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# Sewing terror: price dynamics of the strawberry needle crisis

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In September 2018, a former strawberry-picking supervisor for a strawberry farm in New South Wales was arrested for inserting sewing needles in hundreds of punnets of fresh strawberries sold in retail stores across several Australian states. This paper analyses the 2018 Australian strawberry needle scare as a case study on the market impacts of agro-terrorism events. We develop a novel four-step procedure to estimate the effects of the strawberry needle contamination on wholesale fresh strawberry prices. Our results indicate a drop in wholesale fresh strawberries prices of about 20% while the needle crisis was ongoing. However, public and private supply restrictions caused wholesale prices to rise by up to 94% relative to expected dynamics over several weeks in the immediate wake of the incident.

**Key words:** 2018 strawberry needle scare, agro-terrorism, price analysis.

## 1. Introduction

> A single act of commercial terrorism has brought a multi-million dollar industry to its knees. Adrian Schultz, Vice President, Queensland Strawberry Growers Association (2018).

Adulteration of the food supply requires no particular expertise or technology, and the economic impacts are large. These factors make it unsurprising that agriculture was weaponised in both World Wars. In World War I, German agents used anthrax and glanders to infect livestock and pack animals; in World War II, Japan attempted to use airdrops of infected plants to destroy its enemies' grain and vegetable supplies (Turvey *et al*. 2003). For the same reasons, the agri-food economy is also particularly vulnerable to acts of domestic terrorism (Wheelis, Casagrande and Madden 2002). Agro-terrorism – defined as the intentional contamination of food products with harmful substances – can destroy agricultural businesses and raze entire industries.

The motivations underlying domestic agro-terrorism are diverse. Many modern incidents have been motivated by politics (Howitt 2003; Turvey *et al*. 2003). Examples include the 1978 citrus contamination in Israel,[1] the 1984

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[1] In 1978, a Palestinian group injected mercury into Israeli citrus exports into Europe (Israel's biggest earner of foreign exchange).


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salad bar contamination in Oregon[2] and the 1989 table grape contamination in Chile.[3] Other agro-terrorism incidents were motivated by personal revenge such as 1997 pet food contamination in Wisconsin[4] and a 2002 incident in Nanjing, China.[5] Yet, in spite of the complex motivations and global costs imposed by agro-terrorism, economic literature on the issue is scarce. In the aforementioned examples, the intentional, deliberate nature of the incident is unknown to the public at the outset of the scare. Instead, only the resulting symptoms and (possibly) the affected product or locale are known while the threat is ongoing. Consequently, market reactions to this type of event are likely similar to those of an unintentional food scare, such as the recent *E. coli* outbreaks in romaine lettuce across the United States (KSBW 2018).

In contrast to these crises, this paper is intended as a case study on the market impacts of an agro-terrorism incident in which the terrorist motivations are known at the time of the event. We estimate the impact of the 2018 Australian strawberry needle scare on wholesale strawberry prices. In September 2018, hundreds of punnets of fresh strawberry were found to contain partially embedded sewing needles. The scare spanned several weeks and concerned several Australian states. The incident also spawned several hoaxes and copycats. In November 2018, a former strawberry-picking supervisor for a strawberry farm in New South Wales – allegedly motivated by spite against her former employer – was arrested for the contamination. Investigators assert that the needles were planted with the intent to cause economic harm to the Australian strawberry industry.

We construct an econometric model using weekly price data for calendar years 2014–2018 to estimate price impacts. We believe this research makes both substantive and methodological contributions to the literature. To the authors' knowledge, this is the first paper to study the economic impacts of an actual agro-terrorism incident. Oladosu *et al.* (2013) use a computable general equilibrium framework to simulate the hypothetical impacts of the intentional introduction of foot-and-mouth disease virus in the United States. The majority of the previous agricultural economics research on agro-terrorism focuses on the relationship between fear and risk perceptions from the consumer resulting from a nonspecific agro-terror threat (Govindasamy *et al.* 2006; Turvey *et al.* 2007, 2003). Govindasamy *et al.* (2006) focus on consumers' substitution towards locally grown produce as a risk-avoidance strategy due to terrorism fears. Turvey *et al.* (2007) conduct a consumer survey on attitudes regarding food-system vulnerability to agro-terrorism.

[2] In 1984, members of a fringe religious group in Oregon contaminated salad bars with Salmonella to prevent residents from voting in local elections.

[3] Anti-Pinochet separatists injected Chilean table grapes with cyanide in 1989.

[4] The owner of a failing animal food processing facility deliberately contaminated Purina feed supplies with harmful pesticides in Wisconsin in 1997.

[5] In 2002, a shopkeeper in Nanjing, China, driven by jealousy spiked foods from a rival business with rat poison, killing 38 people.

