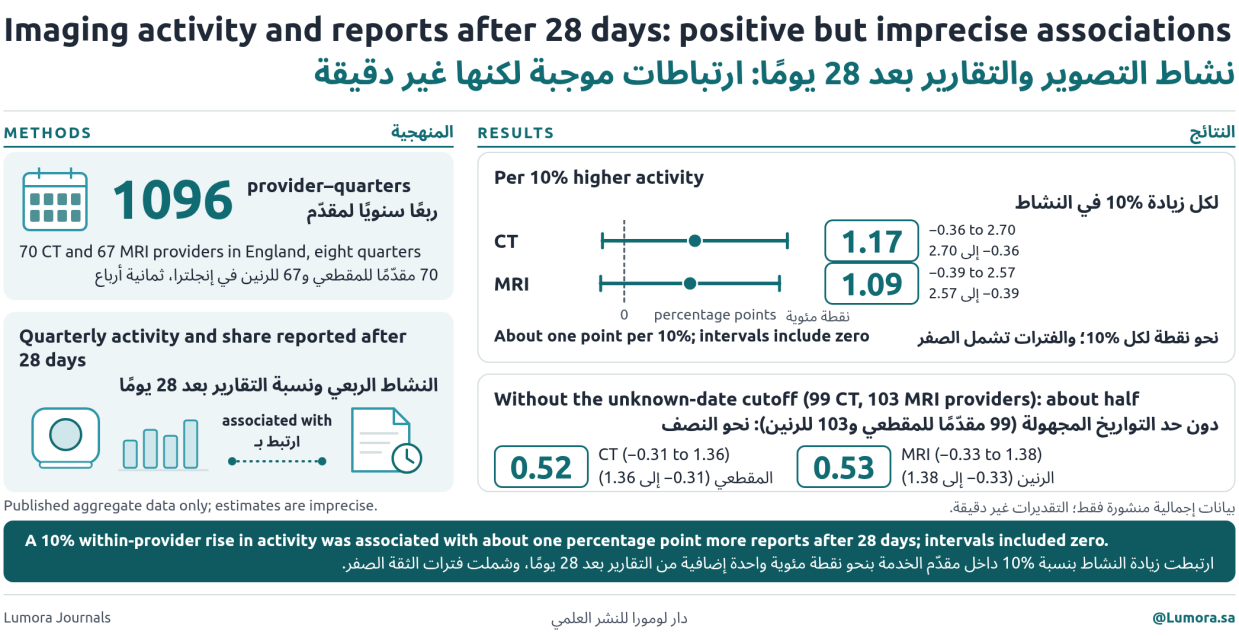
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Graphical abstract



Graphical abstract: A 10% within-provider rise in activity was associated with about one percentage point more reports after 28 days; intervals included zero. Designed by the Lumora Journals production team using hand-coded HTML/SVG rendered in Google Chrome.

Abstract

Introduction: Examination counts and report-date completeness are distinct organisational measures. We examined whether within-provider changes in computed tomography (CT) and magnetic resonance imaging (MRI) activity accompanied changes in the share of examinations reported after 28 days in England, separately from differences between providers.

Methods: We analysed eight quarters (April 2023–March 2025) of the finalised public Diagnostic Imaging Dataset (DID) Table 13 workbooks. The panel was restricted to complete R-coded providers. Linear models with provider and quarter fixed effects and acute-setting share were fitted separately for CT and MRI, with provider-cluster inference. Unknown-date bounds, rounding, cohort definitions, influential provider-quarters, denominator weighting and a fractional logit model were examined as sensitivity analyses.

Results: The primary panels comprised 70 CT and 67 MRI providers (1096 provider–quarters). A 10% within-provider increase in activity was associated with 1.17 (−0.36 to 2.70) percentage points more examinations reported after 28 days for CT and 1.09 (−0.39 to 2.57) for MRI (8.49 and 7.92 per doubling). Within-provider activity varied little (standard deviation 6.2% and 9.8%). Without the unknown-date cutoff, the cohorts expanded to 99 and 103 providers and the estimates roughly halved (3.79 and 3.83 per doubling). Unknown-date bounds were 1.41–14.36 and 2.99–12.79; these are not confidence intervals.

Conclusion: Within-provider estimates were positive but imprecise, compatible with zero, and sensitive to the report-date completeness rule and to a small number of provider-quarters. The published aggregates do not establish causal congestion, urgency-specific compliance, national performance or patient benefit.

Keywords: Diagnostic Imaging; Tomography, X-Ray Computed; Magnetic Resonance Imaging; Health Services Research; Time Factors; State Medicine

Introduction

Timely reporting is an important part of diagnostic imaging services, but acquiring an examination, issuing its report and acting on it are different processes. More completed examinations can improve access to scanning without shortening the time until reports are available, and a provider can absorb more examinations through additional reporting capacity or changed workflow. England's Diagnostic Imaging Dataset (DID) publishes examination activity and reporting intervals for each provider, which allows these quantities to be examined together at organisational level [1–6]. The published aggregates do not identify who reported the examinations or the consequences for patients.