Because the prior analyses focus on nebulous and nonspecific threats, their value regarding active or immediately resolved agro-terror incidents may be limited. It is straightforward to anticipate consumer substitution away from a food product targeted by an agro-terror event, thereby causing prices to drop. However, the expected duration of this downward price pressure is unknown. Moreover, consumer substitution is only part of the total market effect. Also important to the market price effects are producer and government responses. Producers take steps to preserve reputation, and government efforts attempt to mitigate economic and public health consequences. We document these latter responses in the light of the strawberry needle contamination incident and show that – in this case – they dominate the total market effect. We identify a drop in wholesale fresh strawberries prices of about 20% while the needle crisis was ongoing. However, perhaps counter to expectations, wholesale prices rose by up to 94% relative to expected dynamics over several weeks in the immediate wake of the incident, likely due to public and private damage mitigation strategies.

From a methodological perspective, our estimation procedure extends state-of-the-art event study techniques. We develop a four-step procedure to estimate the effects of the strawberry needle contamination on wholesale fresh strawberry prices. Similar to the relative-price-of-a-substitute (RPS) technique developed by Carter and Smith (2007), our model relies on identification by comparing price dynamics for the subject product relative to those for unaffected substitute products. However, in contrast to the RPS approach, our methodology allows for highly seasonal price data and an endogenously defined group of multiple substitute products. The method can be applied to study the impacts of policies and other events affecting fresh produce markets or other seasonal products.

The remainder of the paper is organised as follows. Section 2 provides a background of the 2018 strawberry needle incident. Section 3 presents our price impact estimation methodology. Section 4 presents results, and Section 5 concludes.

## 2. 2018 strawberry needle scare

Australian farmers grow around 93.5 thousand metric tonnes of strawberries each year, with an annual value of production of approximately $500 million (Horticulture Innovation Australia Ltd 2019). Strawberries are grown in almost all Australian states, and the two largest producing regions – Beerwah in Queensland and Yarra Valley in Victoria – account for more than 77% of national production (Horticulture Innovation Australia Ltd 2019). Almost three-fourths of Australian households regularly purchase fresh strawberries, and over 80% of national strawberry production goes to domestic fresh supply. The remaining portion goes to processing (14%) and fresh export (5%) (Horticulture Innovation Australia Ltd 2019). Imports (typically from China or South Korea) equate to < 0.1% of fresh supply.

Climatic conditions in Australia allow for year-round availability of fresh strawberries (Horticulture Innovation Australia Ltd 2019). Harvest in

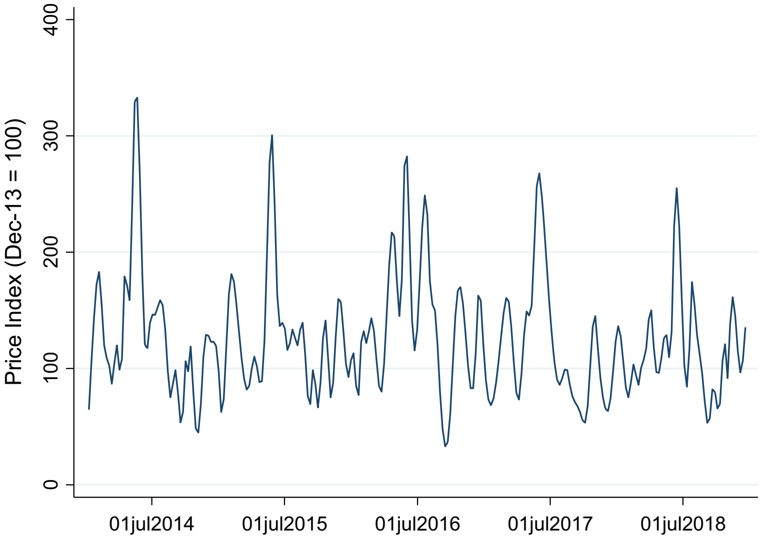



**Figure 1** Seasonality of wholesale strawberry prices in Melbourne.
Source: Australian Bureau of Agricultural and Resource Economics (2019), 'Movements in selected fruit and vegetable prices' *Weekly Australian Climate, Water and Agricultural Update*, available at http://www.agriculture.gov.au/abares/publications/weekly_update/movements-selected-fruit-vegetable-prices#kiwifruit-hayward-strawberry-apple-royal-gala-and-avocado-hass. [Colour figure can be viewed at wileyonlinelibrary.com]

Queensland – the highest producing region (42% of national production) – begins in March and continues through late November, with highest availability between July and September. Harvest in Victoria (36% of national production) immediately follows Queensland harvest, beginning in October and continuing through June, with highest availability between November and March. This staggered availability results in highly seasonal price patterns for fresh strawberries. Figure 1 plots weekly wholesale prices for fresh strawberry prices in Melbourne for calendar years 2014–2018. Prices tend to reach a peak around the 21$^{st}$ week of the calendar year (mid-May, following Mother's Day) and a trough around the 37$^{th}$ week of the year (early September).

On 9 September 2018, a man from Moreton Bay, New South Wales, was taken to hospital after swallowing part of a needle found in a punnet of Berry Obsession brand strawberries (grown in Queensland) purchased from Woolworths. In the days following this first incident, dozens of contaminated strawberries grown in Queensland and Western Australia were discovered in New South Wales, South Australia, Victoria and Tasmania (Williams 2019). Within one week of the initial incident, at least eight brands of strawberries were found to have been contaminated (Maunder 2018).[6] In total, there were 186 reports of contamination nationally, including several hoaxes and copycat incidents (BBC News 2018).

[6] These brands were Australian Choice, Berry Licious, Berry Obsession, Donnybrook Berries, Delightful Strawberries, Love Berry, Mal's Black Label and Oasis Brands (Maunder 2018).

Many affected growers deemed the incident to be catastrophic. Queensland and Western Australian farmers disposed of their harvested strawberries. Some farmers burned off unsellable strawberries still in the fields. Others simply ceased operations for the year and left unharvested strawberries to rot. Still other growers chose to proceed with harvest after installing metal detectors to certify that their strawberries were free from contamination (Weber and Prendergrast 2018). Responses at the retail level were mixed. In the wake of the initial incidents, Coles and Aldi removed all strawberries from their shelves, but restocked after 18 September. Woolworths continued to stock strawberries throughout and removed only affected brands; it also removed sewing needles from sale for a short time (Bavas 2018).

In response to the contamination, the premiers for Queensland and Western Australia each issued $100,000 rewards for information leading to the arrest of the saboteur (Colvin 2018; Crockford 2018). The Queensland Government also issued a $1 million assistance package for strawberry growers in the state (Doherty, 2018). The Federal Government issued a recall of affected brands from distribution centres and wholesalers. It did not issue consumer food recalls. In fact, even in the midst of the incident, government officials continued to encourage the purchase and consumption of strawberries. Queensland Chief Health Officer, Dr. Jeannette Young, said ‘strawberries are a great fruit, a great produce to eat. But at this stage, please cut them up and just look to make sure they haven’t been contaminated’ (Williams 2019).

On 20 September 2018, the Australian parliament voted to increase the maximum jail term from 10 to 15 years for offenders convicted of food tampering (Reuters 2018). On 11 November 2018, My Ut Trinh – a 50-year-old woman, who had formerly worked as a strawberry-picking supervisor for the Berry Obsession fruit farm north of Brisbane, was arrested in Brisbane and charged with seven counts of contaminating goods. The arrest was made after her DNA was found on contaminated strawberries in Victoria. Queensland police allege that Ms. Trinh was motivated by spite at her former employer (BBC News 2018).

## 3. Methodology

We employ a four-step procedure to estimate the effects of the strawberry needle contamination on wholesale fresh strawberry prices. At any given time $t$, fresh strawberry prices and prices for other fresh fruits are comprised of three components: $Ln\, P_t^k = \theta_t^k + \eta_t^k + \varepsilon_t^k$ where $P_t^k$ is the price of product $k$ in time $t$. The first component is a seasonal oscillation, resulting from natural agroecological and climate patterns in Australia and abroad. For a fresh fruit product $k$, we denote this component of the price $\theta_t^k$. The second component (denoted $\eta_t^k$) generates deviations from the seasonal oscillation as a result of economic information linked to time-specific supply and demand conditions. The final component (denoted $\varepsilon_t^k$) is a random, independent and identically distributed shock with zero mean.

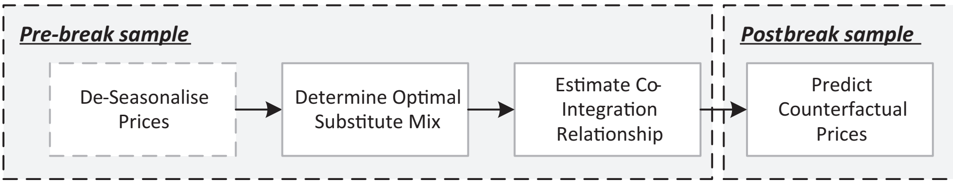



**Figure 2** Impact estimation methodology.

Similar to the relative-price-of-a-substitute (RPS) technique developed by Carter and Smith (2007), our model relies on identification by comparing price dynamics for fresh strawberries relative to those for unaffected substitute products. As discussed in Carter and Smith (2007), this methodology requires a stable relationship between the affected strawberry price and the benchmark substitute price. Here, the relationship between fresh strawberry prices and those for substitute products is undoubtedly nonstable due to seasonal oscillation ($\theta_t^k$). Accordingly, our analysis relies on the second component of the strawberry price, $\eta_t^S$, as the component of primary interest for the purposes of our empirical strategy. Model identification hinges on the existence of some set of imperfectly substitutable fresh fruits ($j$) for which the correlation between $\eta^S$ and $\eta^j$ is nonzero.

The schematic diagram in Figure 2 explains our four-step estimation procedure. We first de-seasonalise our strawberry price series with a nonlinear least-squares regression using a series of annual and semi-annual Fourier terms. Second, we choose the optimal bundle of imperfectly substitutable products (for which Corr $(\eta_t^S, \eta_t^j) \neq 0$) to use as controls against which to measure the price impact. Third, we estimate the historical co-integrating relationship between the affected strawberry price component ($\eta_t^S$) and those of the imperfect substitutes ($\eta^j$) observed prior to the needle event. Finally, we deduce the dynamic price impact by comparing actual strawberry prices with those that would have resulted had the historical co-integrating relationship continued in the wake of the needle event.