National guidance introduced in August 2023 set a maximum reporting timeframe of four weeks, with shorter expectations for particular pathways [7]. The Care Quality Commission had found variation in radiology reporting arrangements [8], and workforce and recovery reports describe a persistent gap between imaging demand and reporting capacity [9,10]. Imaging use rose steadily in high-income health systems over two decades [11]. Services have responded with outsourcing, teleradiology and other delivery changes: a systematic review of 57 studies found promising but mainly uncontrolled evidence that such initiatives shorten radiology waiting times [12]; a UK audit found 11% of departments in cross-organisation out-of-hours reporting rotas [13]; remote reporting increased reporting output in one NHS department [14]; and outsourced CT reports at a UK centre recommended inappropriate follow-up more often than internal reports [15]. Published activity is therefore not a measure of workload per radiologist, because the same count can be reported by different staff, in-house or outsourced.

Studies of radiology workflow show why an aggregate queue cannot be interpreted without its operational context. Lagzi and colleagues examined interactions between priority classes on a radiology workflow platform [16]; Becker and colleagues and Law and colleagues evaluated subspecialty workflow management and automatic examination assignment at a tertiary cancer centre [17,18]; and Lockwood and colleagues evaluated facilitation of radiographer reporting of musculoskeletal radiographs across London [19]. These studies use detailed operational records that the public quarterly workbooks do not contain.

The official 2024/25 DID release already reports that longer median test-to-report periods did not occur preferentially in providers with higher or lower activity, for the chest radiograph, brain MRI and abdominal and pelvic ultrasound tests that could contribute to early cancer diagnosis [3]. That is a comparison between providers at one time. It cannot show whether a provider's own changes in activity coincide with changes in delayed reporting, because larger and smaller providers differ persistently in case mix, configuration and record quality. Panel methods separate the within-provider relationship from the between-provider one [20,21]. Our contribution is to estimate the within-provider association for CT and MRI, with the between-provider contrast as a secondary comparison against the official statement.

The report-date field creates a measurement problem. An examination with an unknown report date is not equivalent to one reported promptly, and excluding unknown dates changes the denominator. Published aggregates allow this to be examined without assuming that dates are missing at random, by bounding the regression slopes that the unknown dates permit [22].

We asked whether within-provider changes in CT and MRI activity accompanied changes in the published share of the outpatient/general-practitioner group reported after 28 days, after accounting for provider and quarter effects and the acute-setting share, and how unknown report dates, rounding and cohort

restrictions affected the answer. The study offers a bounded organisational association with explicit measurement sensitivity, not a causal estimate or a validated performance indicator.

Methods

Design, data and reporting

This retrospective ecological panel study used public aggregate DID statistics for England's financial years 2023/24 and 2024/25 [1–6]. The unit was the provider–modality quarter, and CT and MRI were analysed separately. The target was the panel of R-coded providers with complete eligible records in all eight quarters. Reporting follows the STROBE and RECORD statements [23,24]; the completed checklist is provided as a supplementary file.

The analysis plan was not registered or deposited before analysis; all analyses are reported, and additional analyses are identified as such. Provider identity was clarified after the workbooks showed independent-sector codes repeated across regions.

Eight independent-sector codes occurred as regional segments in the same quarter. Their fractions could not be combined without an unavailable late-event numerator, and no R-coded provider had a duplicated code–modality–quarter. All panels were therefore restricted to R-coded providers. Official coding guidance describes the legacy R form and later code formats [25]; the restriction is a selection boundary, not an assertion that all NHS providers carry an R code.

Data and variables

We used Table 13 (report turnaround bands by modality and broad patient-source setting) for each year: the finalised 2023/24 workbook published in July 2025 and the 2024/25 workbook published in November 2025 [2,4]. Only the worksheets for 2023–24 Q1–Q4 and 2024–25 Q1–Q4 were extracted; legacy worksheets for earlier periods were excluded. Provisional 2025/26 data were outside the study.

For each provider, modality and quarter we extracted the acute-setting total (accident and emergency, inpatients and day cases), the outpatient/general-practitioner (OP/GP) group total, and the OP/GP fractions reported after 28 days and with report date not known. The OP/GP group also includes Other Health Care Provider and Other sources, and the whole-day intervals do not identify urgent referrals. The outcome is therefore the published broad-group share exceeding 28 days, not adherence to every pathway-specific target [5–7].

Total activity was the sum of the two setting totals. Its base-2 logarithm was the exposure, so coefficients are percentage points per doubling of activity; because within-provider activity varied much less than a doubling, we also report the association per 10% higher activity (coefficient × $\log_2 1.1$ = coefficient × 0.1375) and per within-provider standard deviation. The acute-setting share (acute total divided by combined total) represents a coarse setting mix, not clinical severity or urgency. The outcome was the published late fraction × 100.

Fractions are published to 0.001 and totals are rounded. Suppressed, missing and non-numeric cells were treated as unavailable and never set to zero.

Fixed eligibility and source-quality exclusions

The primary cohort required a valid row in all eight quarters, numeric non-negative totals, at least 100 OP/GP and 1,000 combined examinations in every quarter, valid late and unknown fractions (sum ≤1.001), and an unknown-date fraction ≤10% in every quarter. These are analytical support criteria, not clinical standards.

Whole providers were excluded if the publisher's provider-issue notes for either year recorded activity shortfalls or duplication, report-date problems, broad-setting miscoding or an acquisition; 32 provider codes were named, 31 of them R-coded. A GP-to-outpatient recoding within the same broad group, or an ultrasound-specific issue, did not lead to exclusion. Exclusions were applied in the sequence shown in Table 1 and Figure 1.