For this analysis, we combine the weekly, Melbourne wholesale fresh strawberry price data from the Australian Bureau of Agricultural and Resource Economics (ABARES) *Weekly Australian Climate, Water and Agricultural Update* depicted in Figure 1 with weekly, Melbourne wholesale prices for fresh blueberries and kiwifruit, also obtained from ABARES.[7] All prices run from January 2014 to December 2018 and are indexed such that the price in December 2013 equals 100. Our raw dataset includes 14 missing observations for the kiwifruit price series (≈5% of observations) and 4 missing observations for blueberry prices (≈2% of observations). Because vector error correction (VEC) modelling – used to generate the historical co-integrating relationship – requires that the dataset does not

[7] Kiwifruit and blueberries were chosen as substitutes as they both exhibit high degrees of seasonality and perishability like strawberries.

contain time gaps, we interpolate these missing observations using cubic splines.

### 3.1 De-seasonalising prices

For step one of the estimation procedure, we seek to remove the seasonal oscillation component ($\theta_t$) from each price series in order to test for a stable relationship between price series. We approximate $\theta_t$ for each price series by fitting prices to a set of annual and semi-annual Fourier terms. We estimate the following equation via nonlinear least squares:

$$\begin{aligned} Ln\,P_t = v_t + b_0 + b_1 \cos\left(\frac{2\pi}{52}t + b_2\right) + b_3 \sin\left(\frac{2\pi}{52}t + b_4\right) \\ + b_5 \cos\left(\frac{4\pi}{52}t + b_6\right) + b_7 \sin\left(\frac{4\pi}{52}t + b_8\right) \end{aligned} \tag{1}$$

where sinusoidal terms with frequency $\frac{2\pi}{52}$ generate an annual oscillation and terms with frequency $\frac{4\pi}{52}$ generate a semi-annual oscillation. Note that this specification does not specifically identify the component of interest: $\eta^t$ Instead, variable $v_t$ from Equation (1) is equal to $\eta_t + \varepsilon_t$.

Nonlinear least-squares estimations require the designation of initial guesses for all parameters. We initialise the regression by first estimating Equation (1) via ordinary least squares (OLS) under the assumption that the phase shift is zero for each oscillation term; that is, initial values (denoted superscript *I*) are $b_i^I = \hat{b_i^{OLS}} \forall i \in \{0, 1, 3, 5, 7\}$ and $b_i^I = 0 \forall i \in \{2, 4, 6, 8\}$. We repeat this procedure for strawberry, blueberry and kiwifruit prices.

Table 1 reports coefficient estimates for Equation (1) and the initial values obtained via OLS. Note that, as reported in Columns (2), (4) and (6), we cannot obtain standard error estimates for amplitude and phase shift parameters separately in the NLLS specifications. This is of little concern because our interest is in the precision of fit of the overall oscillation pattern ($\hat{\theta}_t$), not the specific coefficient estimates. Using the coefficient estimates reported in Columns (2), (4) and (6) of Table 1, we generate oscillation patterns for fresh strawberry, kiwifruit and blueberry prices. These oscillation patterns (and corresponding confidence intervals) are plotted in Figure 3.

We de-seasonalise each price series by subtracting the estimated seasonal oscillation patterns depicted in Figure 3 ($Ln\,P_t - \hat{\theta}_t$). Arithmetically, this leaves $\hat{v}_t$—a nonseparable estimate of the final two components of each price series $\eta_t + \varepsilon_t$.

### 3.2 Tests for optimal substitute mix

We perform the remainder of the econometric analysis on the de-seasonalised prices $\hat{v}_t$, comprised of price components $\eta_t$ and $\varepsilon_t$. The next

**Table 1** Seasonal price oscillation estimates

| Coefficient | Strawberries | | Kiwifruit | | Blueberries | |
|---|---|---|---|---|---|---|
| | (1) OLS | (2) NLLS | (3) OLS | (4) NLLS | (5) OLS | (6) NLLS |
| b0 | 4.753*** | 4.753*** | 4.497*** | 4.497*** | 5.121*** | 5.121*** |
| | (0.021) | (0.029) | (0.014) | (0.014) | (0.013) | (0.012) |
| b1 | −0.207*** | −0.204*** | −0.067*** | −0.074*** | −0.447*** | −0.445*** |
| | (0.029) | (0.029) | (0.022) | (0.020) | (0.019) | (0.017) |
| b2 | | 0.004 | | −0.063 | | −0.031 |
| | | – | | – | | – |
| b3 | 0.181*** | 0.180*** | 0.270*** | 0.275*** | 0.083*** | 0.097*** |
| | (0.030) | (0.030) | (0.022) | (0.025) | (0.017) | (0.018) |
| b4 | | −0.020 | | 0.025 | | −0.032 |
| | | – | | – | | – |
| b5 | 0.197*** | 0.198*** | 0.068*** | 0.071*** | 0.110*** | 0.108*** |
| | (0.028) | (0.029) | (0.022) | (0.020) | (0.020) | (0.018) |
| b6 | | 0.012 | | 0.011 | | 0.029 |
| | | – | | – | | – |
| b7 | −0.077** | −0.075*** | −0.045** | −0.044** | −0.100*** | −0.097*** |
| | (0.031) | (0.029) | (0.018) | (0.020) | (0.016) | (0.0180) |
| b8 | | 0.016 | | 0.084 | | −0.024 |
| | | – | | – | | – |
| Observations | 239 | 239 | 237 | 237 | 239 | 239 |