Statistical model and uncertainty

The primary linear model related the late percentage to $\log_2$(total activity) and the acute-setting share, with provider and quarter fixed effects and equal weight for each provider–quarter. The coefficient is therefore an unweighted conditional organisational association. Provider effects absorb time-invariant differences between organisations, and quarter effects absorb changes shared by all providers, including the national seasonal pattern; time-varying staffing, outsourcing and case mix are not controlled [20,21]. A provider-cluster sandwich covariance with the CR1 correction $G/(G-1) \times (N-1)/(N-K)$ was used, where G is the number of providers, N the number of provider–quarters and K the model rank, with 95% intervals from a t distribution with G−1 degrees of freedom [26].

A provider-resampling percentile bootstrap (2,000 resamples of whole eight-quarter provider sequences) was planned because it keeps each provider's series intact and gives intervals for the between-provider coefficient and the between-minus-within difference. Because the cited bootstrap guidance concerns the wild cluster bootstrap [27], we added a wild cluster restricted bootstrap-t for the within coefficient (Rademacher weights, 9,999 replicates, seed 20261101; interval by test inversion on a 0.05-point grid). The seeds of the original provider-bootstrap run were not recorded; Supplementary Code S1 reruns the provider bootstrap with documented seeds (596 for CT, 597 for MRI), giving −0.98 to 20.16 for CT and −1.86 to 18.35 for MRI. Both modality analyses were exploratory, without a confirmatory significance threshold.

The secondary between-provider model related each provider's mean late percentage to mean $\log_2$(activity) and mean acute share. The full within-model coefficients, including provider and quarter effects, are provided in Supplementary Code S1.

Additional analyses. We described the within-provider exposure distribution and identified influential provider-quarters by DFBETA, refitting after removing the most influential observation, all observations with $|\text{DFBETAS}| > 2/\sqrt{N}$, the five most extreme residualised exposures, and the provider with the largest activity range. Because the outcome is a bounded fraction near zero and each provider-quarter has equal weight, we fitted the primary model weighted by the OP/GP denominator and a fractional logit model with the same fixed effects and provider-cluster standard errors, reporting the average marginal effect of a 10% increase [28]. We also allowed provider-specific seasonal effects (provider × quarter-of-year), and, in the cohort without the unknown-date cutoff, regressed the unknown-date share and the combined late-or-unknown share on activity to examine the direction of selection.

Unknown dates, rounding and specification sensitivity

A known-date outcome divided the late fraction by one minus the unknown fraction, using the same cohort and regressors. It estimates the late share among examinations with a known report date; no missingness mechanism was assumed.

For unknown-date bounds, each quarter's true late percentage was allowed to lie between the observed value and the observed value plus the unknown percentage (capped at 100%). Because the fixed-design least-squares coefficient is a signed linear combination of the outcomes, assigning the unknown mass according to the sign of each weight gives the attainable minimum and maximum slopes. These are identification bounds, not confidence intervals [22].

A display-precision sensitivity moved each observed late percentage by at most 0.05 percentage points, half the displayed unit, through the same signed operator.

Planned robustness analyses relaxed the unknown-date limit to 20% and then removed it, each changing the cohort, and, within the primary cohort, omitted acute share, used first differences with quarter-change effects, and left out each provider in turn. The condition number of the within-transformed regressors is reported for the 2 × 2 cross-product matrix, which is the square of the condition number of the design matrix.

Verification and code

Workbook cells were extracted independently of the main spreadsheet reader, and coefficients, cluster covariances and unknown-date bounds were checked with a separate full-matrix implementation and by refitting the model with endpoint outcomes. Supplementary Code S1 contains the analysis code, which downloads the two workbooks, reproduces the cohort flow and every reported estimate, and writes the full coefficient set.

Results

Source flow and selected panels

The eight quarters contained 1,352 CT and 1,442 MRI provider rows from 148 and 151 provider codes, of which 137 and 138 were R-coded. After the fixed source-quality and acquisition exclusions, 107 providers remained for each modality, all with eight quarters. Numeric-cell requirements retained 106 CT and 104 MRI providers, volume and fraction requirements 99 and 103, and the unknown-date criterion 70 and 67 (560 and 536 provider–quarters; Table 1, Figure 1).