***$P < 0.01$, **$P < 0.05$, *$P < 0.1$.
Robust standard errors in parentheses.

step in our analysis is to determine the optimal set of substitutes to use as controls in the price impact determination. Recall that our identification strategy hinges on identifying one or more substitute products $j$ for which the correlation is nonzero between the strawberry price component containing economic information and that of the substitute products (Corr $(\eta_t^S, \eta_t^j) \neq 0$). Should we specify as control prices kiwifruit only ($\hat{v}_t^K$), blueberry prices only ($\hat{v}_t^B$) or both kiwifruit and blueberry prices?

We begin this assessment by first analysing the time-series properties of the de-seasonalised prices. Vector error correction (VEC) modelling – used to measure the price impact – and the concept of price co-integration require that the treated and control price series are integrated of the same order (Engle and Granger 1987). Table 2 reports results of augmented Dickey–Fuller tests for nonstationarity on the de-seasonalised strawberry, kiwifruit, and blueberry prices. These tests are conducted only on the pre-event data (i.e. the subsample running from the beginning of the sample period through the first week of August 2018). This subsample is chosen to avoid weeks that were affected by the needle contamination in order to omit possible changes to the data generating process resulting from the event. As shown in Table 2, we reject the null hypothesis of nonstationarity at 99.99% for the strawberry and blueberry series and at 90% for the kiwifruit series. This is strong

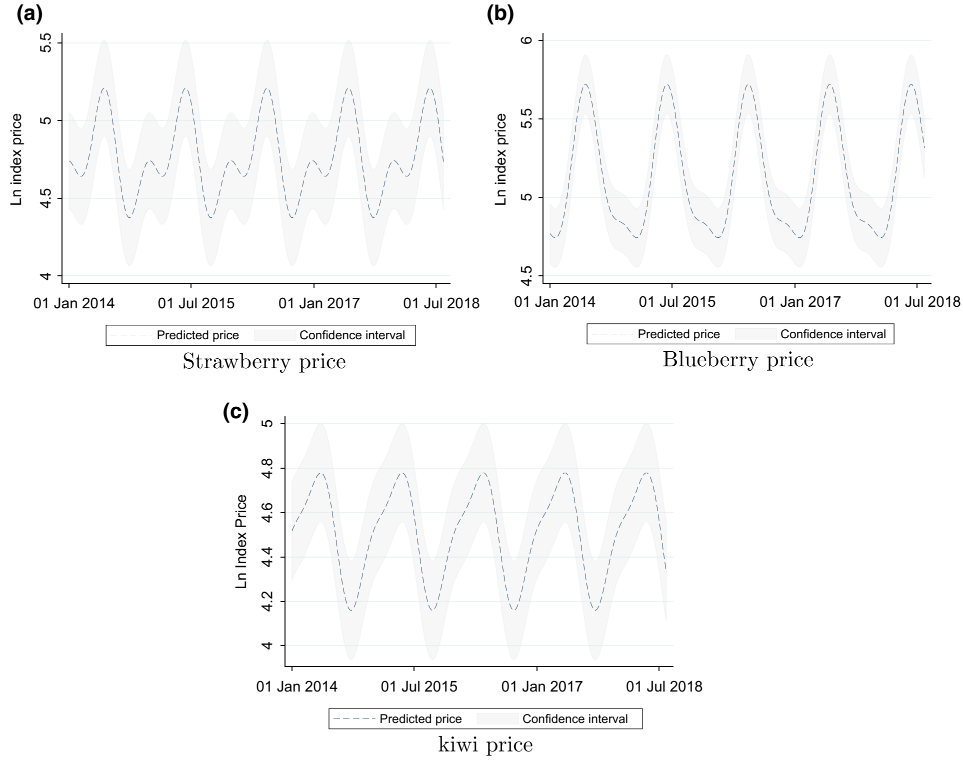



**Figure 3** Seasonal price oscillation estimates. [Colour figure can be viewed at wileyonlinelibrary.com]

**Table 2** Augmented Dickey–Fuller tests for de-seasonalised prices

| | Strawberries ($\hat{v}^S$) | Kiwifruit ($\hat{v}^K$) | Blueberries ($\hat{v}^B$) |
|---|---|---|---|
| Obs | 234 | 232 | 234 |
| $T$-Statistic | −7.350 | −2.748 | −5.661 |
| $P$-value | 0.000 | 0.066 | 0.000 |
| Conclusion | Reject $H_0$ | Reject $H_0$ | Reject $H_0$ |

All tests conducted with 4-lag specification.
5% MacKinnon Critical Value is − 2.881.

evidence that all series are integrated of order I(0). The VEC procedure is consistent for stationary variables (Engle and Granger 1987).

To formally determine the optimal substitute mix, we estimate a VEC model, including our de-seasonalised fresh strawberry prices ($\hat{v}_t^S$) and the de-seasonalised price series for both blueberries and kiwifruit. Lag length is specified as three, as prescribed by the Schwarz–Bayesian information criterion (SBIC) (Schwarz 1978). The analysis is run only on the precontamination sample, which runs through the first week of August 2018. We estimate the following VEC model (Engle and Granger 1987):

$$\Delta \hat{v}_t^S = \alpha^S z_{t-1} + \sum_{i=1,2} \left( \gamma_i^S(L) \Delta \hat{v}_{t-i}^S + \delta_i^S(L) \Delta \hat{v}_{t-1}^B + \xi_i^S(L) \hat{v}_t^K \right) + e_t^S \quad (2)$$

$$\Delta \hat{v}_t^B = \alpha^B z_{t-1} + \sum_{i=1,2} \left( \gamma_i^B(L) \Delta \hat{v}_{t-i}^S + \delta_i^B(L) \Delta \hat{v}_{t-1}^B + \xi_i^B(L) \hat{v}_t^K \right) + e_t^B \quad (3)$$

$$\Delta \hat{v}_t^K = \alpha^K z_{t-1} + \sum_{i=1,2} \left( \gamma_i^K(L) \Delta \hat{v}_{t-i}^S + \delta_i^K(L) \Delta \hat{v}_{t-1}^B + \xi_i^K(L) \hat{v}_t^K \right) + e_t^K \quad (4)$$

where $\Delta \hat{v}_t^k; k \in \{S, B, K\}$ is the difference between the de-seasonalised price at time $t$ and $t-1$ for strawberries, blueberries and kiwifruit, respectively. Functions $\gamma_i(L), \delta_i(L), \xi_i(L) \in \{1, 2\}$ are polynomials in the lag operator, and $z_{t-1} = \hat{v}_{t-1}^S - \hat{v}_{t-1}^B - \hat{v}_{t-1}^K$ is the error correction term. Interpreted in words, this specification allows the economic component embedded in the fresh strawberry price and the equivalent components for kiwifruit and blueberries to move together according to a long-term equilibrium. However, in each week, each market experiences an exogenous shock. The other markets adjust to this exogenous shock over the two subsequent weeks. Short-run and long-run coefficients are estimated using the Johansen (1995) maximum-likelihood method.

To determine whether blueberry prices should be included as controls, we run a postestimation F-test of the hypothesis that the short-run adjustment parameters on blueberry variables are jointly zero ($\delta_S = 0$) in the strawberry equation (Equation 2). We run a similar F-test for kiwifruit prices ($\xi_S = 0$). Results of these F-tests are reported in Table 3. We reject at 99% that the kiwifruit price variables should be excluded from the short-run strawberry price equation. Similarly, we reject at 90% that the blueberry price variables should be excluded. Thus, we conclude that both blueberry and kiwifruit prices should be included as controls for the co-integration and impact analyses.

After identifying the set of substitute products to include both kiwifruit and blueberries, we confirm that strawberry prices moved together with this substitute set prior to the needle contamination event using the Johansen test for co-integration (Johansen, 1991). Again, this test is run with the three-lag specification. At the 99% level ($\chi^2$ statistic 9.62), we find the economic component of the strawberry price is co-integrated with that of the substitute products. Accordingly, we proceed with the impact estimation.

**Table 3** Tests for optimal substitute mix

| Substitute | $\chi^2$ | *P*-value |
|---|---|---|
| Kiwifruit | 11.140 | 0.004 |
| Blueberry | 5.700 | 0.058 |

### 3.3 Historical co-integrating relationship

To determine the impact of the needle contamination on fresh strawberry prices, we now ask two questions: (1) does the data support the hypothesis that the needle contamination event caused a break in the co-integrating relationship between fresh strawberry prices and prices for kiwifruit and blueberries; and (2) what would strawberry prices have been had the break not occurred? To answer the first question, we rerun the VEC model described above over the entire sample period. We perform a test of the hypothesis that there was a structural break in the short-run adjustment parameters $(\gamma_S, \delta_k, \xi_k)$ in the strawberry Equation (2) beginning in the week of 6 September 2018 – the date the first needle was discovered. We test this hypothesis alternatively using the Wald and likelihood-ratio tests for a single structural break (Linden 2017). Results of this analysis are presented in Table 4. For both the Wald and likelihood-ratio tests, we reject at 99% the null hypothesis of no structural break on 6 September 2018. This suggests the data generating process for de-seasonalised strawberry prices did change around the time of the needle contamination and is strong evidence that the event did impact strawberry prices.