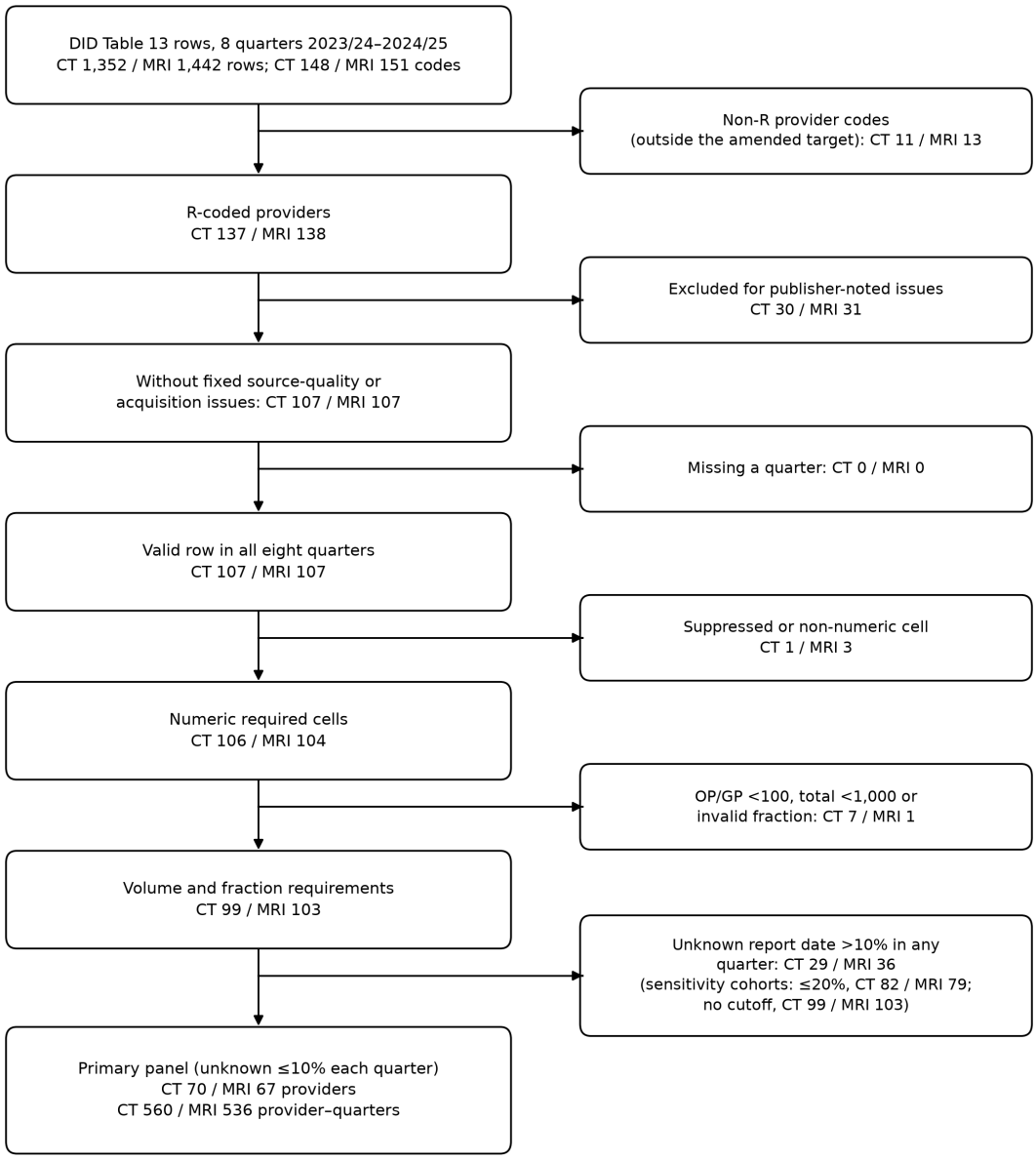
Table 1. Sequential source flow into the selected eight-quarter panel

Flow step	CT	MRI
Original source rows	1352	1442
All published provider codes	148	151
R-coded providers	137	138
After fixed source-quality/acquisition exclusions	107	107

Flow step	CT	MRI
After complete eight-quarter requirement	107	107
After numeric required-cell requirement	106	104
After volume/fraction requirements	99	103
Primary: unknown ≤10% each quarter	70	67
Primary provider–quarter observations	560	536

Notes for all tables and figures. CT, computed tomography; MRI, magnetic resonance imaging; DID, Diagnostic Imaging Dataset; OP/GP, outpatient/general-practitioner group (including Other Health Care Provider and Other); acute setting, accident and emergency, inpatients and day cases; CI, confidence interval; SE, standard error; $\log_2$, base-2 logarithm; pp, percentage points. n counts providers or provider–quarters, never patients; examination counts are rounded published totals. The late share uses the whole OP/GP denominator, including unknown report dates. Volume coefficients are percentage points per doubling of published activity unless labelled per 10%; acute-share coefficients refer to a change of the fraction from 0 to 1. Within models include provider and quarter fixed effects and acute share, with equal provider–quarter weight; cluster intervals use CR1 standard errors and t with providers − 1 degrees of freedom; provider-bootstrap intervals use 2,000 whole-provider resamples. Unknown-date bounds are fixed-design identification ranges, not confidence intervals. Rows marked † are additional analyses.

Figure 1. RECORD flow diagram of providers into the primary panels



Sequential exclusions as in Table 1; counts are providers unless stated.

The selected CT panel's summed published activity was 8,959,125 examinations (3,985,250 in the OP/GP group) and the MRI panel's 4,362,085 (3,589,375) (Table 2). Within providers, activity varied little: the within-provider standard deviation of $\log_2$ activity was 0.087 for CT and 0.134 for MRI (about 6.2% and 9.8%), no CT provider and one MRI provider changed by a full doubling across the eight quarters, and 98.6% of CT and 93.1% of MRI residualised exposures lay within $\pm\log_2 1.1$, that is, within a 10% change.

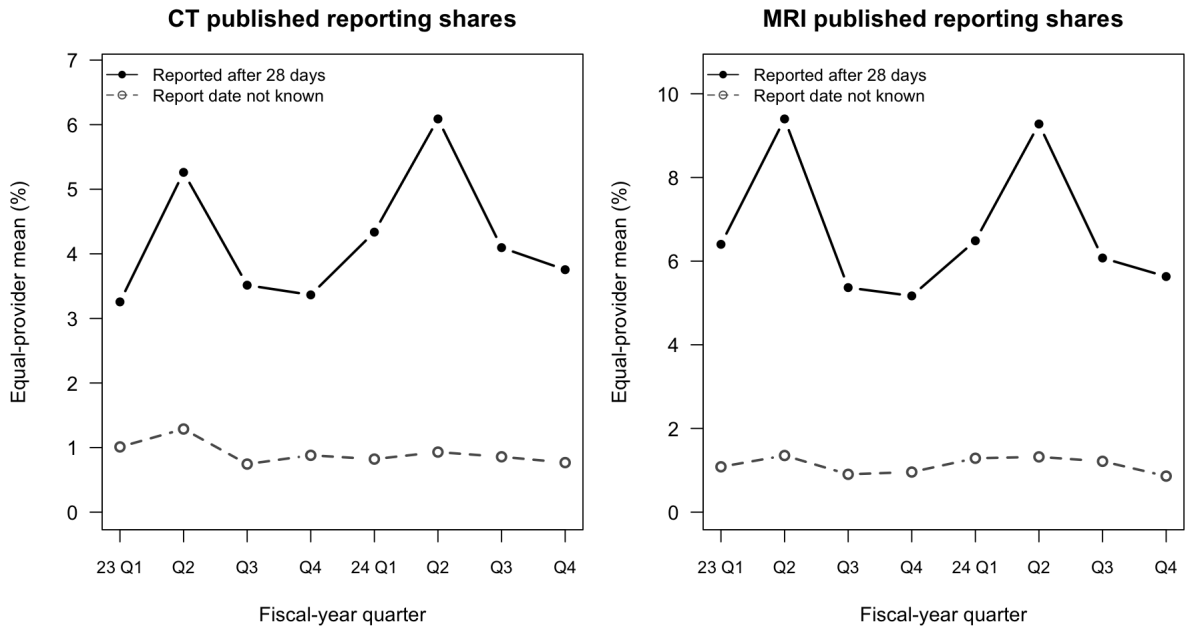
Table 2. Published activity and reporting fractions in the primary panels

Quantity	CT	MRI
Providers / provider–quarters	70 / 560	67 / 536
Summed combined examination activity	8,959,125	4,362,085
Summed OP/GP activity	3,985,250	3,589,375
Median quarterly combined activity	13,995.0	6,737.5

Quantity	CT	MRI
Quarterly activity range	1,115–46,830	1,475–25,300
Mean published late share (%)	4.21	6.73
Mean unknown-date share (%)	0.91	1.12
Within-provider SD of $\log_2$ activity (≈% change) †	0.087 (6.2%)	0.134 (9.8%)
Provider range of $\log_2$ activity, median (IQR) †	0.21 (0.17–0.27)	0.32 (0.22–0.43)
Providers changing by ≥ a doubling (range ≥1) †	0 of 70	1 of 67
Residualised exposures within $\pm\log_2 1.1$ (±10%) †	98.6%	93.1%
Residualised exposure, minimum to maximum †	−0.48 to 0.14	−0.41 to 0.45

The mean published late share was 4.21% for CT and 6.73% for MRI, and the mean unknown-date share 0.91% and 1.12%. Late shares peaked in the July–September quarter of both years (CT 5.26% and 6.09%; MRI 9.40% and 9.28%; Figure 2). The quarter effects absorb this shared seasonal pattern, and allowing each provider its own seasonal pattern changed the estimates little (see robustness analyses below).

Figure 2. Quarterly published late and unknown-date shares



Equal-provider mean OP/GP percentages for 70 CT and 67 MRI providers. Solid circles, reported after 28 days; open circles, unknown report date. Quarters: Q1 April–June, Q2 July–September, Q3 October–December, Q4 January–March.

Within-provider and between-provider associations

The adjusted within-provider coefficient was 8.49 percentage points per doubling of CT activity (95% cluster CI −2.64 to 19.62) and 7.92 for MRI (−2.86 to 18.70). Expressed for realistic changes, a 10% increase in activity was associated with 1.17 (−0.36 to 2.70) and 1.09 (−0.39 to 2.57) percentage points,