We answer the second question by measuring the historical co-integration relationship that occurred prior to the event. The parameter results from estimating the VEC model described by equations (2–4) on the precontamination sample are reported in Table 5. The lagged error correction term associated with the strawberry equation ($\alpha^S$) is negative and significant at 99%. Therefore, the weekly strawberry price adjusts downward to correct short-run deviations from the long-run trend. The error correction terms associated with the blueberry and kiwifruit equations ($\alpha^B$, $\alpha^K$) are also negative; however, only the kiwifruit correction term is statistically different from zero. The significance of the strawberry and kiwifruit error correction parameters indicates that these fresh fruit prices can deviate from their long-run relationship for extended periods of time.

Parameters $\delta_2^S$, $\xi_1^S$, and $\xi_2^S$ are all statistically significant at 95%, indicating the long-run adjustments in the strawberry price equation are significant. The estimated co-integrating vector (normalised with a coefficient of unity on the strawberry price) has an associated coefficient of 0.352 on the blueberry price (significant at 90%) and a coefficient of 0.443 on the kiwifruit price (significant at 99%). The significance of the lagged error terms and the significant coefficient estimated in the co-integrating vector indicate that a

**Table 4** Structural break test

| Test method | Date | Obs | $\chi^2$ statistic | *P*-value |
|---|---|---|---|---|
| Wald | 6 Sept 2018 | 254 | 87.982 | 0.000 |
| Likelihood ratio | 6 Sept 2018 | 254 | 19.859 | 0.006 |

**Table 5** VECM coefficient estimates

| Equation | Parameter | Estimate | SE | *P*-value |
|---|---|---|---|---|
| $\Delta\hat{v}_t^S$ | $\alpha^S$ | −0.304 | 0.041 | 0.000 |
| | $\gamma_1^S$ | 0.785 | 0.048 | 0.000 |
| | $\gamma_2^S$ | −0.291 | 0.064 | 0.000 |
| | $\delta_1^S$ | −0.081 | 0.103 | 0.434 |
| | $\delta_2^S$ | 0.240 | 0.105 | 0.022 |
| | $\xi_1^S$ | −0.344 | 0.164 | 0.036 |
| | $\xi_2^S$ | 0.545 | 0.165 | 0.001 |
| | Const | 0.000 | 0.009 | 0.989 |
| $\Delta\hat{v}_t^B$ | $\alpha^B$ | −0.016 | 0.022 | 0.473 |
| | $\gamma_1^B$ | 0.052 | 0.026 | 0.040 |
| | $\gamma_2^B$ | −0.056 | 0.034 | 0.101 |
| | $\delta_1^B$ | 0.894 | 0.055 | 0.000 |
| | $\delta_2^B$ | −0.536 | 0.056 | 0.000 |
| | $\xi_1^B$ | −0.041 | 0.088 | 0.643 |
| | $\xi_2^B$ | 0.056 | 0.088 | 0.521 |
| | Const | −0.001 | 0.005 | 0.917 |
| $\Delta\hat{v}_t^K$ | $\alpha^K$ | −0.047 | 0.015 | 0.002 |
| | $\gamma_1^K$ | −0.004 | 0.018 | 0.806 |
| | $\gamma_2^K$ | 0.059 | 0.023 | 0.012 |
| | $\delta_1^K$ | −0.010 | 0.038 | 0.796 |
| | $\delta_2^K$ | −0.007 | 0.038 | 0.861 |
| | $\xi_1^K$ | 0.719 | 0.060 | 0.000 |
| | $\xi_2^K$ | −0.391 | 0.060 | 0.000 |
| | Const | 0.001 | 0.003 | 0.761 |

Long-run equation

| | Coefficient | SE | *P*-value |
|---|---|---|---|
| $\hat{v}^S$ | 1.000 | | |
| $\hat{v}^B$ | 0.352 | 0.205 | 0.086 |
| $\hat{v}^K$ | 0.443 | 0.146 | 0.002 |
| Constant | 0.004 | | |

Johansen normalisation restriction imposed in long-term equation.

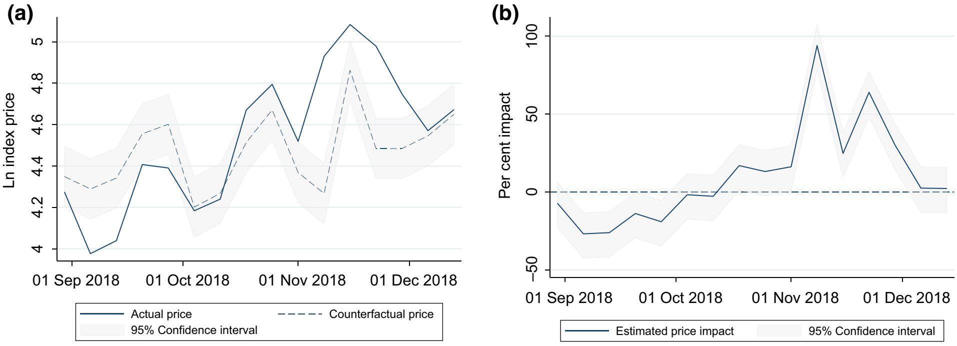
vector autoregressive (VAR) specification would yield inconsistent estimates due to misspecification (Schmitz 2018).