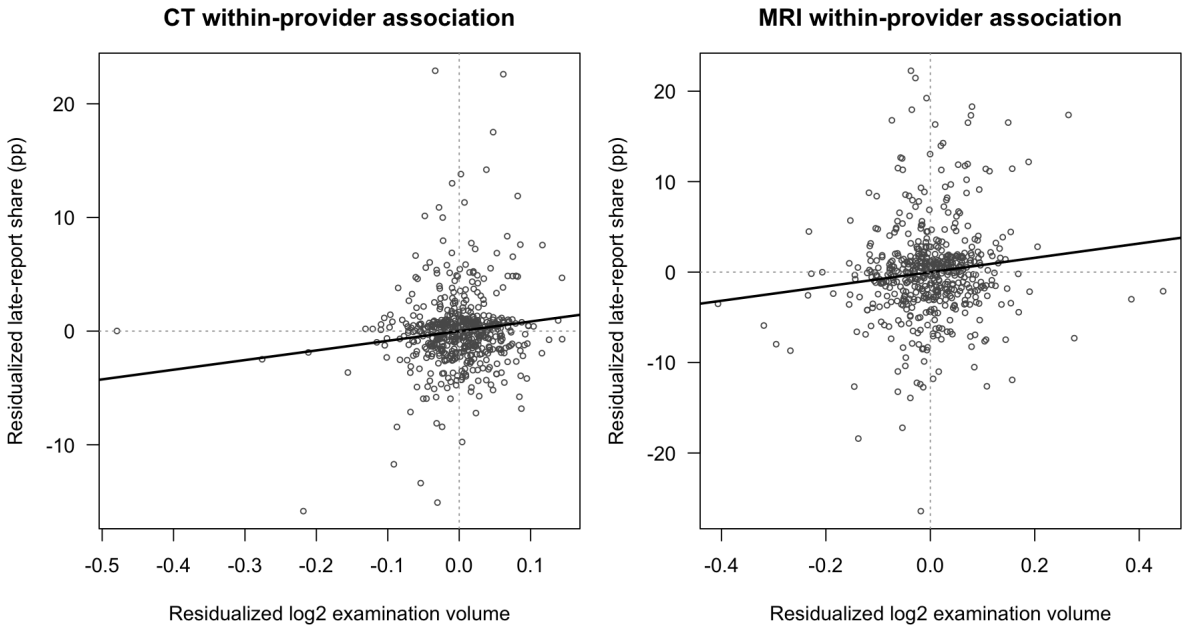
and a one-standard-deviation within-provider increase with 0.74 (−0.23 to 1.70) and 1.06 (−0.38 to 2.51) (Table 3). The seeded provider-bootstrap intervals (−0.98 to 20.16 and −1.86 to 18.35; Supplementary Code S1) and the wild cluster bootstrap-t intervals (−1.30 to 21.60, p = 0.12; −2.45 to 19.10, p = 0.14) also included zero. Without the acute-setting share the coefficients were 6.82 (−0.54 to 14.18) for CT and 9.28 (1.20 to 17.37) for MRI; the MRI interval excluded zero.

Table 3. Within-provider and between-provider associations

Model quantity	CT	MRI
Within volume coefficient per doubling (95% cluster CI)	8.49 (−2.64 to 19.62)	7.92 (−2.86 to 18.70)
Within volume coefficient per 10% higher activity (95% cluster CI) †	1.17 (−0.36 to 2.70)	1.09 (−0.39 to 2.57)
Within volume coefficient per within-provider SD (95% cluster CI) †	0.74 (−0.23 to 1.70)	1.06 (−0.38 to 2.51)
Within, acute share omitted, per doubling (95% cluster CI) †	6.82 (−0.54 to 14.18)	9.28 (1.20 to 17.37)
Within volume: 95% provider-bootstrap interval	−0.98 to 20.16	−1.86 to 18.35
Within volume: 95% wild cluster bootstrap-t interval; p for zero †	−1.30 to 21.60; p = 0.117	−2.45 to 19.10; p = 0.138
Acute-share coefficient (cluster SE)	9.15 (21.17)	−12.53 (30.72)
Between volume coefficient (95% bootstrap interval)	0.61 (−0.40 to 1.68)	2.62 (1.05 to 4.23)
Between minus within: 95% paired bootstrap interval	−18.53 to 1.14	−15.89 to 4.97
Cluster degrees of freedom / full model rank	69 / 79	66 / 76

The acute-share coefficients were imprecise (CT 9.15, SE 21.17; MRI −12.53, SE 30.72). Figure 3 plots the residualised exposure and outcome. The most influential provider-quarter was RTE 2023-24 Q4 for CT (coefficient without it 5.78) and RJN 2024-25 Q2 for MRI (6.50). Removing the 29 CT and 35 MRI provider-quarters with |DFBETAS| > 2/√N reduced the coefficients to 1.55 (−3.04 to 6.14) and 4.59 (−0.83 to 10.00), whereas removing the five most extreme residualised exposures gave 6.38 and 8.74 (Table 5).

Figure 3. Residualised within-provider activity and reporting share



Each point is a provider–quarter, after removing provider and quarter effects and acute share; the line is the primary least-squares fit. Most points lie within ±0.1 on the horizontal axis (about ±7% activity): 96% for CT and 84% for MRI.

The between-provider coefficients were 0.61 percentage points for CT (bootstrap interval −0.40 to 1.68) and 2.62 for MRI (1.05 to 4.23). The paired intervals for the between-minus-within difference (−18.53 to 1.14 and −15.89 to 4.97) included zero.

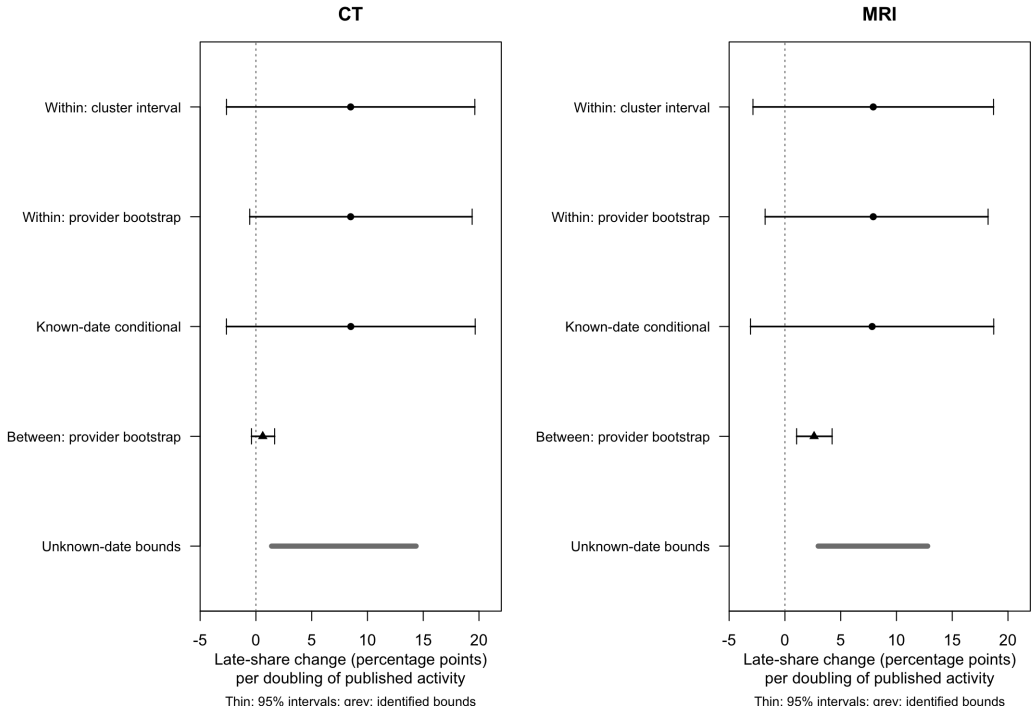
Measurement bounds and planned robustness

Conditioning on known report dates changed the coefficients little (CT 8.51, −2.65 to 19.66; MRI 7.82, −3.08 to 18.72). Unknown-date bounds were 1.41 to 14.36 for CT and 2.99 to 12.79 for MRI, and the half-display-unit ranges 7.85 to 9.13 and 7.48 to 8.36 (Table 4; Figure 4).

Table 4. Alternative date denominators, measurement bounds and expanded cohorts

Analysis or measurement contrast	CT	MRI
Known-date coefficient (95% cluster CI)	8.51 (−2.65 to 19.66)	7.82 (−3.08 to 18.72)
Unknown-date identified coefficient bounds	1.41 to 14.36	2.99 to 12.79
Half-display-unit coefficient sensitivity	7.85 to 9.13	7.48 to 8.36
Unknown ≤20%: providers; coefficient (95% cluster CI)	82; 6.45 (−3.60 to 16.51)	79; 4.83 (−2.67 to 12.32)
No unknown cutoff: providers; coefficient (95% cluster CI)	99; 3.79 (−2.29 to 9.87)	103; 3.83 (−2.41 to 10.07)
No unknown cutoff: per 10% higher activity †	0.52 (−0.31 to 1.36)	0.53 (−0.33 to 1.38)
No unknown cutoff: unknown-date share on activity †	−1.34 (−8.66 to 5.98)	3.82 (−0.88 to 8.52)
No unknown cutoff: late-or-unknown share on activity †	2.45 (−7.11 to 12.01)	7.65 (−0.08 to 15.38)

Figure 4. Sampling uncertainty and missing-date identification ranges



Thin bars, 95% cluster or provider-bootstrap intervals; circles, within estimates; triangles, between estimates; thick grey bars, unknown-date identification bounds. Horizontal units are percentage points per doubling; the dotted line marks zero.

Relaxing the unknown-date limit to 20% gave 82 CT and 79 MRI providers and coefficients of 6.45 and 4.83; removing it gave 99 and 103 providers and 3.79 and 3.83, about half the primary estimates (0.52 (−0.31 to 1.36) and 0.53 (−0.33 to 1.38) per 10%). The 29 CT and 36 MRI providers excluded by the cutoff had a mean unknown-date share of 12.3% and 12.7%. In the expanded cohorts, the unknown-date share was not clearly associated with activity (CT −1.34 (−8.66 to 5.98); MRI 3.82 (−0.88 to 8.52)), and the combined late-or-unknown share gave 2.45 (−7.11 to 12.01) and 7.65 (−0.08 to 15.38). Weighting by the OP/GP denominator gave 11.19 (−4.13 to 26.51) and 7.76 (−4.96 to 20.49); in the fractional logit model the average marginal effect of a 10% increase was 1.36 (0.16 to 2.56) and 1.15 (−0.31 to 2.61) percentage points, the CT interval excluding zero. First differences gave 10.46 and 8.94, and leave-one-provider-out estimates ranged from 4.46 to 10.87 and 5.33 to 9.60 (Table 5).

Table 5. Specification, influence and weighting checks

Specification or diagnostic	CT	MRI
First-difference coefficient	10.46	8.94
Leave-one-provider-out coefficient range	4.46 to 10.87	5.33 to 9.60
Without the most influential provider-quarter (DFBETA) †	RTE 2023-24 Q4: 5.78	RJN 2024-25 Q2: 6.50
Without provider-quarters with |DFBETAS| > 2/√N (n removed) †	1.55 (−3.04 to 6.14); n = 29	4.59 (−0.83 to 10.00); n = 35
Without the five most extreme residualised exposures †	6.38 (−4.75 to 17.52)	8.74 (−3.84 to 21.32)
Without the provider with the largest activity range †	RJ6: 10.87 (−1.62 to 23.37)	RNQ: 6.02 (−4.45 to 16.49)