## 4. Price impact results

We assess the impact of the needle contamination on strawberry prices by comparing actual strawberry prices during and in the aftermath of the scandal with the counterfactual strawberry prices implied by the historical co-integration relationship and seasonal oscillation pattern that occurred prior to the event. To generate the series of counterfactual strawberry prices (denoted $Ln\,\hat{P}_t^S$) that would have resulted from September to December 2018 had the contamination event not happened, we combine the VEC parameters from Table 5 with our estimate of the seasonal oscillation for strawberry prices in Table 1. Panel (a) of Figure 4 plots this counterfactual price series (and associated confidence intervals) versus actual strawberry prices from September to December 2018. Panel (b) of the Figure plots the implied price impact – expressed as a percentage of the counterfactual price – in each week.

Referring to Figure 4, we see that in the four-week period at the height of the strawberry needle crisis, actual wholesale fresh strawberry prices were statistically significantly below counterfactual levels. This reduction is largest in magnitude on the week of 6 September 2018 – the week in which the first needles were discovered. In this week, actual prices were 26.8% less than counterfactual levels. Five weeks later (11 October), the price impact had reduced to −2.7%. We note that in the latter weeks of September and early weeks of October, statistical significance of these impacts is diminished. We attribute this price reduction in the midst of the needle scare primarily to substitution by consumers and other downstream actors away from fresh strawberries.

However, as shown in Figure 4, in the aftermath of the needle scare – rather than returning to historical levels – fresh strawberry prices rose well above historical levels. In the week of 11 November 2018 – the week in which My Ut Trinh was arrested for the September contamination events – wholesale fresh strawberry prices were 94% higher than the counterfactual comparator price. Prices remained statistically significantly above counterfactual prices into December. We attribute this secondary effect to producer and government responses. As discussed in Section 2, strawberry growers and retailers took many steps to preserve brand reputation, including installing metal detectors to certify that strawberries were free from contamination and removing fresh strawberries from retail shelves. Moreover, the Government issued recalls of all affected brands from distributors and wholesalers. All of these reduce the supply of fresh strawberries, thereby driving prices up after the threat has been eliminated. Thus, comparing the initial price impacts (likely the result of downstream substitution) with the




Actual and predicted price residuals — Implied price impact

**Figure 4** Price impact estimates. [Colour figure can be viewed at wileyonlinelibrary.com] latter impacts (likely the result of producer and government responses), we see that the latter effect dominated the market.

## 5. Conclusion

In September 2018, a former strawberry-picking supervisor for a strawberry farm in New South Wales was arrested for inserting sewing needles in hundreds of punnets of fresh strawberries sold in retail stores across several Australian states. This paper develops a four-step procedure to estimate the effects of the 2018 Australian strawberry needle contamination on wholesale fresh strawberry prices. Our results indicate a drop in wholesale fresh strawberries prices of about 20% while the needle crisis was ongoing. However, public and private supply restrictions caused wholesale prices to *rise* by up to 94% relative to expected dynamics over several weeks in the immediate wake of the incident.

Our findings are not without qualification. First, the analysis relies on only a single market in Melbourne, Victoria, while the scare originated in New South Wales and spread to several locations across several states. Reliance on a single Australian market is the result of convenience rather than desirability. Melbourne is the only locale for which weekly fresh produce price data are publicly available. Yet, one could imagine the price dynamics in locations nearer to the contamination 'hot spot' or locations further afield may have exhibited very different price dynamics from those in Melbourne.

Second, our analysis is conducted at the wholesale, as opposed to the farm gate or retail level. Again, this is purely the result of data availability. One could imagine prices at different stages of the supply chain exhibit disparate impacts in the midst and immediate wake of a terrorist event. For example, we hypothesise farm-gate prices would be much more sensitive to agro-terrorism than wholesale prices, and retail prices less so still. Finally, we cannot assign revenue or more complex surplus impacts resulting from the strawberry contamination because we are unable to access fresh strawberry sales data of sufficient granularity and reliability to undertake the analysis. Such data may allow for spatial and temporal welfare calculations that would aid in the understanding of the full costs of the 2018 needle crisis to the Australian strawberry industry.

### Data availability statement

The data that support the findings of this study are available in from the Australian Bureau of Agricultural and Resource Economics (2019), 'Movements in selected fruit and vegetable prices', Weekly Australian Climate, Water and Agricultural Update, available at http://www.agriculture.gov.au/abares/publications/weekly_update/movements-selected-fruit-vegetable-prices#kiwi-fruit-hayward-strawberry-apple-royal-gala-and-avocado-haas

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