Specification or diagnostic	CT	MRI
Provider × quarter-of-year effects added †	7.27 (−7.61 to 22.16)	8.57 (−10.34 to 27.48)
Weighted by OP/GP denominator, per doubling †	11.19 (−4.13 to 26.51)	7.76 (−4.96 to 20.49)
Weighted by OP/GP denominator, per 10% †	1.54 (−0.57 to 3.64)	1.07 (−0.68 to 2.82)
Fractional logit: log-odds per doubling †	2.61 (0.31 to 4.92)	1.45 (−0.39 to 3.30)
Fractional logit: average marginal effect per 10%, pp †	1.36 (0.16 to 2.56)	1.15 (−0.31 to 2.61)
Condition number of the 2 × 2 within cross-product matrix (design-matrix value squared)	13.48 (3.67²)	90.33 (9.50²)

Discussion

Principal findings

In complete R-coded reporting panels, a 10% within-provider increase in CT or MRI activity was associated with about one percentage point more OP/GP examinations reported after 28 days, against mean late shares of 4.2% and 6.7%. The intervals were wide and included zero, and the estimates were sensitive to the report-date completeness rule and to a small set of provider-quarters. The results are compatible with no association as well as with a moderate one, and do not show that reporting congestion rises whenever activity grows.

The between-provider coefficient was small for CT and positive for MRI, and the between-minus-within intervals included zero. The CT result agrees with the official 2024/25 release, which found no tendency for longer test-to-report periods in higher- or lower-activity providers [3]; for MRI our between-provider estimate differs, although the release examined brain MRI only and median intervals rather than the 28-day tail. What the within-provider design adds is a comparison of each provider with itself, which removes stable differences in case mix, configuration and record quality. It does not remove changes in staffing or technology during the eight quarters.

Comparison with other evidence

Delivery changes such as outsourcing, teleradiology and remote reporting can raise reporting capacity [12,14], and cross-organisation reporting is common [13]; such arrangements may explain why a provider's reporting tail moves little when its own activity changes, but they are not recorded in the DID. The quality of outsourced reports is a separate concern [15]. Detailed workflow studies [16–19] and national workforce reports [9,10] provide operational context that the quarterly workbooks cannot supply. Linking provider-level staffing, outsourcing and priority data to the DID would be needed to test these mechanisms.

Limitations

Unknown report dates. The known-date analysis changes the denominator but does not correct missing data; dates may be absent because a report is incomplete, was not incorporated, or was submitted incorrectly. The unknown-date bounds hold the realised design fixed and reallocate only the unknown

share; their positive lower endpoints do not contradict the confidence intervals that include zero, which describe sampling uncertainty. The half-display-unit sensitivity was narrower than the unknown-date ranges.

Selection. Requiring complete records and a low unknown-date share selected providers with better report-date recording, and removing that requirement halved the estimates. Report-date completeness is itself a reporting-process attribute, so the direction of bias is uncertain. In the expanded cohorts the unknown-date share did not clearly rise with activity, but for MRI the combined late-or-unknown outcome (7.65) was close to the primary estimate, consistent with some late reports being recorded as unknown when activity rises, which would bias the expanded-cohort estimate towards zero. For CT the combined outcome (2.45) did not support this explanation; the difference may instead reflect a weaker association among providers with less complete recording, to whom the primary estimate would then not generalise. The restriction to R-coded providers excludes independent-sector segments and later code formats [25].

Exposure and specification. Completed activity reflects both demand and acquisition capacity, and reporting capacity can respond through scheduling or outsourcing. Within-provider variation was small, so per-doubling coefficients extrapolate well beyond the data; the estimate depended on a few provider-quarters, and one of the added sensitivity analyses (fractional logit for CT) and the MRI model without acute share had intervals excluding zero. Given the number of analyses, we do not interpret these as confirmation. Published fractions and totals are rounded, and the 28-day tail cannot monitor hour-based or cancer-pathway targets or communication of results [7,29].

Implications

Public monitoring should present examination volume, the late share, the unknown-date share and the eligible denominator together, and state whether a contrast is within providers, between providers or a change in national mix. The present estimates are too imprecise to support a provider-capacity threshold. A completed report is not proof of communication or follow-up, and the study has no patient-level endpoint [29]. Future work could link verified staffing, outsourcing, complexity and priority data to a longer provider series, with explicit report-date completeness.

Conclusion

Within complete R-coded CT and MRI panels, higher quarterly activity had positive but imprecise associations with the outpatient/general-practitioner share reported after 28 days, about one percentage point per 10% more activity, with intervals including zero. Without the unknown-date cutoff the estimates roughly halved, and they depended on a few provider-quarters. The findings support monitoring activity, late share, unknown dates and denominator together, and do not establish a causal congestion effect, national performance, referral-specific compliance or patient benefit.

Declarations

This study used only published aggregate official statistics; no individual-level or confidential data were accessed.

Data availability: The data are publicly available from NHS England as the Diagnostic Imaging Dataset Table 13 workbooks for 2023/24 (finalised version published July 2025; https://www.england.nhs.uk/statistics/wp-content/uploads/sites/2/2025/07/DID-Table-13-2023-24-Modality-Provider-Report-turnaround-XLSX-868KB.xlsx) and 2024/25 (published November 2025; https://www.england.nhs.uk/statistics/wp-content/uploads/sites/2/2025/11/DID-Table-13-2024-25-Modality-Provider-Report-

Turnaround-XLSX-866KB.xlsx). The SHA-256 checksums of the files analysed, the analysis code, its output and the full coefficient set are in Supplementary Code S1.

Ethics: This study was exempt from ethics review because the data are public (secondary analysis of public, aggregate data).

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